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THE AUTONOMY OF TRADE ELASTICITIES: CHOICE AND CONSEQUENCES

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ABSTRACT

Fifty years of econometric work on trade assumes that trade elasticities are invariant to changes in spending patterns, that prices can be taken as given, and that expenditures on domestic and foreign goods can be studied independently of each other. To relax these assumptions, this paper assembles a simultaneous model explaining trade among Canada, Japan, and the United States. 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 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 undermines our effectiveness in addressing questions relevant to economic interactions among nations. Specifically, the estimates from the Rotterdam model predict that asymmetries in income elasticities, which were important once, have vanished.

The Autonomy of Trade Elasticities: Choice and Consequences

Jaime Marquez[1]

1. Introduction

Virtually all of our knowledge about the role of international trade in shaping macroeconomic interdependencies rests on formulations that neglect the implications of optimization and simultaneity: income and price elasticities are viewed as invariant to changing spending patterns, domestic and foreign prices are taken as given, and purchases of domestic and foreign products are modeled independently of each other. These assumptions are restrictive indeed and their persistence for the last fifty years creates the presumption that both optimization and simultaneity are more important in theory than in practice. This paper documents the practical implications of abstracting from these two considerations when studying trade among Canada, Japan, and the United States.

The analysis begins in section 2 by reviewing the drawbacks of formulations treating trade elasticities as autonomous parameters: if these formulations are to be consistent with optimizing behavior, then expenditure shares must be both fixed and equal across spending categories, a requirement not supported by the data. Section 3 assembles a model that both avoids this restriction and does not take prices as given. Specifically, spending decisions are assumed to follow the Rotterdam model developed by Barten (1964, 1966) and Theil (1965) which embodies the properties associated with utility maximization. Pricing decisions are assumed to follow the model developed

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by Gagnon and Knetter (1990) which allows foreign producers to exercise price discrimination and to update their pricing decisions in response to changes in demand. Section 4 estimates the parameters of the spending and pricing equations using the Full Information Maximum Likelihood (FIML) estimator, a choice that avoids simultaneity biases, incorporates the implications associated with optimization, and allows pricing decisions to internalize explicitly the structure of preferences. This estimation uses annual observations for 1965-1987 and relies on bilateral-price data. This reliance recognizes the exporters' ability to vary prices across foreign destinations, a possibility ignored by most previous empirical studies of international trade.

To examine the implications of the assumed autonomy in trade elasticities, section 5 uses the elasticity estimates of both the Rotterdam model and the log-linear model to re-examine the asymmetry of income elasticities noted by Houthakker and Magee (1969). Based on the evidence, the estimates from the Rotterdam model predict that asymmetries in income elasticities, which were important once, have vanished whereas the estimates from the log-linear model predict that these asymmetries are permanent. This difference in predictions is not intended to replace the existing elasticity estimates with those of the Rotterdam model but to redirect the econometric work on trade away from estimating elasticities that are invariant to changes in spending patterns.

2. Autonomy of Trade Elasticities: Drawbacks

Inspecting fifty years of econometric work on trade reveals a propensity to relate international trade to income and prices using a log-linear formulation:[2]

$$\text{(1)} \qquad \ln q_i = \alpha_i + \eta_i \ln(y / P) + \Sigma_j \, \varepsilon_{ij} \ln p_j ,$$

where q_i represents imports of the ith product, p_i is the associated price, p_n is the price of the domestic product, $y = \Sigma p_j q_j$, P is the overall price level, η_i is the income elasticity, ε_{ij} is the compensated price

2 Appendix A surveys 110 papers written between 1941 and 1991 that study the behavior of import demand for Canada, Japan, and the United States among other countries. Of these papers, 74 of them rely on a logarithmic formulation. For an early critique of the use of elasticities in international trade see Balogh and Streeten (1951).

elasticity with respect to changes in the price of the jth product, and $\sum_j \varepsilon_{ij} = 0\ \forall i$. By treating these elasticities as autonomous parameters, eq. (1) facilitates addressing questions couched in terms of elasticities such as the similarity of income elasticities for exports and imports and whether the Marshall-Lerner condition holds. In addition, eq. (1) is linear in the parameters, a feature that helps their econometric estimation. But the assumed autonomy of η_i and ε_{ij} creates several difficulties that limit their usefulness, even if the best of data sets and estimation methods were employed.

First, because an elasticity is the ratio between a marginal propensity and an expenditure share, treating the parameters of (1) as autonomous implies either that the associated propensities and shares are fixed or that their changes are mutually offsetting. For example, the income elasticity of the ith good, η_i implied by (1) equals $(\partial p_i q_i / \partial y)/(p_i q_i / y)$ where the numerator is the marginal propensity to purchase the ith good and the denominator is the associated expenditure share. As section 4 shows, these shares have changed substantially and, therefore, the assumed invariance of η_i requires $(\partial p_i q_i / \partial y)$ to respond in order to offset changes in the ith expenditure share, a response with no theoretical basis.

Second, if the parameters of (1) are constant and individuals optimize their spending decisions subject to a linear budget constraint, then these parameters are known:[3] $\eta_i = 1$, $\varepsilon_{ii} = -1$, and $\varepsilon_{ij} = 0$ $\forall j \neq i,n$, a knowledge that makes their econometric estimation redundant. Moreover, the estimates of η_i and ε_{ii} found in the literature depart from these classical benchmarks,[4] a gap that turns the seemingly redundant role of econometrics into a conflictive one. Third, eq. (1) is not suitable for studying interdependencies among spending decisions because the symmetry of the Slutsky matrix is not equivalent to $\varepsilon_{ij} = \varepsilon_{ji}$ but to $w_i \varepsilon_{ij} = w_j \varepsilon_{ji}$ where w_i is the share of spending devoted to the ith product. Thus for (1) to be consistent with such symmetry, purchases must be evenly distributed

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[3] See Deaton and Muellbauer (1980, p. 17) and Koopmans and Uzawa (1991, p. 3).

[4] See Chang (1946, tables 4 and 5); Houthakker and Magee (1969, tables 1 and 4); Thursby and Thursby (1984, table 3); and Marquez (1990, table 1).

across spending categories (and thus fixed), a requirement contradicting the evidence of the last twenty years.[5]

In addition to facing these limitations, many of the existing estimated elasticities are based on formulations that take prices as given and abstract from the interactions among spending decisions.[6] Price-taking behavior is not, however, a reasonable assumption for explaining trade among Canada, Japan, and the United States. For example, arguing that changes in U.S. imports from Canada leave the associated import price unchanged is questionable when more than half of Canadian exports are destined for the U.S. market. Similarly, setting $\varepsilon_{ij} = 0 \; \forall j \neq i,n$ introduces an omitted-variable bias and removes an important interdependency in spending decisions. Finally, most previous empirical formulations of 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 Brenton, 1989). Specifically, the presence of a budget constraint implies that $\sum_j w_j \eta_j = 1$ and ignoring this constraint affects the associated coefficient estimates even if prices could be taken as given and cross-price effects were zero.

Overall, to say that eq. (1) embodies important limitations is stating the obvious. Moreover, these limitations have been noted before but very little empirical work addresses them despite a persistent professional dissatisfaction:

> 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]
>
> 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]

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5 Appendix C examines how ignoring changes in the composition of expenditures affects the least-squares estimates of trade elasticities.

6 Table A1 of appendix A reports that three-quarters of the analyses surveyed here rely on single-equation estimation methods that abstract from the implications of optimization. (Orcutt (1950) is among the first to criticize using ordinary least squares for estimating trade elasticities.) The remaining one quarter is split evenly between studies recognizing simultaneity and studies incorporating optimization; studies recognizing the statistical implications of both optimization and simultaneity are not available, which is the gap filled by this paper.

To address these limitations, I rely on a model where consumers substitute between domestic and foreign products, exporters develop pricing policies that differentiate across export markets, and parameter estimates recognize the interdependence between spending and pricing decisions. This analysis, however, has limitations of its own which I will highlight throughout the discussion.

3. Empirical Formulation

3.1 Spending Decisions

The analysis assumes that individuals determine their spending on foreign and domestic products by maximizing a utility function $u(q_o...q_n)$ subject to $y=\Sigma\, p_j\, q_j$ where q_j represents purchases of the product made in country j and p_j is the associated domestic-currency price. 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 (Barten 1964, 1966; Theil 1965):

$$(2) \qquad w_{it}\, \mathrm{dln}q_{it} = [\partial(p_{it}\, q_{it}) / \partial y_t]\, \mathrm{dln}(y/P)_t + \sum_{j=0}^{n} [w_{it}\, (p_{jt} / q_{it})(\partial q_{it} / \partial p_{jt})]\, \mathrm{dln}p_{jt},$$

where

$$w_{it} = p_{it}\, q_{it} / y_t,$$

$$p_{it} = (1+\tau_{it})\, p_{mit},$$

$$\mathrm{d}P_t = \sum_{j=0}^{n} w_{jt}\, \mathrm{dln}p_{jt},$$

τ_i is the tariff rate (zero for purchases of domestic products, q_n) and p_{mit} is the domestic-currency import price of products made in country i, $i<n$.[7]

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[7] This formulation assumes that individuals differentiate products according to the place of production (Armington, 1969) and that spending decisions are governed by utility maximization alone. The literature offers analyses treating imports as an input to the production process but the existing work focuses exclusively on U.S. aggregate imports (see Burgess 1974, Appelbaum and Kohli 1979, Aw and Roberts 1985, Kohli 1990, Kohli 1991). Viewing imports solely as either factor inputs or consumer goods is not satisfactory, but the literature has yet to offer separate estimates of the structure of factor and consumer demands. I assume that the distortions arising from classifying an import item as a consumer good instead of a productive input are of secondary importance relative to the distortions introduced by ignoring optimization altogether.

In the absence of further restrictions, (2) cannot be rejected by the data and thus is not suitable for applied analysis. To implement (2) empirically, the Rotterdam model restricts $\mu_i = [\partial(p_{it} q_{it}) / \partial y_t]$ and $\pi_{ij} = [w_{it}(p_{jt} / q_{it})(\partial q_{it} / \partial p_{jt})]$ to be invariant to changes in income and prices. 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 ith good.

Treating both μ_i and π_{ij} as autonomous transforms (2) into

$$(3) \qquad w_{it}\,\mathrm{dln}q_{it} = \mu_i\,\mathrm{dln}(y/P)_t + \sum_{j=0}^{n} \pi_{ij}\,\mathrm{dln}p_{jt} + r_{it}\ , \quad i=0,\ldots, n,$$

where r_{it} is a random disturbance containing the (second-order) approximation terms induced by the assumed autonomy of μ_i and π_{ij}.[8] 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 (3) are constant, as suggested by Byron (1984) and Barnett (1984). The income and price elasticities associated with (3) are μ_i / w_{it} and π_{ij} / w_{it}, respectively, and they vary in response to reallocations of expenditures.[9]

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

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[8] Equation (3) corresponds to the *Absolute Price Rotterdam System*. For further developments and applications of this model see Barten (1964, 1966); Theil (1965); Theil (1971), chapters 7 and 11; Deaton (1974); Clements and Theil (1978); and Theil and Clements (1978, 1987).

[9] Critics of the Rotterdam model (Phlips 1974 and Goldberger 1987) argue that the invariance of μ and π 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 (per-capita) parameters. Indeed Barnett (1979, 1891) derives the implications for per-capita 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 per-capita 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 (1980, p. 73) note.

