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Have U.S.-Japan Trade Agreements Made a Difference?

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The few existing empirical studies of U.S.-Japan trade agreements have relied primarily on descriptive statistics or univariate time series methods. We conduct a more powerful test by evaluating agreements in the context of well-specified econometric models. Consistent with trade theory, import demand is modeled as a cointegrating relationship with income and relative price variables, where a trade agreement may cause a structural break in the cointegrating vector. In several cases, we find evidence that market-opening trade agreements may have increased the volume of Japanese imports, while other agreements appear to have had no significant impact.

Key words: structural break tests; U.S.-Japan trade agreements; import promotion policies

JEL F13, F14, C32

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1. Introduction

International trade is governed by a complicated patchwork of international treaties, bilateral agreements, and unilateral actions, often directed at trade in very specific commodities. A determination of the impact of trade measures begins with an assessment of their direct effect on trade volumes and prices in targeted sectors. While in some cases—such as the recent U.S. anti-dumping duties on steel—the measures have readily identifiable impacts, in others cases the effects are more difficult to determine.

Examples of the latter are measures during the 1980s and 1990s to "open" Japanese markets to foreign products. For many years, foreign companies and governments have complained that Japanese markets were unfairly protected from foreign competition by a range of regulations, licensing requirements and government procurement policies that are not addressed by traditional trade liberalization measures.[1] Over the past two decades, the Japanese government entered into a number of bilateral trade agreements designed to increase import access. Among the most visible of the market opening policies were the so-called Voluntary Import Expansion policies, Japanese commitments (or targets, depending on whom you ask) to raise foreign presence in key Japanese markets. These included the U.S.-Japan semiconductor accord of 1986 and the auto parts agreements reached in 1992 and 1995. But, they also included licensing concessions, agreements on standards, and changes in government procurement designed to raise foreign import access.[2]

While the wide range of import expansion agreements looks impressive on paper, it is less clear how big an impact they have had in fact. Critics have charged that some measures are simply window dressing, intended to mollify foreign governments without requiring painful adjustments by Japanese firms. In other cases, the apparent magnitude of incentives appears too small to create incentives for change. Our objective is to look for empirical evidence of import expansion effects for several prominent industries that were subject to U.S.-Japan market-opening agreements.

There are several existing studies of U.S.-Japan trade policy that bear on this issue. These studies range from survey evaluations of American businesses operating in Japan to univariate statistical tests of import change to a few econometric studies. In this paper, we look for evidence that trade agreements altered import patterns using a model-based time series methodology. Our approach permits us to test for changes in import behavior within the context of a traditional econometric trade model. We estimate a cointegrating relationship describing the evolution of imports in each sector as a function of economic fundamentals, and test for a structural break in that relationship at the time of the policy implementation. An indication of a structural break following an import expansion agreement provides some evidence that the agreement may have significantly altered import behavior relative to the pre-intervention period.

The paper is organized as follows. In section 2, we briefly describe the U.S.-Japan market-opening trade agreements that are the focus of our research, and we review existing evidence on the effectiveness of such agreements. Section 3 presents our empirical model. Section 4 presents results, and concluding observations are given in section 5.

2. U.S.-Japan Market-Opening Trade Agreements

The American Chamber of Commerce in Japan (ACCJ) has catalogued and undertaken opinion surveys of the effectiveness of major negotiated policy changes affecting U.S. trade with Japan. ACCJ (1997) identified over 45 significant “market-opening” trade agreements between 1980 and 1996, a survey later extended (ACCJ, 2000) to include an additional 18 agreements made in subsequent years. We examine several of the most notable agreements reached prior to 1996, including autos, auto parts, tobacco, semiconductors, paper, medical products, and lumber. A timeline of these measures is given in Table 1. Among these, the Semiconductor Trade Agreement (STA) has received the most attention in the literature, because of its inclusion of explicit foreign market share targets.