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

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

(6) *symmetry*: $\pi_{ij} = \pi_{ji} \; \forall \, i,j \ni i \neq j$.

I will not be testing, however, whether these parametric relations are supported by the data because they cannot exist independently of utility maximization, which is the maintained hypothesis of this paper.[10] Moreover, the alternative hypothesis for such tests is not generally well defined from an economic standpoint. For example, the alternative hypothesis to eq. (5) involves individuals having permanent money illusion, which is hard to justify.[11] Instead of undertaking such testing, I examine whether the eigenvalues of the Slutsky matrix support the second-order conditions for utility maximization--namely, whether this matrix is negative semidefinite. I find this alternative appealing because the Slutsky coefficients are estimated without sign-restrictions which allows the associated eigenvalues to reject the model. Moreover, finding that the data "support" homogeneity and symmetry is not useful if the estimated Slutsky matrix is not negative semidefinite because it would mean that individuals are *minimizing* utility. Finally, by focusing on these eigenvalues, I avoid technical problems involving insufficient degrees of freedom, which have plagued the econometric testing of demand systems.[12]

As developed here, the Rotterdam model has several limitations. First, I abstract from intertemporal substitution and treat decisions regarding labor supply and asset holdings as separable.

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[10] See Keuzenkamp and Barten (1991) for a history of the philosophical issues associated with testing these relations.

[11] In contrast, testing the absence of cross-price effects has a well-defined alternative hypothesis--namely, $\pi_{ij} = \pi_{ji} \neq 0$--that does not violate utility maximization.

[12] Barten and Geyskens (1975) emphasize the importance of examining the second-order (negativity) conditions; see Theil and Clements (1987) for a discussion of the technical problems that arise when testing homogeneity and symmetry.

These two limitations are important but they apply with the same force to the existing literature, which does not enjoy the advantages offered here. Second, I model the behavior of composite goods and thus ignore their product composition. 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. Third, the model abstracts from dock-strikes, trade treaties, "voluntary" quota restrictions on trade, and other non-tariff trade barriers which affect international trade regardless of changes in income and prices. Finally, the Rotterdam model need not offer the best representation of demand behavior for these countries but the analysis uses it to emphasize comparability of results.

3.2 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

$$\ln(p_{nt} / \rho_{nt}) = \theta_n + \beta_n \ln(\omega_{nt} / \rho_{nt}) + e_{nt} ,$$

where ρ_{nt} is the rental rate of the capital stock, ω_{nt} is the wage rate, and e_{nt} is a random disturbance.[13]

Following the pricing-to-market model developed by Gagnon and Knetter (1990), the analysis explains import prices with

$$\ln p_{mit} = \theta_i + \beta_i \ln C_{it} + \delta_i \ln E_{it} + \gamma_i \ln p_{nt} + e_{it} , \tag{7}$$

where p_{mit} is the domestic-currency import price of products made in country i, $i<n$, 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/domestic currency), and e_{it} is a random disturbance.[14] According to (7), p_{mit}

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[13] Although both labor costs and rental rates are treated as given, I retain this equation for parameter estimation because there is no a-priori reason to believe that the disturbances of the spending-pricing system are orthogonal among themselves.

[14] Equation (7) is strictly derived from profit maximization but I omit the details to conserve space; see Gagnon and Knetter (1990) for these details.

increases in response to increases in foreign marginal costs ($\beta_i>0$) and depreciations of the domestic currency ($dE_{it}<0$, $\delta_i<0$) where δ_i is known as the "passthrough coefficient." Eq. (7) also says that foreign firms anchor their export price to p_{nt}, the price of their destination market. Thus I interpret γ_i as the degree of pricing-to-market.

As formulated, (7) assumes that exporters pass a fixed percentage δ_i of changes in exchange rates to their prices even if the demand for their product becomes price elastic. To relax this restriction, Gagnon and Knetter (1990) recognize that this passthrough is sensitive to changes in the price elasticity of demand and find that

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

where $\psi_{ii,t}$ is the uncompensated price elasticity in absolute terms. For the Rotterdam model, this elasticity equals $(-\pi_{ii} / w_{it} + \mu_i)$ which turns eq. (8) into

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

In addition to internalizing the structure of preferences into pricing decisions, (9) has several properties of interest.[15] First, the passthrough of the ith supplier responds to changes in prices from all other foreign suppliers because w_{it}, via q_{it}, depends on *all* prices.[16] Second, δ_{it} approaches $\{[\pi_{ii}/(2\pi_{ii} - \mu_i)] - 1\}$ as w_{it} approaches one meaning that even a monopolist faces a limit on its ability to pass changes in the exchange rate to p_{mit}. Third, δ_{it} approaches -1 as the own-price effect approaches zero ($\pi_{ii} \to 0$) implying that sellers stabilize prices in their currencies and that buyers fully absorb changes in exchange rates.

Substituting (9) into (7) gives

(10) $$\ln p_{mit} = \theta_i + \beta_i \ln C_{it} + \{[\pi_{ii}/(2\pi_{ii} - \mu_i w_{it})] - 1\}\ln E_{it} + \gamma_i \ln p_{nt} + e_{it},$$

which says that import prices depend on the level of sales, as reflected in the expenditure share, and this dependence makes the spending-pricing system truly simultaneous: Prices influence quantities

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[15] See Marquez (1991b, appendix B) for the derivation of equation (9).

[16] Changes in domestic spending also induce changes in expenditure shares with a corresponding effect on the passthrough coefficient.

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.[17]

4. Econometric Estimation

4.1 Data

The differentiating feature of this study's data set is the availability of bilateral prices for trade among Canada, Japan, and the United States. But to recognize that these three countries have other trading partners, I allow consumers from these three countries to purchase products from both Germany and the rest-of-the-world. The feature that makes bilateral prices useful is their recognition of exporters' potential for price discrimination across foreign destinations. The alternative to using bilateral prices is to rely on multilateral prices which assume that exporters either face only one foreign market or ignore elasticity differentials across export markets.

To measure domestic spending on foreign products, in real terms, I deflate the value of bilateral imports by the corresponding bilateral price. Domestic spending on domestic products, in real terms, is measured as real GNP minus exports in real terms. Thus domestic spending on foreign and domestic products adds up to domestic spending. The data for prices and trade are annual and come from the Commission of the European Communities; data for other variables come from national sources with the details appearing in appendix B.

Inspecting the evolution of expenditure shares for these countries reveals three interesting features (table 1). First, the share of spending devoted to foreign products has changed substantially since 1965 increasing by six percentage points for Canada and five percentage points for the United States but declining three percentage points for Japan. Second, spending is not evenly distributed across spending categories. Specifically, the average expenditure share on domestic products ranges

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[17] See Baldwin (1988), Knetter (1989), and Marston (1990).

Table 1
Expenditure and Import Shares

Canada		Expenditure Shares					Import Shares			
		US	Japan	Germany	R.O.W.	Domestic	US	Japan	Germany	R.O.W.
	1965	11.91	0.45	0.41	3.23	83.99	74.41	2.83	2.58	20.18
	1970	12.93	0.76	0.48	3.05	82.78	75.08	4.41	2.81	17.70
	1975	15.26	0.78	0.51	3.48	79.97	76.19	3.90	2.55	17.36
	1980	17.29	1.02	0.53	3.67	77.49	76.82	4.54	2.34	16.30
	1985	17.28	1.44	0.64	3.87	76.77	74.40	6.20	2.74	16.66
	1986	16.71	1.68	0.76	4.15	76.70	71.71	7.22	3.26	17.81
	1987	15.91	1.53	0.72	4.07	77.77	71.58	6.90	3.23	18.29
Japan		**Canada**	**US**	**Germany**	**R.O.W.**	**Domestic**	**Canada**	**US**	**Germany**	**R.O.W.**
	1965	0.44	2.93	0.28	3.63	92.72	6.06	40.22	3.79	49.93
	1970	0.50	3.00	0.33	3.67	92.50	6.67	39.94	4.43	48.97
	1975	0.55	2.55	0.25	3.91	92.74	7.56	35.14	3.44	53.85
	1980	0.50	2.57	0.26	3.82	92.85	6.94	35.92	3.67	53.47
	1985	0.41	2.24	0.25	3.23	93.88	6.64	36.50	4.11	52.75
	1986	0.28	1.53	0.25	2.34	95.60	6.28	34.78	5.63	53.31
	1987	0.27	1.49	0.29	2.40	95.54	6.09	33.53	6.47	53.91
US		**Canada**	**Japan**	**Germany**	**R.O.W.**	**Domestic**	**Canada**	**Japan**	**Germany**	**R.O.W.**
	1965	0.76	0.38	0.21	1.63	97.02	25.53	12.76	7.09	54.62
	1970	1.20	0.63	0.34	1.80	96.04	30.16	15.98	8.50	45.36
	1975	1.55	0.80	0.38	2.61	94.67	29.04	14.97	7.09	48.91
	1980	1.63	1.31	0.49	4.13	92.44	21.54	17.35	6.53	54.58
	1985	1.78	1.89	0.56	3.76	92.02	22.25	23.73	6.96	47.07
	1986	1.65	2.12	0.65	3.66	91.92	20.38	26.30	8.01	45.32
	1987	1.65	2.06	0.65	3.90	91.74	19.98	24.91	7.93	47.18

Sources: Data for International Trade flows are from the Volimex Data Tape provided by the Commission of the European Communities. Appendix B contains the details and sources for data on domestic variables. R.O.W. stands for the Rest of the World which, in the present study, excludes OPEC and Centrally Planned Economies and includes Asian NICs (exc.Indonesia), Latin America (exc. Venezuela), Other European Countries. 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.

from 80 percent for Canada to 94 percent for both Japan and the United States. Third, international competition has been substantial with Japan's share of import markets increasing largely at the expense of the U.S. market share; competition for the U.S. market has been pronounced also with Japan displacing all other sources of imports except Germany.

4.2 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 over 1965-1987, one country at a time. The estimation drops the equation for purchases from the rest-of-the-world, q_{ot}, to avoid the singularity of the covariance matrix of the disturbances.[18] Overall, the estimating system for each country has eight equations explaining prices and per-capita purchases of four expenditure categories: domestic products, imports from Germany, and imports from the other two countries explicitly considered in this paper. The exogenous variables are nominal income, four nominal wage rates, three nominal exchange rates, the tariff rate, the rental rate for capital, and the price of imports from the rest of the world, p_{ot}, which is the numeraire.[19]

To examine parameter constancy, the analysis tests whether the expected, one-step ahead prediction error is zero for each year from 1973 to 1987. Implementing this test involves including in each equation a dummy variable that equals one for the year being tested (zero otherwise) and testing whether the coefficients of these eight dummies are jointly equal to zero. Figure 1 shows the

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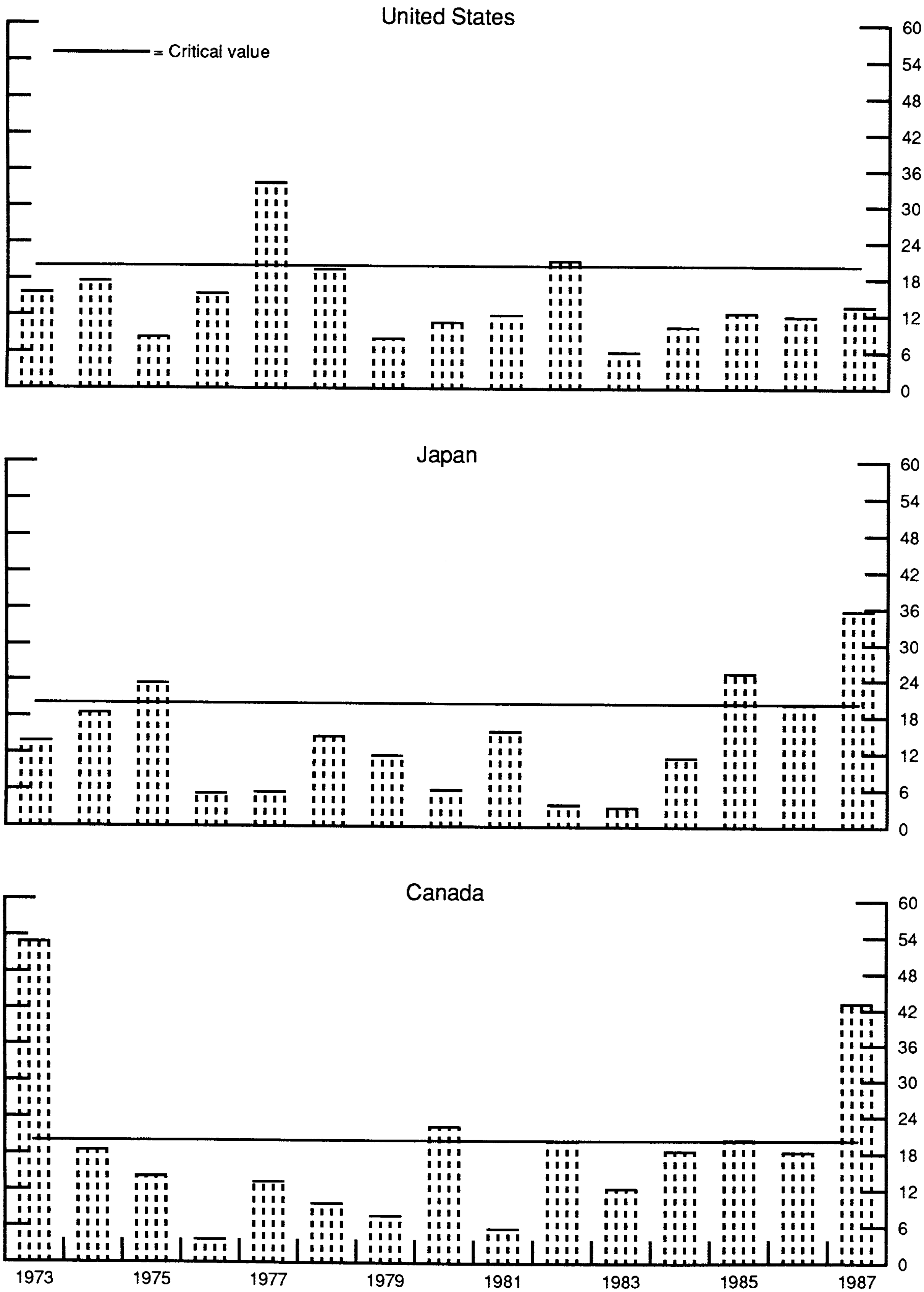
[18] The parameter estimates for the equation explaining q_{ot} can be recovered using equations (4)-(6).

[19] 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 (3). This substitution introduces an additional approximation error into the disturbance of (3). 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. The estimates are robust to changes in initial conditions.

Figure 1

Parameter Constancy: Log-likelihood Ratio Tests



values of the associated log-likelihood ratio test and the critical value for a χ^2 with eight degrees of freedom and a one percent significance level. Overall, the evidence supports the hypothesis of parameter constancy but there are a few exceptions: Out of fifteen years, the data show significant prediction errors in three years for Canada and Japan and two years for the United States. Besides being few in number, these prediction errors could arise from dock-strikes, trade treaties, and "voluntary" trade quotas which affect international trade regardless of changes in income and prices. To confirm that these errors are not due to changes in preferences, appendix C reports the 95 percent confidence bands for the time-series of the coefficient estimates associated with figure 1. The results point to a remarkable constancy of coefficient estimates. This support for the hypothesis of parameter constancy means that the (second-order) approximation errors in the disturbances of (3) are sufficiently small to treat the Rotterdam model as an adequate approximation to the per-capita demand functions for these three countries.

The parameter estimates for the spending and pricing equations (table 2) exhibit several features of interest. First, marginal budget shares ($\hat{\mu}_i$s) are positive, significant, and vary from 0.004 for Japanese imports from Germany to 0.85 for U.S. purchases of U.S. products. Second, purchases of domestic products have the largest marginal budget share ranging from 0.51 for Canada to more than 0.80 for Japan and the United States. Third, own-price effects ($\hat{\pi}_{ii}$s) are negative and vary from -0.003 for U.S. purchases from Germany to -0.083 for Canadian purchases of U.S. products; the sole exception to this pattern is the own-price effect for imports of Japan from Canada which has the wrong sign but is not significant. Fourth, substitutability among foreign products is quantitatively small and often not statistically significant. The few cross-price effects that are statistically significant, however, convey useful information. For example, Germany appears as the most important competitor to the United States in the Canadian and Japanese markets.

Sixth, the most important factor explaining bilateral export prices of both Japan and the United States is the price of their respective destination markets. Japan and the United States differ, however, in their pricing of domestic products with the share of labor costs being substantially smaller in Japan than in the United States: 38 percent and 85 percent respectively. The behavior of

Table 2

Coefficient Estimates for Spending and Pricing Equations: FIML 1965-87*

		Spending Decisions						Pricing Decisions	
Country	Source	Marginal Budget Shares μ_i	Slutsky Coefficients π_{ij}					Pricing to Market γ_i	Labor Costs β_i
			US	**Japan**	**Germany**	**Canada**	**Eigenvalues**		
Canada	**US**	0.347 0.039	-0.079 0.021	0.009 0.007	0.004 0.002	0.035 0.023	-0.099 0.022	0.955 0.427	0.052 0.452
	Japan	0.029 0.009		-0.016 0.005	-0.0001 0.0008	0.004 0.007	-0.025 0.004	0.772 0.225	0.158 0.329
	Germany	0.015 0.003			-0.004 0.001	0.002 0.002	-0.010 0.016	0.154 0.228	0.874 0.337
	Canada	0.511 0.054				-0.039 0.032	-0.002 0.034		0.745 0.012
			Canada	**US**	**Germany**	**Japan**	**Eigenvalues**		
Japan	**Canada**	0.008 0.002	0.0004 0.0003	0.002 0.002	0.0001 0.0005	-0.002 0.002	-0.039 0.016	-0.025 0.284	0.404 0.182
	US	0.039 0.008		-0.015 0.006	0.002 0.001	-0.010 0.004	-0.011 0.005	1.020 0.319	-0.024 0.190
	Germany	0.004 0.002			-0.003 0.0004	0.0002 0.001	-0.003 0.010	-0.191 0.472	0.945 0.469
	Japan	0.811 0.041				-0.034 0.017	0.0009 0.011		0.379 0.071
			Canada	**Japan**	**Germany**	**US**	**Eigenvalues**		
US	**Canada**	0.041 0.006	-0.013 0.002	-0.0009 0.0010	-0.002 0.001	0.009 0.002	-0.020 0.003	1.402 0.478	0.027 0.450
	Japan	0.030 0.009		-0.009 0.002	0.002 0.001	0.005 0.003	-0.009 0.001	1.421 0.076	-0.041 0.108
	Germany	0.013 0.004			-0.002 0.001	0.001 0.002	-0.002 0.003	1.191 0.173	-0.221 0.237
	US	0.847 0.025				-0.008 0.010	0.0003 0.015		0.853 0.022

* The estimates for Canada are based on data for 1966-87. For a given category, the top entry represents the point estimate and the bottom entry represents the associated standard error.

German and Canadian bilateral export prices, in contrast, differs depending on the destination market: For products destined for the United States, the U.S. domestic price is the most important factor; prices for products destined elsewhere are governed by domestic costs. Seventh, the estimated passthrough coefficients are virtually constant over time and range from -0.63 for Canadian imports from the United States to -0.48 for Japanese imports from Canada (figure 2).[20] This constancy contradicts the view that foreign producers reacted to the post-1984 depreciation of the dollar by lowering the degree of exchange-rate passthrough to maintain their market shares in the United States (see Hooper and Mann, 1989).

Finally, the results support the second-order conditions for utility maximization in these three countries, as evidenced by the eigenvalues of the associated Slutsky matrices. Specifically, all of the eigenvalues for Canada's Slutsky matrix are negative; for Japan and the United States, all of eigenvalues except the fourth are negative (table 2). This fourth eigenvalue has a positive sign but a magnitude virtually equal to zero.[21] Overall, the results suggest that these Slutsky matrices can be treated as being negative *semi*definite.

5. The Autonomy of Trade Elasticities

To evaluate the consequences of assuming that trade elasticities are invariant to spending patterns, figure 3 compares the income elasticities implied by the Rotterdam model with those of the log-linear model, eq. (1).[22] According to the evidence, income elasticities from the Rotterdam model are positive, significantly greater than zero, and fluctuate over time. For the United States, these elasticities decline throughout the sample: from 8.0 to 1.5 for imports from Japan (panel 1) and from 5.5

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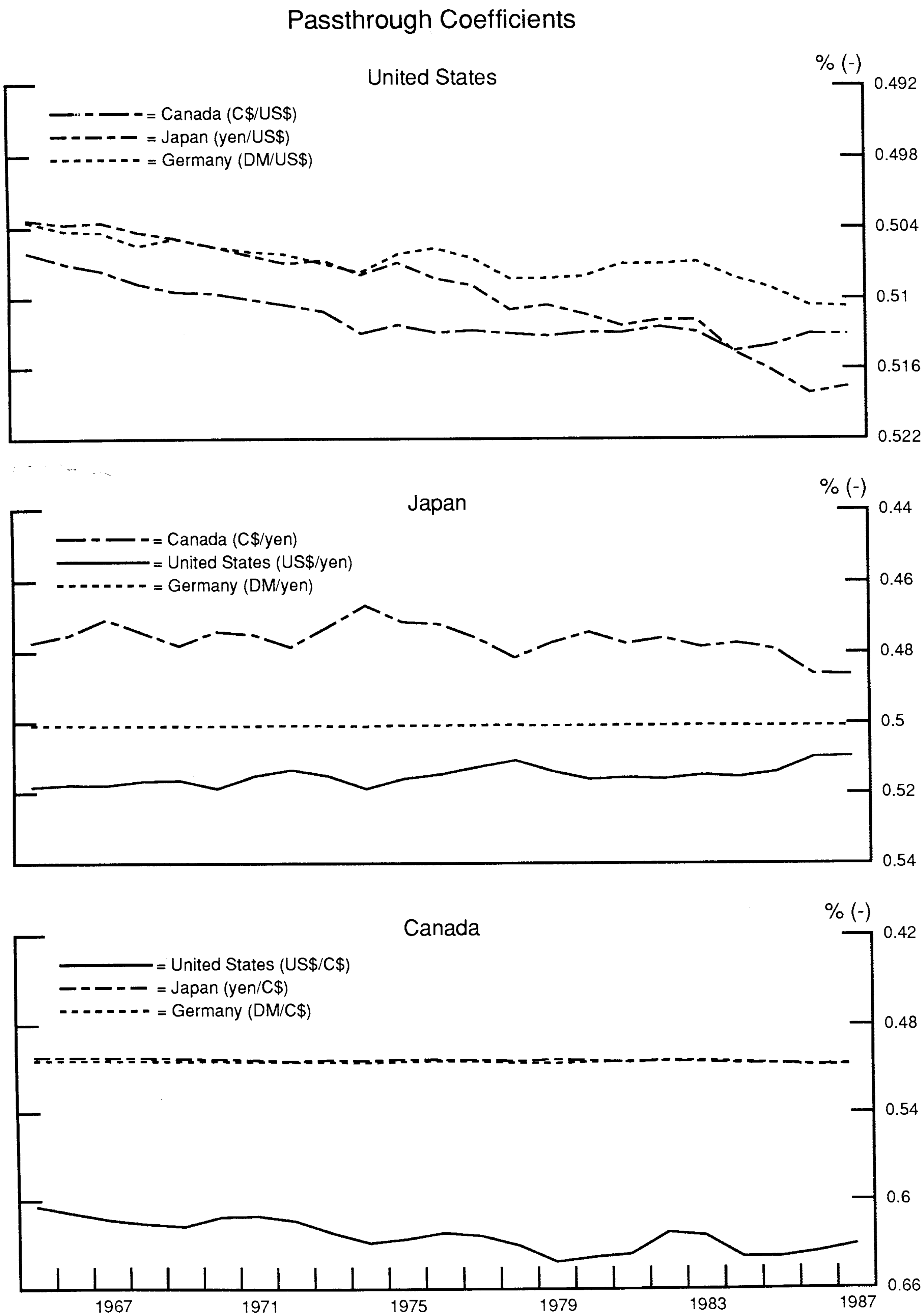
[20] These passthrough coefficients are calculated using both equation (9) and the $\hat{\mu}$'s and $\hat{\pi}$'s from table 2.