While not comprehensive, the seven sectors considered here represent an appropriate sample for evaluating the impact of U.S.-Japan trade agreements. As a group, they constitute a significant share of

Japanese non-oil, non-food imports (rising from 7% in 1980 to 16% in 2000), and each has played an important role in the political economy debate on Japanese market access. Other industries targeted for import promotion have included computers and electronics, satellites, leather, apples, telecommunications, building and contracting, flat glass, beef and citrus, and a number of service areas. In most cases, applying the econometric methodology of this paper to these other sectors is complicated by a lack of sufficiently disaggregated high frequency data, or by the existence of other quantitative trade restrictions that may obscure detection of policy effects. See ACCJ (1997, 2000) for a complete description of U.S.-Japan trade agreements.[3]

In addition to the ACCJ opinion surveys of American companies doing business in Japan, Bayard and Elliot (1994) and Elliot and Richardson (1997) use data from the U.S. Trade Representative's office and interviews to evaluate the effects of actions taken under section 301 of the U.S. Trade Law.

Other researchers have used descriptive statistics to evaluate the success of trade agreements. Gold and Nanto (1991) compare post-agreement growth rates of U.S. exports to Japan with growth rates to the world over the 1985-1990 period and find that targeted sectors did grow faster. Greaney (2001) looks for evidence that a range of measures identified in the ACCJ study stimulated import growth by comparing pre- and post-implementation growth rates for imports in targeted sectors, and by comparing growth rates in targeted sectors to overall import performance. She finds no consistent pattern of real increases in Japanese imports of U.S. goods targeted for increases, although she does find some evidence that U.S.-Japan agreements benefited third countries.

The Semiconductor Trade Agreement has been the subject of considerable recent research. Theoretical studies (Greaney, 1996; Krishna, Roy, and Thursby, 1996) have suggested that the agreement may have facilitated collusive behavior by semiconductor producers, with possibly ambiguous effects on import volumes. Indeed most empirical studies have found that the fair market value requirements (FMVs) of the STA acted as a price support for Japanese exports (see Flamm, 1996; Dick, 1994; Irwin, 1994; Tyson, 1992). There is little consensus on whether the STA was effective in increasing the foreign market share. Several authors have argued (Sumita and Shin, 1996; Flamm, 1996; Bergsten and Noland, 1993) that although the STA may have had some effect on foreign market presence, other factors such as the shift in demand for those integrated circuits for which U.S. firms had a comparative advantage may have explained at least some of the gain in market share. However, Bergsten and Noland (1993) find in their calculations that not all of the shift towards foreign made chips can be attributed to changes in industry patterns alone and attribute the residual to the STA. Flamm (1996), using similar calculations, arrives at the same conclusion. Parsons (2002) estimates an import demand equation for semiconductors using cointegration techniques which simultaneously search for a possible endogenous break. No break is found.

We are aware of very few econometric studies of U.S.-Japan trade agreements. In the paper mentioned above, Greaney (2001) conducts univariate Chow tests for structural breaks in import time series. Noland (1997) uses a gravity model of U.S. trade patterns and finds little evidence of trade agreement impacts. Baker, Gross and Tower (1997) estimate trend models for several measures of U.S. export performance in wood products and find no evidence that Super 301 trade sanctions increased U.S. wood product exports to Japan.

A problem with univariate analyses of trade agreements is that it is difficult to interpret a detected break in the time series. As Greaney (2001) acknowledges, finding a break in an import series only indicates a regime shift, but can say nothing about the source of the observed change. In particular, import volumes can change dramatically if there are sharp changes in fundamental variables upon which they depend. A multivariate econometric model of trade is needed to capture these effects.

3. The Empirical Trade Model

Real import demand in industry *k* is modeled as a function of income and relevant prices. In log form and assuming homogeneity, we estimate,[4]

$$m_{k,t} = \alpha_k + \beta_{1,k} y_t + \beta_{2,k} p^m_{k,t} + \beta_{3,k} p^d_{k,t} + v_{k,t} \,, \qquad (1)$$

where m is real imports, y is real income and p^m and p^d are import and domestic industry prices relative to aggregate domestic prices, and v_t reflects measurement error.[5] Consumer theory predicts that the income and "other" price elasticities should be positive, and the relative import price elasticity will be negative.