[21] The standard errors of the eigenvalues are calculated using a Taylor expansion (see appendix C). Note that the fifth eigenvalue of the Slutsky matrix (not shown) is zero (see Barten and Geyskens, 1975).

[22] Computing the estimated income elasticities for a given period involves taking the ratio between the marginal budget share, μ_i (from table 2) and the associated expenditure share w_{it} (from table 1); the own-price elasticities (not shown) are equal to the ratio between π_{ii} and w_{it}.

Figure 2


Passthrough Coefficients
United States
% (-)
= Canada (C$/US$)
= Japan (yen/US$)
= Germany (DM/US$)
0.492
0.498
0.504
0.51
0.516
0.522
Japan
% (-)
= Canada (C$/yen)
= United States (US$/yen)
= Germany (DM/yen)
0.44
0.46
0.48
0.5
0.52
0.54
Canada
% (-)
= United States (US$/C$)
= Japan (yen/C$)
= Germany (DM/C$)
0.42
0.48
0.54
0.6
0.66
1967
1971
1975
1979
1983
1987

to 2.5 for imports from Canada (panel 2). These declines confirm earlier reports of parameter instability in constant-elasticity models estimated with U.S. data:

> The results reveal significant upward shifts in estimates of both income and price elasticities of U.S. non-oil imports during the early 1960's, followed by significant downward shifts in those estimates during the early 1970's. [Hooper, 1978, p. 3]

> 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]

But unlike previous studies, the elasticity declines shown in figure 3 are an outcome of optimization and not a defective feature of the model.

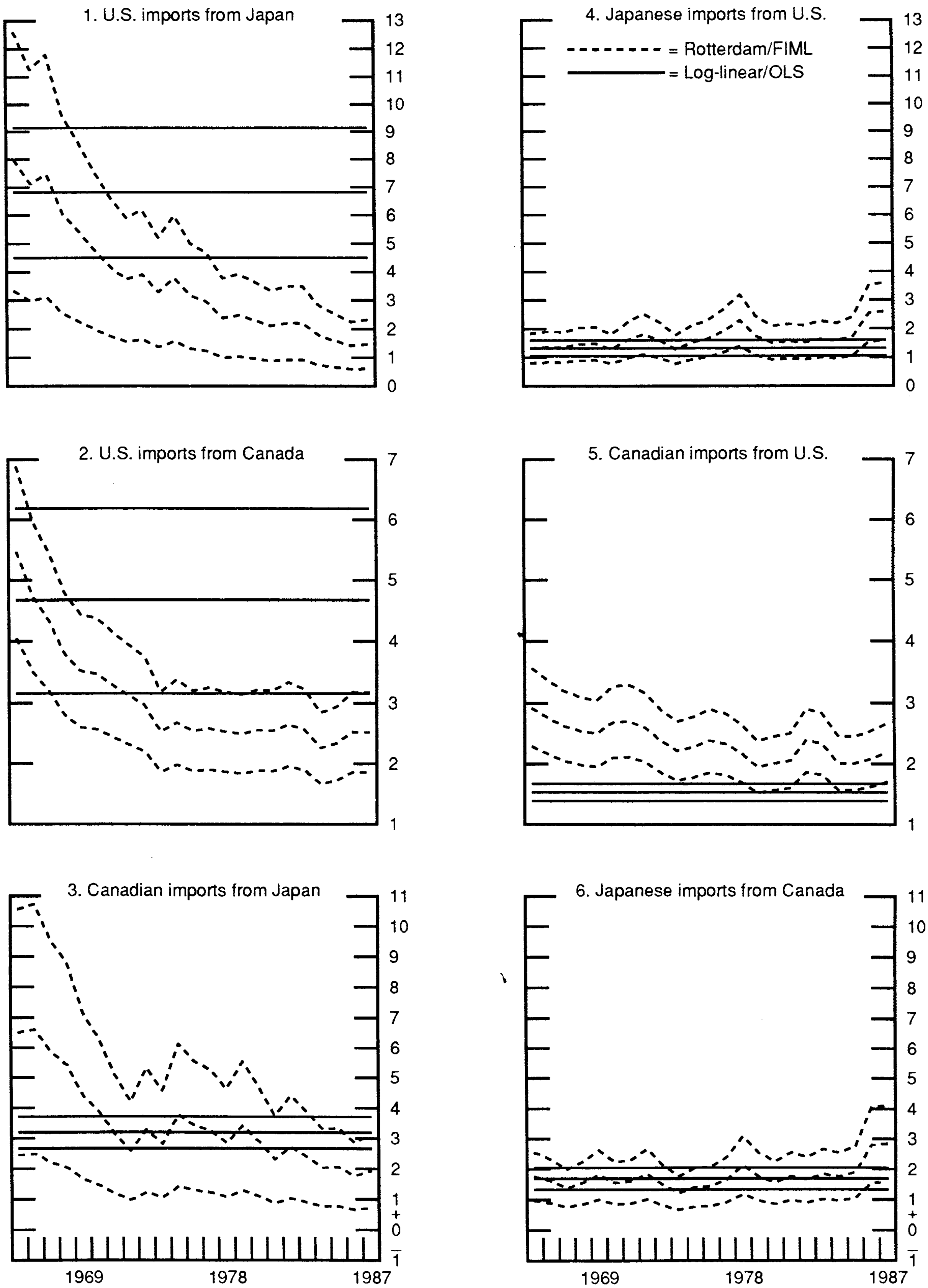
The estimated income elasticities for Canada also exhibit a secular tendency to decline (panels 3 and 5); income elasticities of Japan (panels 4 and 6), however, change little from 1965 to 1985 because the associated expenditure shares are virtually unchanged over this period (table 1). The subsequent decline in these shares raises income elasticities from 1.9 to 2.8 for Japanese imports from Canada and from 1.7 to 2.6 for Japanese imports from the United States. These results emphasize the sensitivity of trade elasticities to changes in expenditures shares -- that is, the elasticity estimates from the Rotterdam model are not constrained to follow any particular path because this model does not restrict the behavior of expenditure shares (other than adding up to one) to follow predetermined paths: Changes in the allocation of expenditures will change the corresponding elasticities.

The elasticity estimates associated with eq. (1) share with the literature (see appendix A) the choice of both functional form and estimation method but differ in three respects: the use of bilateral prices, the allowance of cross-price effects, and the presence of domestic absorption in real terms instead of real GNP. Inspection of the estimates reveals that the income elasticities from the log-linear model are significantly greater than one (figure 3), suggesting that this formulation is inconsistent with utility maximization. Moreover, for trade between the United States and both Canada and

Figure 3

Bilateral Income Elasticities

95% Confidence Intervals



Japan, the gap in elasticity estimates between the Rotterdam and log-linear models is statistically significant. In other words, treating income elasticities as autonomous entails a loss of information--namely, the information stemming from ignoring changes in spending patterns.[23]

To emphasize that the choice between the estimates of the Rotterdam model and those of the log-linear model is not solely a statistical matter, I re-examine the asymmetry in income elasticities noted by Houthakker and Magee (1969).[24] This asymmetry has both played an important role in studying international interactions and shaped much of the econometric work on trade for the last two decades.[25] According to the Rotterdam model, asymmetries in income elasticities for U.S. trade were quite important until the mid-1970s but have virtually vanished since then.[26] For example, the asymmetry in income elasticities for trade between Japan and the United States declines from -7.0 (=1.0 for exports minus 8.0 for imports) in 1965 to 1.0 in 1987; a similar reversal takes place for trade between the United States and Canada. The estimates of the log-linear model also replicate the asymmetry in elasticities noted by Houthakker and Magee but these estimates require for this asymmetry to persist despite changes in the composition of expenditures. Finally, figure 3 highlights the risks of relating the estimates from (1) to the structure of preferences. Specifically, the elasticity estimates from (1) conceal the similarities that exist between Japan and the United States in marginal budget shares for both their own and for each other's products (see table 2). Overall, the evidence

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[23] Based on the residuals of these models (figure C.1, appendix C), these differences in elasticity estimates are not due to differences in predictive accuracy.

[24] Houthakker and Magee find that the income elasticity for U.S. exports is smaller than the income elasticity for U.S. imports. As a result, if all countries were to grow at the same rate, then the United States would develop a growing trade deficit which must, eventually, lead to either a depreciation of the real exchange rate or a slowdown in the rate of growth for the United States. Although Stern (1989) offers a collection of essays examining trade interactions among Canada, Japan and the United States, that collection does not examine the issue of asymmetries in trade elasticities.

[25] For example, based on the *Social Science Citation Index*, Houthakker and Magee (1969) is cited 213 times from 1972 (the first year of citations for this paper) to 1990. The annual average number of citations is 11.2 with a standard deviation of 3.

[26] From a historical standpoint, these results suggest that Houthakker and Magee (1969) would have found an asymmetry in income elasticities even if they had not treated elasticities as autonomous and prices as given. Their finding, however, would not have been interpreted by the ensuing literature as an estimation anomaly or an immutable feature of the structure of international trade but rather as a manifestation of its evolution.

suggests that the choice of elasticities as autonomous parameters has important consequences for understanding the role of trade in transmitting disturbances from one country to another.

6. Conclusions

This paper estimates trade elasticities without relying on the restrictive assumptions embodied in most of the empirical work in trade of the last fifty years: exogeneity of prices, 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 expenditures in Canada, Japan, and the United States. Spending behaves according to the Rotterdam model which, 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 and simultaneity undermines our ability to address key questions such as the responsiveness of trade imbalances to changes in income. These findings, however, ignore intertemporal considerations and aggregation biases arising from the product composition of trade. These limitations are important but the process of eliminating them strengthens the point of the paper: that optimization and simultaneity are just as important in practice as in theory.

Appendix A: Chronology of Empirical Studies of Imports

This appendix offers a chronology of papers reporting econometric estimates of trade elasticities for import demand of Canada, Japan, and the United States, among other countries. In assembling this chronology, I have benefited from the surveys of Stern et al. (1976), Goldstein and Khan (1985), and Kohli (1991). Table A1 classifies existing studies using two considerations: whether the associated formulations embody the implications of optimization and whether the estimation method is single-equation, limited information, or full-information. The implications from optimization are adding-up, homogeneity, and symmetry of cross-price effects. Following the taxonomy of Intriligator (1978, p.373), I classify studies treating prices exogenously as using single-equation estimation methods. Similarly, studies treating prices endogenously but excluding the information pertaining to the model explaining prices are classified as using limited-information estimation methods. Finally, studies incorporating the model determining prices for parameter estimation are classified as using full-information methods.

Overall, three-quarters of the analyses rely on single-equation estimation methods that abstract from the implications of optimization (table A1). The remaining one quarter is split evenly between studies recognizing simultaneity and studies incorporating optimization; studies recognizing the statistical implications of both optimization and simultaneity are not available, which is the gap filled by this paper. The chronology also reveals that of the 95 papers that abstract from optimization, 74 of them treat trade elasticities as invariant to changes in the composition of expenditures (i.e. autonomous).

Table A2 lists specific studies and the most relevant features of their design for this paper: trade disaggregation, functional form, estimation method, and the presence of cross-price effects.[27] The appendix also reports whether the data set includes bilateral or multilateral prices, the frequency of observation, and the period covered in estimation. The mnemonics used in abbreviating these features are listed below:

Functional Form	*CES*: Constant Elasticity of Substitution; *CRESH*: Constant Ratios of Elasticity of Substitution Homogenous-Homothetic; *SGM*: Symmetric Generalized McFadden. *Linear:* Linear in the variables. *Log-linear:* Linear in the logarithms of the variables.
Estimation Method	*GLS*: Generalized Least Squares; *IV*: Instrumental Variables; *FIML*: Full Information Maximum Likelihood; *ILS*: Indirect Least Squares; *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.
Countries	C: Canada, G: Germany, J: Japan, U: United States, UK: United Kingdom, DC: Developed countries; O: the associated study reports estimates for Other countries.
Data Frequencies	A: Annual; Q: Quarterly; S: Semi-annual.

[27] Thus by design, this chronology excludes papers studying the structure of 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).

Table A1

Classification of Econometric Studies According to their Allowance for Optimization and Simultaneity: 1941-1991[1]

Implications from Simultaneity[3]	Implications from Optimzation[2]: Absent	Allowed	Total
Absent	81	13	94
Limited Allowance	11	2	13
Full Allowance	3	None	3
Total	95	15	110

[1] Entries in this table represent number of studies; the data for this table are the chronology of the 110 studies listed in appendix A.

[2] The implications from optimization are said to be absent if the estimates do not allow for adding-up, homogeneity, and symmetry of cross-price effects.

[3] The implications from simultaneity are said to be *absent* if prices are treated as exogenous for parameter estimation; the estimation has *limited allowance* for simultaneity if prices are treated as endogenous but the estimation ignores the model explaining prices; parameter estimates *fully allow* for simultaneity if they explicitly take into account the model determining prices.

Table A2
Chronology of Selected Studies for Imports of Selected Countries:
Canada, Japan, Germany, and the United States

Study	Trade Disaggregation	Functional Form	Estimator/ Price Behavior	Cross-Price Effects	Countries	Price Data	Frequency; Sample Range
1. de Vegh (1941)	Countries	Linear and log-linear	OLS Exog.	No	C, U	No Price Effects	A; 1919-38
2. Adler (1945)	Dutiable and Duty-free	Linear	OLS Exog.	No	U	Multilateral	A; 1922-37
3. Hinshaw (1945)	Countries	Linear	OLS Exog.	No	U	Multilateral	A; 1922-37
4. Chang (1946)	Countries and Commodities	Log-linear	OLS Exog.	No	C, J, U, O	Multilateral	A; 1924-38
5. Tinbergen (1949)	Commodities	Log-linear	OLS Exog.	No	U	Multilateral	A; 1879-1914
6. Adler, Schlesinger, and Westerborg (1952)	Countries and Commodities	Linear and log-linear	OLS Exog.	No	U	Bilateral	A; 1923-37
7. Harberger (1953)	Commodities	Log-linear	ILS Endo.	No	U	Multilateral	A; 1923-39
8. Lovasy and Zassenhaus (1953)	Commodities	Linear	OLS Exog.	No	U	Multilateral	Q; 1928-38 and 1947-52
9. Neisser (1953)	Commodities	Linear	OLS Exog.	No	U	Multilateral	A; 1919-37
10. Neisser and Modigliani (1953)	Commodities	Linear	OLS Exog.	No	U, O	Multilateral	A; 1925-39
11. Liu (1954)	Countries and Commodities	Linear	OLS Exog.	Yes	U	Bilateral	A; 1923-37
12. Polak (1954)	Commodities	Linear	OLS Exog.	No	C, J, U, O	Multilateral	A; 1923-37
13. Sasaki (1959)	Countries and Commodities	Log-linear	OLS Exog.	No	U, J	Bilateral	A; 1950-56
14. Kreinin (1960)	Commodities	Linear	OLS Exog.	No	U	No Price Effects	Q; 1947-58
15. Ball and Marwah (1962)	Commodities	Linear	OLS Exog.	No	U	Multilateral	A; 1948-58
16. Krause (1962)	Commodities	Log-linear	OLS Exog.	No	U	Cross-sec.	A; 1947-58
17. Reimer (1964)	Countries and Commodities	Linear	OLS Exog.	Yes	U	Bilateral	A; 1923-60

Table A2 (continued)

Study	Trade Disaggregation	Functional Form	Estimator/ Price Behavior	Cross-Price Effects	Countries	Price Data	Frequency; Sample Range
18. Rhomberg and Boissonneault (1964)	Countries	Linear	OLS Exog.	No	U, O	Multilateral	A; 1948-62
19. Davis (1966)	Countries	Linear	OLS Exog.	No	C, U, O	Multilateral	A; 1948-62
20. Kreinin (1967)	Countries and Commodities	Log-linear	OLS Exog.	No	C, J, U, O	Cross-sec. Multilateral	A; 1955-75 and A; 1954-64
21. Branson (1968)	Commodities	GNP Shares	OLS Exog.	No	U	Multilateral	Q; 1955-66
22. Floyd and Hynes (1968)	None	Log-linear	OLS Exog.	No	U	Multilateral	A; 1925-39 and 1949-62
23. Heien (1968)	None	Log-linear	OLS Exog.	No	C,U,O	Multilateral	A; 1951-65
24. Robinson (1968)	Commodities	Linear	OLS Exog.	No	C	Multilateral	Q; 1952-65
25. Adams, Eguchi, and Meyer-zu-Schlochtern (1969)	None	Linear	OLS Exog.	No	C, J, U, O	Multilateral	A; 1955-65
26. Detomasi (1969)	Countries and Commodities	Log-linear	ILS Exog.	No	C, U	Cons. Prices	A; 1948-65
27. Houthakker and Magee (1969)	Countries and Commodities	Log-linear	OLS Exog.	No	C, J, U, O	Multilateral	A; 1951-66
28. Officer and Hurtubise (1969)	Countries and Commodities	Linear	OLS Exog.	No	C, U	Bilateral	Q; 1953-65
29. Artus (1970)	None	Log-linear	OLS Exog.	No	G	Multilateral	A; 1955-69
30. Barten (1971)	Countries	Log-linear	NLS Exog.	No	Western Europe	Multilateral	A; 1959-67
31. Gregory (1971)	None	CES	OLS Exog.	No	U	Multilateral	Q; 1948-68
32. Marston (1971)	Commodities	Log-linear	OLS Exog.	No	UK	Multilateral	Q; 1955-67
33. Hooper (1972)	Countries and Commodities	Log-linear	OLS Exog.	Yes	U	Multilateral	Q; 1954-70
34. Kwack (1972)	Commodities	Linear	OLS Exog.	No	U	Multilateral	Q; 1960-67
35. Magee (1972)	Commodities	Log-linear	OLS Exog.	Yes	U	Bilateral	A; 1951-69

Table A2 (continued)

Study	Trade Disaggregation	Functional Form	Estimator/ Price Behavior	Cross-Price Effects	Countries	Price Data	Frequency; Sample Range
36. Price and Thornblade (1972)	Countries and Commodities	Log-linear	OLS Exog.	Yes	U	Bilateral	Q; 1964-69
37. Kreinin (1973)	Countries and Commodities	Log-linear	OLS Exog.	No	U, O	Multilateral	Q; 1964-70
38. Resnick and and Truman (1973)	Countries	Log-linear	OLS Exog.	No	Western Europe	Bilateral	A; 1953-68
39. Taplin (1973)	Commodities	Log-linear	OLS Exog.	No	C, J, U, O	Multilateral	A; 1953-70
40. Burgess (1974)	None	Translog	I3SLS Exog.	Yes	U	Multilateral	A; 1947-68
41. Clark (1974)	Commodities	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1963-73
42. Leamer (1974)	Commodities	Log-linear	Bayes Exog.	No	C, U, O	No Price Effects	Cross-sec. 1958
43. Miller and Fratianni (1974)	None	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1956-72
44. Ahluwalia and Hernández-Catá (1975)	Commodities	Log-linear	ILS Endo.	No	U	Multilateral	Q; 1960-73
45. Joy and Stolen (1975)	None	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1953-71
46. Khan and Ross (1975)	None	Log-linear	OLS Exog.	No	C, J, U, O	Multilateral	S; 1960-72
47. Yadav (1975)	Commodities	Log-linear	OLS Exog.	No	C	Multilateral	Q; 1956-72
48. Beenstock and Minford (1976)	None	Log-linear	OLS Exog.	No	C, J, U, O	Multilateral	Q; 1955-71
49. Hooper (1976)	Commodities	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1956-75
50. Marwah (1976)	Countries	Trade Shares	OLS Exog.	No	C, J, U, O	Multilateral	A; 1960-69
51. Murray and Ginman (1976)	None	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1961-68
52. Richardson (1976)	Commodities	Log-linear	ILS Exog.	No	U	Multilateral	A; 1958-71
53. Berner (1977)	Countries and Commodities	Rotterdam	GLS Exog.	Yes	G	Bilateral	A; 1954-70

Table A2 (continued)