Because imports and income (and perhaps relative prices) are likely integrated, OLS estimation of the model in levels may yield a spurious regression. One solution would be to estimate the model in differenced form to eliminate non-stationary regressors, but this approach may neglect important long-run relationships. Instead, we test for a cointegrating system among the model variables, and then proceed to use a consistent estimator due to Saikkonen (1991) and Stock and Watson (1993).[6,7] The dynamic ordinary least squares (DOLS) estimating equation is given by:

$$m_{k,t} = \alpha_k + \boldsymbol{\beta}_k \mathbf{x}_{k,t} + \sum_{i=0}^{p} \gamma_{k,-i} \Delta \mathbf{x}_{k,t+i} + \sum_{i=1}^{p} \gamma_{k,i} \Delta \mathbf{x}_{k,t-i} + v'_{k,t}\,, \tag{2}$$

where $\mathbf{x}_{k,t}$ is a vector of the independent variables: income and relative prices. The dependent variable in time t is regressed upon the independent regressors in time t plus an appropriate number of lead and lag differences of the independent variables (including the contemporaneous difference). The estimate for the parameter vector, $\boldsymbol{\beta}_k$, typically the main parameters of interest, is *super-consistent* so long as the system of I(1) variables are, in fact, cointegrated. The standard errors, however, must be properly scaled upward.

Stock and Watson (1993) used Monte Carlo experiments to evaluate the finite sample properties of six alternative estimators under cointegration. Using samples of 100 and 300 observations, they found that while the Johansen estimates had the smallest bias, the variance was much larger than the other efficient estimators. In the same study they found that the DOLS estimator had the smallest root mean squared error among the estimation methods studied. Therefore, the DOLS was deemed appropriate for this study. For a more detailed explanation of this method, the reader is referred to Hayashi (2000).

Most existing econometric trade papers assume that estimated cointegrating vectors represent stable long-run relationships among modeled variables, so that estimated parameters are taken as constant over time. In this paper, we would like to test whether import expansion agreements caused detectable changes in the import demand relationship. As shown in Hayashi (2000) and supported by results in

Hansen (1992), our system (2) can be augmented to allow for structural breaks by including dummy variables:[8]

$$
\begin{aligned}
m_{k,t} = \alpha_k + \alpha_{k,T+} Dum_{k,T+,t} + \boldsymbol{\beta}_k \mathbf{x}_{k,t} + \boldsymbol{\beta}_{k,\mathrm{T}+} \mathbf{xDum}_{k,T+,t} \\
+ \sum_{i=0}^{p} \gamma_{k,-i} \Delta \mathbf{x}_{k,t+i} + \sum_{i=1}^{p} \gamma_{k,i} \Delta \mathbf{x}_{k,t-i} + v''_{k,t}
\end{aligned}
\quad . \qquad (3)
$$

Here, $\alpha_{k,T+}$ is the coefficient on a dummy variable having unit values beginning in period *T*, and $\boldsymbol{\beta}_{\mathrm{k,T+}}$ is a vector of slope coefficients on dummy variables $\mathbf{xDum}_{\mathrm{k,T+}}$ with non-zero values from period *T* onward.

Market-opening trade agreements may be expected to induce changes in estimated demand parameters to the extent that they alter the incentives or constraints on imports, given economic fundamentals. In a simple case, this may entail only a level shift in imports, but changes in the response of imports to income and relative prices—the familiar trade elasticities—cannot be ruled out. In particular, liberalization of non-price barriers to trade may raise the responsiveness of imports to both income and price developments.

If a structural break is found in the vicinity of a policy change, then it suggests that the trade policy may be responsible for the structural change. Of course such inference is far from clear-cut. It is possible that other (non-modeled) changes in policy, the economic environment, or firm or consumer behavior could also account for structural changes in the estimated equation.

There are several difficulties that arise in implementing this procedure. First, the possibility of structural change raises problems for determining cointegration rank. Rank tests under general structural change have not yet been derived (Hansen, 2003). Nevertheless, we have chosen to report standard rank test results for constant-parameter models as a guide to model specification. In addition, the model of structural change used here assumes once-and-for-all shifts in cointegration or adjustment vectors and does not allow for gradual change in model parameters following a regime shift.

There are a number of previous papers that look for evidence of structural change in Japanese import relationships, including Corker (1989), Moriguchi (1993), and Ceglowski (1996 and 1997). Hamori and Matsubashi (2001) look for structural change in a cointegration framework. Our research differs from these previous studies because it looks for evidence that structural change is linked to specific U.S.-Japan trade agreements, and our analysis is conducted at a disaggregated level appropriate to answering these industry-level questions.

5. Data and Results

Our study focuses on seven industries subject to U.S.-Japan agreements intended to raise Japanese imports. Total non-energy imports were included for comparison purposes. For each sector, a time series of real imports as well as industry-specific domestic and import price series were constructed. The overall price index used was the Japanese GDE deflator. The data sources are summarized in Table 2.