Study	Trade Disaggregation	Functional Form	Estimator/ Price Behavior	Cross-Price Effects	Countries	Price Data	Frequency; Sample Range
54. Geraci and Prewo (1977)	None	Log-linear Gravity	GLS Exog.	No	C, J, U, O	No Price Effects	Cross-sec. 1970
55. Khan and Ross (1977)	None	Log-linear	OLS Exog.	No	C, J, U, O	Multilateral	Q; 1960-72
56. Mutti (1977)	Commodities	Log-linear	OLS Exog.	Yes	U	Multilateral	A; 1958-72
57. Yadav (1977)	Commodities	Log-linear	OLS Exog.	No	C	Multilateral	Q; 1956-73
58. Clements and Theil (1978)	Commodities	Rotterdam	OLS Exog.	Yes	DC	Multilateral	A; 1971-75
59. Deppler and Ripley (1978)	Commodities	Log-linear	OLS Exog.	No	C, J, U, O	Multilateral	S; 1964-76
60. Hooper (1978)	Commodities	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1955-77
61. Goldstein and Khan (1978)	None	Log-linear	FIML Endo.	No	G, J, U, O	Multilateral	Q; 1955-70
62. Lawrence R. (1978)	Commodities	Log-linear	OLS Exog.	No	U	Multilateral	S; 1962-77
63. Roningen (1978)	None	Log-linear Gravity	OLS Exog.	No	C, J, U, O	No Price Effects	Pooled-Cross-sec A; 1967-73
64. Theil and Clements (1978)	Commodities	Rotterdam	ML Exog.	Yes	U	Multilateral	A; 1921-70
65. Appelbaum and Kohli (1979)	None	Generalized Leontief	SUR Exog.	Yes	C	Bilateral	A; 1951-72
66. Fair (1979)	None	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1954-74
67. Stern, Baum, and Green (1979)	None	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1973-76
68. Stone (1979)	Countries and Commodities	Log-linear	2SLS Endo.	No	U	Bilateral	S; 1963-72
69. Wilson and Takacs (1979)	Countries	Log-linear	OLS Exog.	No	C, J, U, O	Multilateral	Q; 1957-71
70. Goldstein, Khan, and Officer (1980)	None	Log-linear	OLS Exog.	Yes	C, U, O	Multilateral	A; 1950-73
71. Leamer (1981)	Commodities	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1947-78

Table A2 (continued)

Study	Trade Disaggregation	Functional Form	Estimator/ Price Behavior	Cross-Price Effects	Countries	Price Data	Frequency; Sample Range
72. Geraci and Prewo (1982)	Countries	Log-linear	OLS Exog.	Yes	G, J, U, O	Bilateral	Q; 1958-74
73. Grossman (1982)	Commodities	Log-linear	2SLS Endo.	Yes	U	Bilateral	Q; 1968-78
74. Haynes and Stone (1983a)	None	Log-linear	Spectral Exog.	No	U	Multilateral	Q; 1955-79
75. Haynes and Stone (1983b)	None	Log-linear	IV Endo.	No	U, O	Multilateral	Q; 1947-79
76. Ueda (1983)	None	Log-linear	IV Endo.	No	J	Multilateral	S; 1966-80
77. Warner and Kreinin (1983)	Commodities	Log-linear	OLS Exog.	No	C, J, U, O	Multilateral	Q; 1957-70
78. Husted and Kollintzas (1984)	Commodities	Linear	IV Endo.	No	U	Multilateral	Q; 1954-80
79. Thursby and Thursby (1984)	None	Linear and log-linear	OLS Exog.	Yes	C, J, U, O	Multilateral	Q; 1955-78
80. Aw and Roberts (1985)	Countries	Translog	I3SLS Exog.	Yes	U	Multilateral	A; 1960-80
81. Kohli (1985)	Countries	Translog	SUR Exog.	Yes	U	Multilateral	A; 1960-79
82. Italianer (1986)	Countries and Commodities	CRESH	NLS Exog.	Yes	U, O	Bilateral	A; 1963-80
83. Italianer and d' Alcantara (1986)	Countries and Commodities	Log-linear	OLS Exog.	No	Western Europe	Bilateral	A; 1965-80
84. Shiells, Stern, and Deardorff (1986)	Commodities	Log-linear	2SLS Endo.	Yes	U	Multilateral	A; 1962-78
85. Lawrence R. (1987)	Commodities	Log-linear	OLS Exog.	No	J	No Price Effects	Pooled-Cross-sec 1970, 1980, 1983
86. Baldwin (1988)	Non-oil	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1967-87
87. Brada and Méndez (1988)	Countries	Log-linear Gravity	OLS Exog.	No	C, J, U, O	No Price Effects	Pooled-Cross-sec A; 1973-77
88. Cushman (1988)	Countries	Log-linear	OLS Exog.	Yes	U	Multilateral	Q; 1974-83
89. Helkie and Hooper (1988)	Commodities	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1969-84

Table A2 (continued)

Study	Trade Disaggregation	Functional Form	Estimator/ Price Behavior	Cross-Price Effects	Countries	Price Data	Frequency; Sample Range
90. Kohli and Morey (1988)	Countries	CES	NLS Exog.	Yes	U	Multilateral	Q; 1960-79
91. Marquez and McNeilly (1988)	Countries and Commodities	Log-linear	2SLS Endo.	No	C, J, U, O	Bilateral	Q; 1974-84
92. Parikh (1988)	Countries	Almost Ideal	ML Exog.	Yes	J, U	Multilateral	A; 1965-80
93. Brenton (1989)	Countries and Commodities	Almost Ideal	OLS Exog.	Yes	UK	Bilateral	A; 1961-82
94. Deyak, Sawyer, and Sprinkle (1989)	Commodities	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1958-83
95. Gagnon (1989)	Countries	Log-linear	ML Endo.	No	U	Multilateral	Q; 1973-85
96. Krugman (1989)	None	Log-linear	OLS Exog.	No	C, J, U, O	Multilateral	A; 1971-86
97. Lawrence D. (1989)	None	SGM	IV Endo.	Yes	C	Multilateral	A; 1961-80
98. Moffet (1989)	None	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1967-87
99. Noland (1989)	None	Log-linear	Grid/OLS Exog.	No	J	Multilateral	Q; 1970-85
100. Shiells (1989)	Countries and Commodities	Log-linear	2SLS Endo.	Yes	U	Multilateral	A; 1963-78
101. Summary (1989)	Countries	Log-linear Gravity	OLS Exog.	No	U	No Price Effects	Cross-sec. 1978 and 1982
102. Cushman (1990)	Countries	Log-linear	OLS Exog.	Yes	U	Multilateral	Q; 1974-83
103. Kohli (1990)	None	Translog	I3SLS Exog.	Yes	U	Multilateral	Q; 1948-87
104. Lawrence R. (1990)	Commodities	Log-linear	OLS Exog.	No	U	Multilateral	Q; 1976-90
105. Marquez (1990)	Countries	Log-linear	OLS Exog.	Yes	C, J, U, O	Multilateral	Q; 1973-85
106. Clarida (1991)	None	Log-linear	NLS Exog.	No	U	Multilateral	Q; 1968-90
107. Kohli (1991)	Countries and Commodities	Flexible Forms	IV Endo.	Yes	U	Multilateral	Q; 1948-87

Table A2 (continued)

Study	Trade Disaggregation	Functional Form	Estimator/ Price Behavior	Cross-Price Effects	Countries	Price Data	Frequency; Sample Range
108. Lawrence R. (1991)	Commodities	Shares	OLS Exog.	No	J	No Price Effects	Cross-sec. 1985
109. Marquez (1991a)	Countries	Log-linear	FIML Endo.	Yes	U	Multilateral	Q; 1973-84
110. Shiells (1991)	Commodities	Log-linear	2SLS Endo.	No	U	Multilateral	Q; 1978-88

Appendix B: 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; Population, Domestic Production, and Prices; and Factor Prices.

B.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 resulting in different prices for the same product. For parameter estimation, these prices are expressed in terms of local currency using bilateral dollar exchange rates and local tariffs (see below). 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, $i=0,..., n-1$) are estimated as the dollar value of trade flows deflated by the corresponding bilateral dollar price and are expressed on a per-capita basis.

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). Bilateral rates among these three countries are based on the corresponding bilateral dollar exchange rates.

Tariffs. Data for the multilateral tariff rate are constructed as the ratio between custom duties and the total value of imports, both in local currency. The sources for data on custom duties are listed below:

Canada

For 1964-75: **Historical Statistics of Canada,** *Statistics Canada*, 1983, series G479.
For 1976-77: **Government Finance Statistics Yearbook,** *International Monetary Fund*, 1980, p. 156, series 6.1.
For 1978-87: **Government Finance Statistics Yearbook,** *International Monetary Fund*, 1989, p. 194, series 6.1.

Japan

For 1960-64: **Japan Statistical Yearbook,** *Bureau of Statistics*, 1968, table 327.
For 1965-67: **Japan Statistical Yearbook,** *Bureau of Statistics*, 1969, table 320.
For 1968-69: **Japan Statistical Yearbook,** *Bureau of Statistics*, 1970, table 319.
For 1970-77: **Government Finance Statistics Yearbook,** *International Monetary Fund*, 1980, p. 296, series 6.1.
For 1978-87: **Government Finance Statistics Yearbook,** *International Monetary Fund*, 1989, p. 371, series 6.1.

United States

Survey of Current Business, U.S. Department of Commerce, National Income and Product Accounts (NIPA), International Transactions, NIPA, table 3.2.

Data for tariffs of Japan and Canada are expressed, in these sources, in terms of fiscal calendars that begin April 1. To convert these figures to calendar year, I assume that tariff collection is uniform over the year and construct tariff revenues of calendar-year t as three quarters of the fiscal-year t plus one-quarter of the tariff revenue of the fiscal-year t-1.

B.2 Population, Domestic Production, and Prices

Population. The sources for data on population are listed below:

Canada	**Bank of Canada Review**, *Bank of Canada,* table H1.
Japan	**Economic Statistics Annual,** Bank of Japan, table 156.
United States	**Employment and Earnings**, U.S. Department of Labor (Bureau of Labor Statistics), table 1.

Domestic Production. 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. 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. 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. The data sources for these four variables are listed below:

Nominal GNP:

Canada	**Canadian Economic Observer**, *Statistics Canada,* table 1.1.
Japan	**Economic Statistics Monthly,** *Bank of Japan,* table 123.
United States	**Survey of Current Business**, *U.S. Department of Commerce,* NIPA, table 1.1.

Nominal value of exports:

Canada	**Canadian Economic Observer**, *Statistics Canada,* table 1.2.
Japan	**Economic Statistics Monthly,** *Bank of Japan,* table 123.
United States	**Survey of Current Business**, *U.S. Department of Commerce,* NIPA, table 1.1.

Real GNP:

Canada	**Canadian Economic Observer**, *Statistics Canada,* table 1.3.
Japan	**Economic Statistics Monthly,** *Bank of Japan,* table 123.
United States	**Survey of Current Business**, *U.S. Department of Commerce,* NIPA, table 1.2.

Real value of exports:

Canada	**Canadian Economic Observer**, *Statistics Canada,* table 1.3.
Japan	**Economic Statistics Monthly,** *Bank of Japan,* table 123.
United States	**Survey of Current Business**, *U.S. Department of Commerce,* NIPA, table 1.2.

B.3 Factor Prices

The wage rate for the United States is constructed as the ratio between Compensation of Employees (**Survey of Current Business,** *U.S. Department of Commerce,* table 1.14) and employment (**Employment and Earnings,** *U.S. Department of Labor,* table 1). 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).

The rental rate of capital is measured as

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

where $pk_t = (1-\tau_{pt} - \tau_{ct})[\, l_t - \kappa - \mathrm{dln}pi_t]$, τ_{pt} is the average personal tax rate constructed as the ratio between personal income tax receipts and personal income; τ_{ct} is the average corporate tax rate measured as the ratio between corporate profit tax accruals and corporate income before taxes; l_t is the long-term market yield on government securities; pi_t is the deflator for non-residential investment; and κ is the depreciation rate of the non-residential capital stock which I treat as constant and equal to 6.0 percent for Canada, 10.6 percent for Japan, and 9.6 percent for the United States. These estimates come from the FRB-Multicountry model.

Personal income taxes:

Canada	**Canadian Economic Observer**, *Statistics Canada*, table 1.7.
United States	**Survey of Current Business**, U.S. Department of Commerce, NIPA, table 3.2.

Corporate taxes:

Canada	**Canadian Economic Observer**, *Statistics Canada*, table 1.7.
United States	**Survey of Current Business**, *U.S. Department of Commerce*, NIPA, table 3.2.

Corporate plus personal taxes for Japan:

Economic Statistics Monthly, *Bank of Japan*, table 180.

Personal income

Canada	**Canadian Economic Observer**, *Statistics Canada*, table 1.1.
Japan	**Annual Report on National Account**, *Economic Planning Agency*, Account 5.
United States	**Survey of Current Business**, *U.S. Department of Commerce*, NIPA, table 1.14.

Corporate income

Canada	**Canadian Economic Observer**, *Statistics Canada*, table 1.1.
Japan	**Annual Report on National Account**, *Economic Planning Agency*, Account 5.
United States	**Survey of Current Business**, *U.S. Department of Commerce*, NIPA, table 1.14.

Interest rates

Canada	Market yield on Government Bellwether bonds with 10 year maturity; **Statistical release H13**; *Board of Governors of the Federal Reserve System.*
Japan	Market yield on Government Bellwether bonds with 10 year maturity; **Statistical release H13**; *Board of Governors of the Federal Reserve System.*
United States	Market yield on U.S. Treasury securities with 10 year maturity; **Statistical release H15**; *Board of Governors of the Federal Reserve System.*

Deflator for non-residential investment

Canada	**Canadian Economic Observer**, *Statistics Canada*, tables 1.2 and 1.3.
Japan	**Economic Statistics Monthly,** Bank of Japan, table 123.
United States	**Survey of Current Business**, U.S. Department of Commerce, NIPA, tables 1.1 and 1.2.

Appendix C: Econometric Issues

This appendix shows the prediction errors for the bilateral trade flows shown in figure 3 and indicates how ignoring changes in the composition of expenditures affect the least-squares estimates of trade elasticities.

C.1 The Econometric Implications of Changes in Spending Patterns for Elasticity Estimation

To show how changes in the composition of expenditures affect the least-squares estimates of trade elasticities, I treat the eqs. in (3) as though they were a collection of Engle curves:

(C.1) $$w_{it}\tilde{q}_{it} = \alpha_i\tilde{y}_t + u_{it}, \quad u_t \sim N(0, \sigma^2)$$

where $\tilde{q}_{it} = \text{dln}q_{it}$, $\tilde{y}_t = \text{dln}(y / P)_t$. To simplify the presentation, I focus on the results for a single commodity and drop the subscript i. Given (C.1), the elasticity of q with respect to y is $\eta_t = \alpha/w_t$, which responds to changes in w_t.

Applying least squares to (C.1) yields

(C.2) $$\hat{\alpha} = [\Sigma_t\, w_t\tilde{q}_t\tilde{y}_t] / [\Sigma_t\, \tilde{y}_t^2],$$

and the estimated elasticity-path is given by $\hat{\eta}_t = \hat{\alpha}/w_t$. If spending patterns are fixed, then the expression for $\hat{\alpha}$ is

(C.3) $$\hat{\alpha}^f = w^f\, [\Sigma_t\, \tilde{q}_t\tilde{y}_t] / [\Sigma_t\, \tilde{y}_t^2],$$

w^f is the fixed expenditure share and the estimated (constant) elasticity is $\hat{\eta}^f = \hat{\alpha}^f / w^f$.

To examine the implications of fixing spending patterns, subtract (C.3) from (C.2) to obtain

(C.4) $$\begin{aligned} \hat{\eta}_t w_t - \hat{\eta}^f w^f &= \hat{\alpha} - \hat{\alpha}^f \\ &= \{[\Sigma_t\, w_t\tilde{q}_t\tilde{y}_t] / [\Sigma_t\, \tilde{y}_t^2]\} - \{w^f\, [\Sigma_t\, \tilde{q}_t\tilde{y}_t] / [\Sigma_t\, \tilde{y}_t^2]\} \\ &= \{[\Sigma_t\, (w_t - w^f)\, \tilde{q}_t\tilde{y}_t]\} / [\Sigma_t\, \tilde{y}_t^2] \\ &= [\Sigma_t\, \Gamma_t\tilde{q}_t\tilde{y}_t] / [\Sigma_t\, \tilde{y}_t^2], \end{aligned}$$

where Γ_t is the gap between w_t and w^f. Re-arranging terms in (C.4) yields

$$\hat{\eta}_t = \hat{\eta}^f\, (w^f / w_t) + (1/w_t)\{[\Sigma_t\, \Gamma_t\tilde{q}_t\tilde{y}_t] / [\Sigma_t\, \tilde{y}_t^2]\},$$

which indicates that differences between $\hat{\eta}_t$ and $\hat{\eta}^f$ are small as long as *percentage* deviations between w_t and w^f are small.

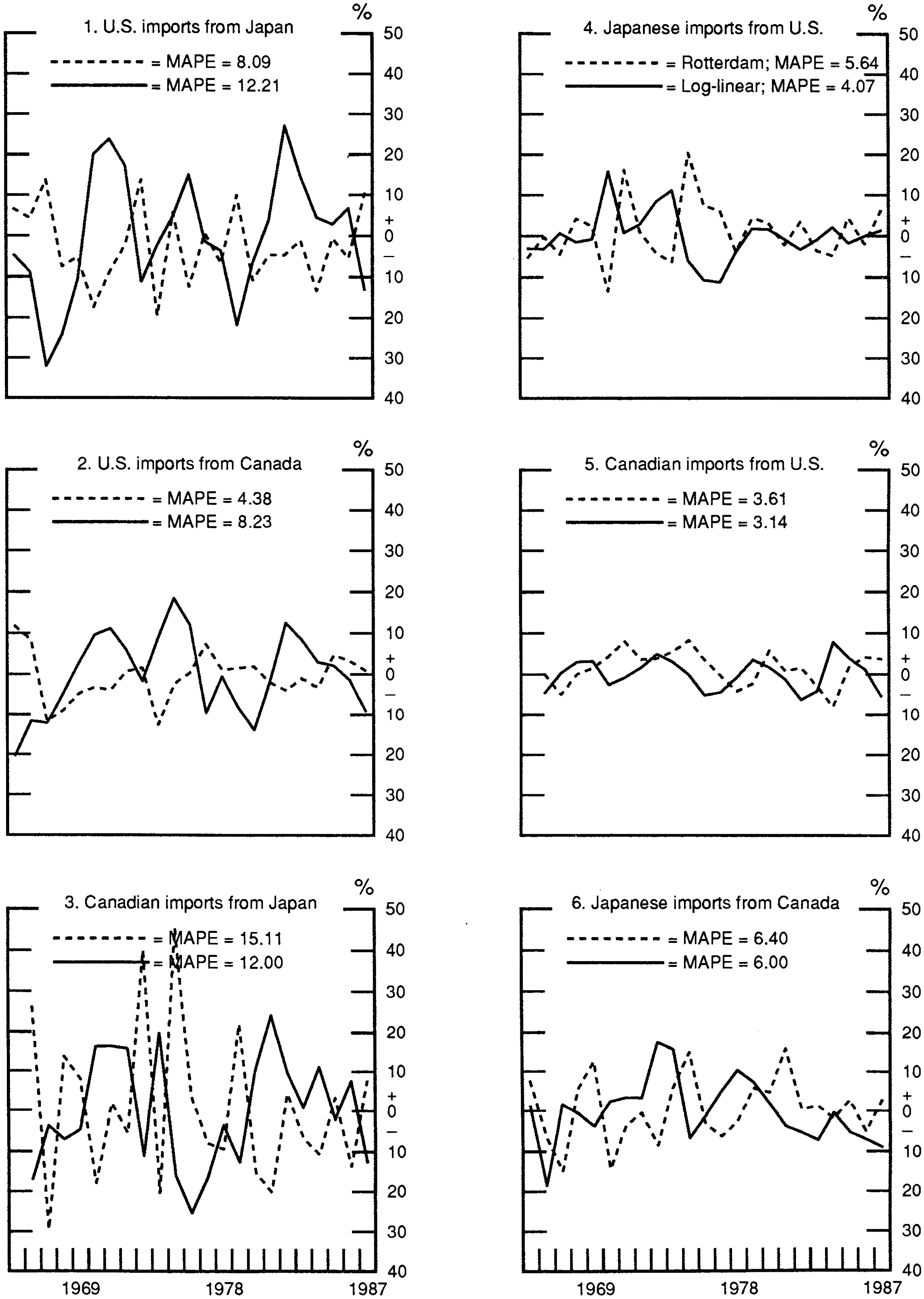
C.2 Prediction Errors for Bilateral Trade

Figure C.1 shows deviations between actual and predicted values, in percentage terms, and the associated Mean Absolute Percentage Error (MAPE) for each of the bilateral trade flows shown in figure 3. Specifically, percentage deviations are calculated as $(q_{it} - \hat{q}_{it})/q_{it}$ where the symbol "^" denotes the predicted value for purchases of the ith product. Given that the estimating equations are nonlinear in the variables, computing these errors involves re-expressing the associated equations into a suitable form. Thus the predicted values for the log-linear model are calculated as

Figure C.1

Prediction Errors for Bilateral Trade

Percent Deviation Relative to Actual



MAPE: Mean Absolute Percentage Error

$$\hat{q}_i = \exp(\hat{\alpha}_i + \hat{\eta}_i \ln(y / P) + \Sigma_j \hat{\varepsilon}_{ij} \ln p_j),$$

whereas the predicted values for the Rotterdam model involves solving

$$(\hat{q}_{it} p_{it} / y_t)(\ln \hat{q}_{it} - \ln q_{i,t-1}) = \hat{\mu}_i \, d\ln(y / P)_t + \sum_{j=0}^{n} \hat{\pi}_{ij} \, d\ln p_{jt}$$

for $\hat{q}_i$. Note that to ensure comparability with the OLS "residuals," I use 1-step ahead prediction errors also for the Rotterdam model.

Inspection of the results suggests that the prediction errors generated by the OLS estimates of the log-linear model are smaller than the FIML estimates of the Rotterdam model. But the differences in predictive accuracy are small suggesting that the gap in elasticity estimates shown in figure 3 is not due to a deterioration of explanatory power of the Rotterdam model.

C.3 Tests of Parameter Constancy

To examine parameter constancy, the analysis includes in each equation a dummy variable that equals one for the year being tested (zero otherwise). This procedure amounts to excluding the given year out of the estimation sample. Figures C.2-C.4 show the associated estimated marginal budget shares (μ_i's) and own-price effects (π_{ii}'s). Overall, the evidence supports the hypothesis of parameter constancy but there are a two exceptions in 1973 for Canadian imports from Japan and Germany.

C.4 Standard Errors of Eigenvalues

The standard errors of the eigenvalues are calculated using a Taylor expansion:

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

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 $\hat{\pi}$.