Quarterly import series were derived from monthly trade statistics of the Japan Tariff Association, reported in the monthly *Summary Report Trade of Japan*. We aggregated the data from monthly into quarterly time series. This removes the very short-term volatility in the underlying series and allows us to use quarterly GDE as the income variable. Real imports were calculated by deflating the nominal series by a matching import price index, described below. The import time series are graphed in Figure 1.

As discussed above, the selection of industries was based in part on their prominence in the U.S.-Japan trade debate and freedom from other obvious quantitative restrictions. However, selection was also constrained in part by the availability of data for the targeted products. For example, while there were closely watched agreements covering telecommunication equipment and cellular phones, it was not possible to obtain trade data of sufficient detail, or that covered the important service components. Obviously these decisions involve an element of judgment.

Import and domestic wholesale prices were constructed from Bank of Japan price series published in the *Price Indexes Annual* and available for recent years on the internet. In each case, the most closely matching disaggregated price series was used, although for one category (auto parts) only a more aggregate import price series was available. The Bank of Japan publishes historical series for only a handful of semi-aggregate categories. We have constructed monthly linked price series from the detailed historical price index reports, following a similar methodology to that used by the Bank. The yen-denominated monthly indices were aggregated to quarterly series by a simple average as is done by the Bank of Japan (BOJ) in their own calculation of quarterly statistics.

As is typically the case, the Japanese import price and import value series are in local currency non-inclusive of tariffs. Since economic decisions are based on relative prices *inclusive* of tariffs, it is important to make adjustments for changes in tariff rates over time. (See Stone, 1979.) We have made such adjustments to each import price and value series using an estimated time series of tariff rates for the product, based on tariff line data from the Customs Tariff Schedules of Japan. Tariff rates are given in Appendix Table A1. While for many categories the decline in tariffs is fairly small, for tobacco tariffs fall from 35% to 20% in a single fiscal year.[9]

Developments in real income are measured using real gross domestic expenditure (GDE) from Japanese national accounts data reported by the Economic and Social Research Institute (formerly Economic Planning Agency). It was necessary to splice real GDE series prior to 1980 in order to obtain a sample extending back into the 1970s. All of the series (except GDE and the GDE deflator which were already adjusted) have been seasonally adjusted using the Census X-12 multiplicative procedure, and log transformations were taken.

Augmented Dickey-Fuller and Phillips-Perron tests for non-stationarity were performed on each data series. Results are available from the authors upon request. In all cases, we are unable to reject the null hypothesis of a unit root in real imports, whether or not a deterministic trend was also included. Unit root tests for the two relative price terms were mixed, with rejections of the unit root hypothesis for one or the other test in some cases.

Cointegration and Structural Change

Table 3 reports the results of trace tests for the rank of the cointegrating space for each of the models.[10] For three of the seven industries (Medicine, Semiconductors, Tobacco) and aggregate non-energy imports, we can reject the null of no cointegrating vector at the 95% confidence level, and the rank does not appear to be greater than one. For Lumber, we cannot reject the null of no rank. Paper is more problematic in that more than one cointegrating vector is supported—in fact the system may have full rank. Passenger cars and auto parts, consistent with poor results elsewhere, show strong evidence rejecting rank=1 and also rank=2, suggesting either strong endogeneity or perhaps poor specification altogether.

The trace test results presented here should be viewed as merely suggestive of possible cointegrating relationships among the variables (or the lack thereof). As noted above, formal tests for rank with structural change have not yet been developed, and tests for cointegration in the presence of structural breaks may lead to incorrect inference (Hansen, 2003). In addition, cointegration tests in small samples may be biased, and there appears to be no consensus in the literature on the how best to address this bias (Maddala and Kim, pp. 214-220). In the following we proceed under the assumption that a single cointegrating vector exists in each case, and that this represents a demand function for imports.[11]

The results of estimation by DOLS with break dummies are reported in Table 4. Note that under this specification the coefficients represent long-run elasticities of real imports with respect to income (real GDP) and the two relative price terms, and the dummy variables test for changes in these elasticities and in the intercept following the break date.