Figure C.2

Tests of Parameter Constancy: 95% Bands
Canada

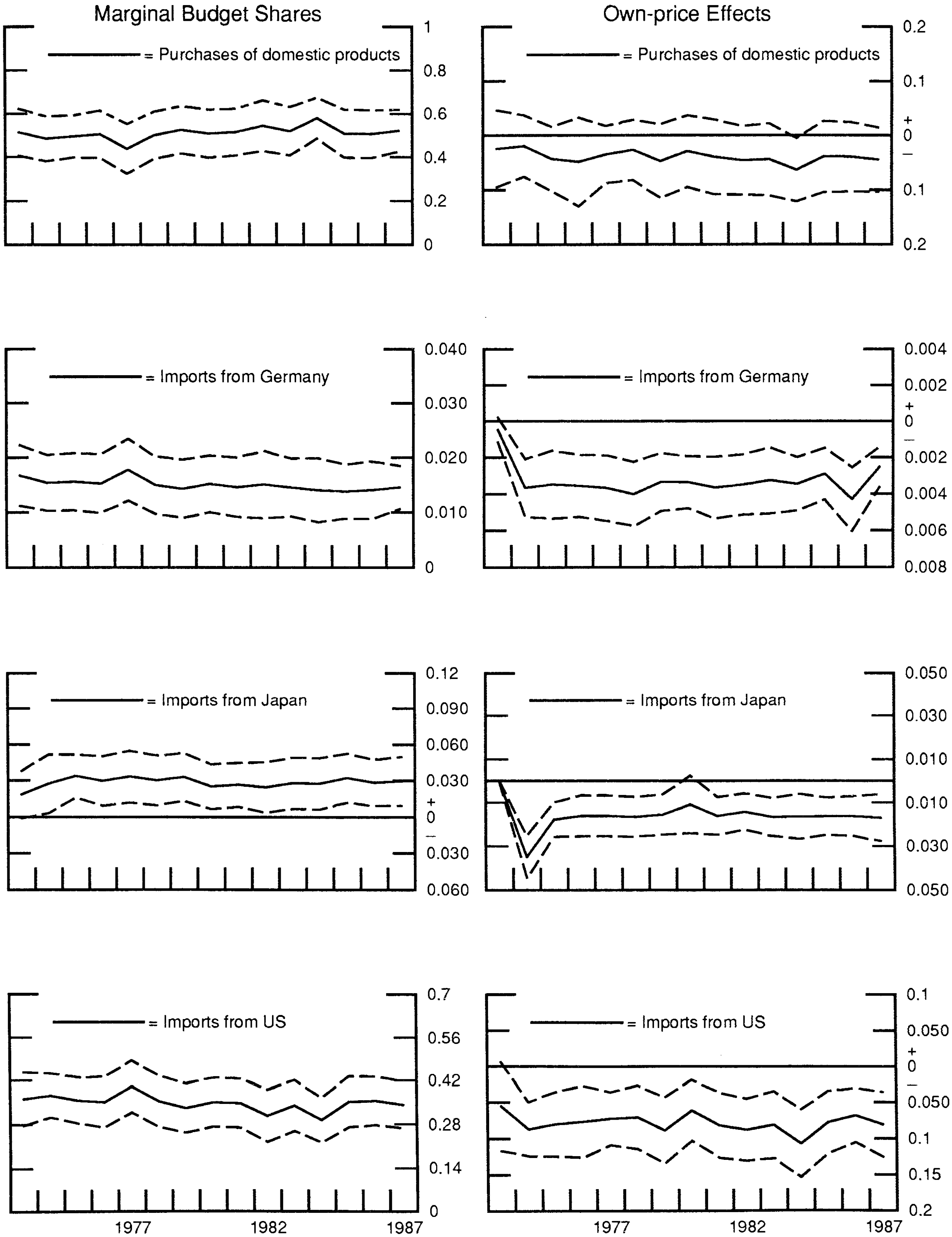


Figure C.3

Tests of Parameter Constancy: 95% Bands

Japan

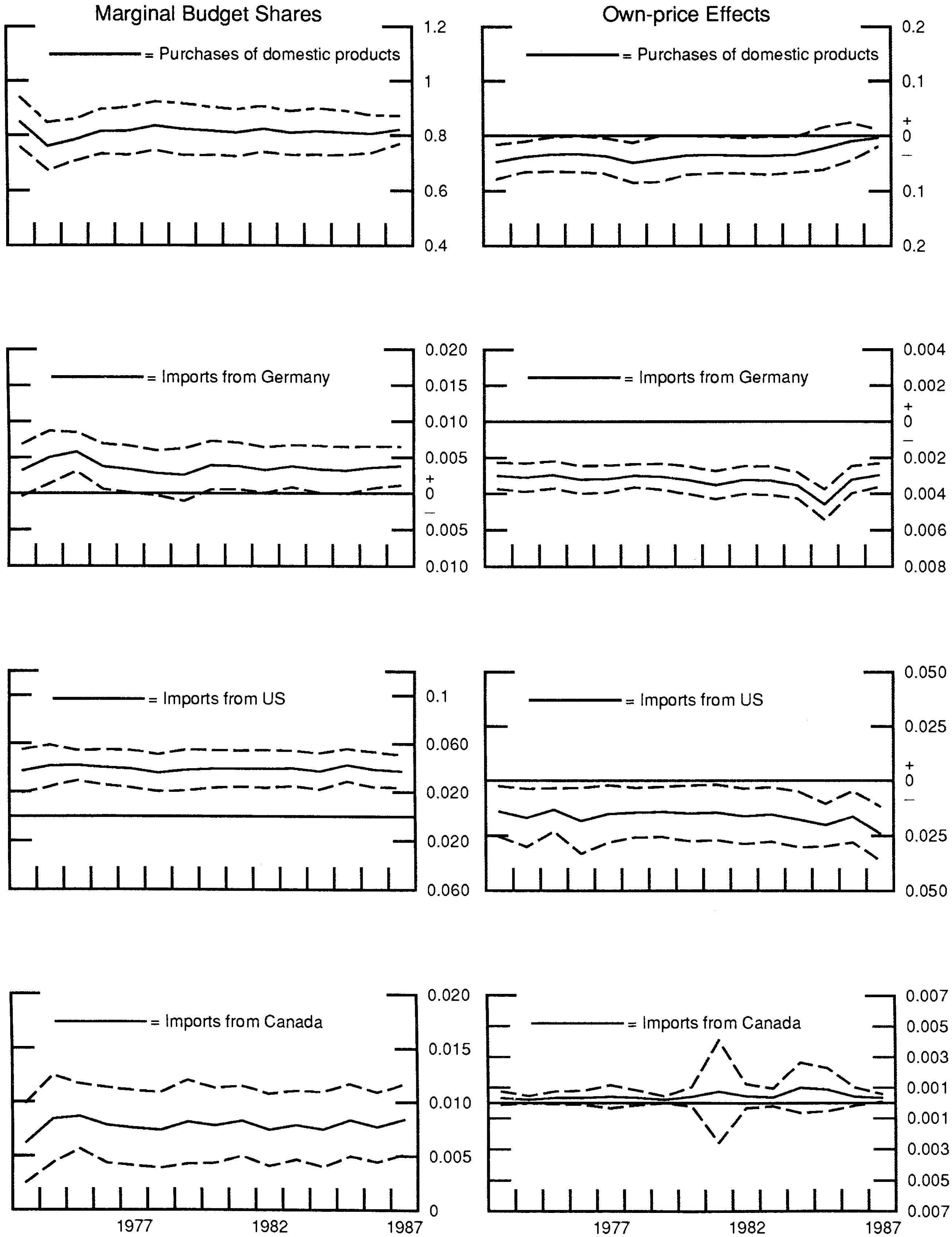
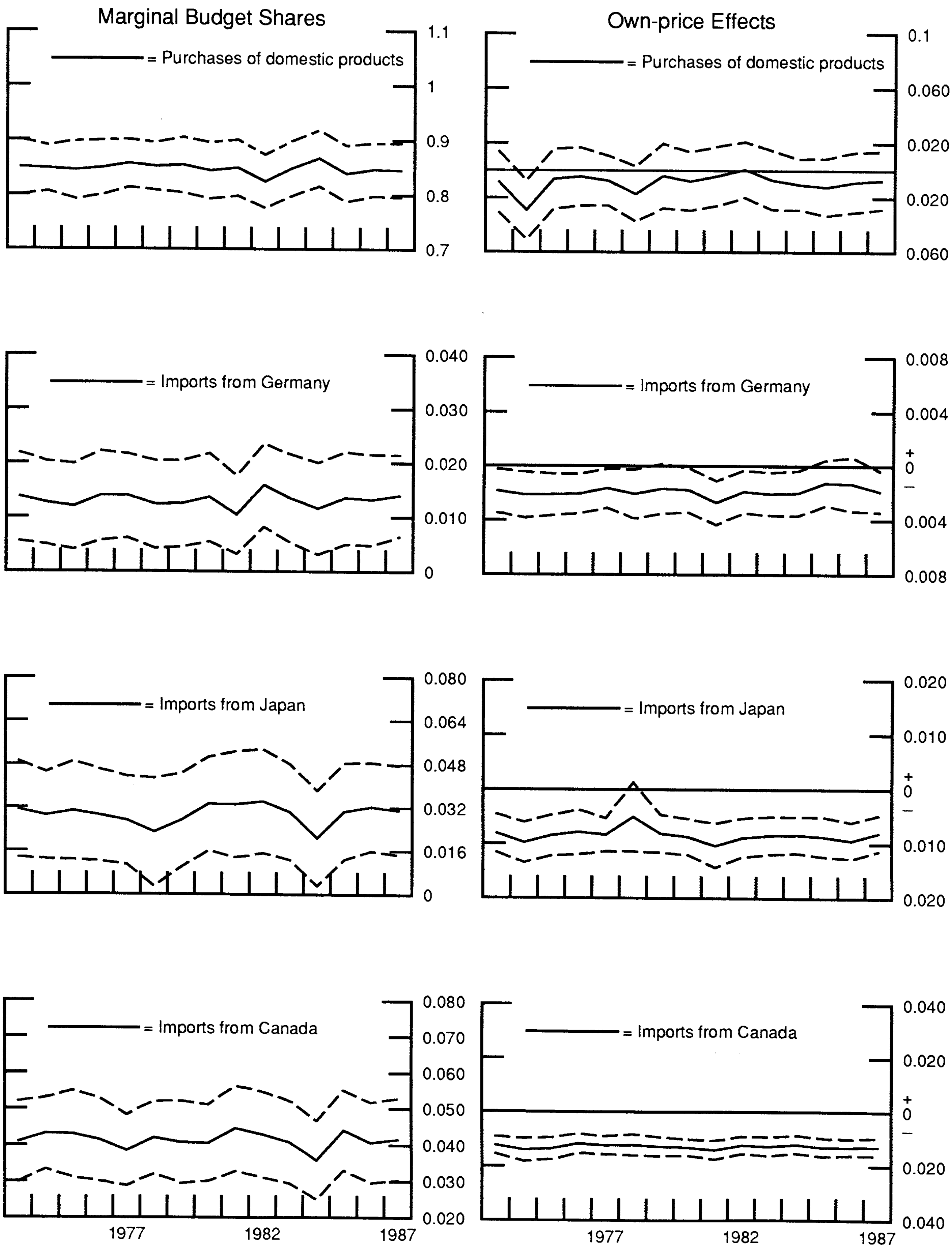


Figure C.4

Tests of Parameter Constancy: 95% Bands

U.S.



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