Looking first at the full-sample parameters, we note that elasticity estimates generally conform to theoretical predictions, although significance levels are low. Income elasticities are large, ranging from 1.0 to 4.5, and all but one of the import relative price elasticities are negative but generally insignificantly different from zero. All but one of the "other" price elasticities are positive, as predicted by theory. For aggregate non-energy imports, the income elasticity of 2.8 is large compared to existing studies, while the relative price elasticity of -1.1 is typical, although these parameters are not significantly different from zero.[12] The low significance levels reflect in part the substantial data demands of DOLS; the DOLS-adjusted standard errors range from 1.1 to 3.3 times larger than standard OLS errors.

We were unable to fit satisfactory models for the two auto-related sectors, Auto imports and Auto Parts imports. We tend to find perverse price effects and poor fit for these two sectors. This failure to find satisfactory auto models using traditional specifications appears to be a rather robust result. (See Gangnes and Parsons, 2002.)

Turning to the intervention dummies, we find evidence of structural breaks at the time of U.S.-Japan trade agreements in some industries, while others exhibit apparent parameter constancy or perverse changes. Tobacco shows the clearest evidence of positive change following the 1986 agreement, perhaps due to its elimination of preferential excise tax treatment for Japan Tobacco and the establishment of a more U.S.-friendly distribution system. This agreement was also rated quite highly (9 out of 10) in the ACCJ survey. The income elasticity increases, and the responsiveness to relative import prices also rises. Paper products also show some evidence of positive impact, although except for the "other" price term the dummies are not significant at the 10% level using DOLS-adjusted standard errors. For the heavily-studied semiconductor industry, we find a large increase in price sensitivity following the 1986 Semiconductor Accord, but no change in the income elasticity.[13]

For Medicine, there is simply no statistically-significant evidence that the 1986 agreement to reduce regulatory red-tape for foreign producers raised import levels over what they would otherwise have been. (This is true even if we apply unadjusted OLS standard errors.) In the case of the lumber agreement, there was an apparent **reduction** in the import demand elasticity following the 1990 accord, which is nearly significant at the 90% level. Finally, partly as a robustness check, we look for shifts in aggregate non-energy imports at the time of the Plaza Accord. No significant break is detected. This is noteworthy, since several of our policies occurred in close proximity to this change in exchange rate management.

In sum, these results provide a mixed picture of the effectiveness of U.S.-Japan trade policies in raising the level or sensitivity of import demand to income growth and changes in relative import prices. Although we look at only a limited number of agreements, it is tempting to observe that agreements making quantifiable changes appear to have more robust effects. That is probably too strong a conclusion to make given the low significance levels of full-sample parameter estimates.

6. Conclusions

We have evaluated the effectiveness of several U.S.-Japan trade agreements by looking for structural breaks in import behavior near the time of the agreements. The method we have employed tests for a change in behavioral parameters within the context of an empirical import demand model. In this way, we are able to look for evidence of policy impacts while controlling for the “normal” influence of income and relative prices on import volumes.

The empirical results we have assembled here paint a mixed picture of the apparent effectiveness of trade agreements. For Tobacco, there is a sharp increase in both income and price elasticities following the 1986 Tobacco Trade Agreement. There is also some evidence of change in Paper demand behavior. For the much-touted Semiconductor Trade Agreement, sensitivity to relative prices appears to have increased following the 1986 accord. For these three sectors, there is some evidence that the agreements may have been a “success,” at least on purely mercantilist grounds. For Medical products, there was no evidence of policy impacts, and lumber may have seen a perverse effect. We were unable to study Autos and Parts because no satisfactory econometric import model could be established.

Reliable detection of an agreement’s impact on trade flows depends on satisfactory modeling of the underlying import behavior. Without a model of how imports depend on economic fundamentals it is simply not possible to tell whether an acceleration of imports is due to a policy change or evolution of the fundamentals themselves. In our view, the results of trade agreement studies that ignore the influence of fundamentals cannot be relied upon. The difficulty we encountered fitting some import equations, and low significance levels obtained with corrected standard errors, illustrate the challenge in applying such tools in circumstances where data samples are limited.

We acknowledge that statistical break analysis of this kind will always be subject to a difficulty of attribution. While a trade agreement or other policy shift could cause a change in model parameters, so too could changes in consumer behavior or in the structure of the industry or economic environment. We try to isolate the effect of trade agreements by testing for change only in the vicinity of the agreement, but certainly that does not eliminate the possibility that changes in import propensities or the rise of an important popular product (say, the Pentium chip) could explain the import change.

Agreements to raise U.S. imports into the Japanese market became a central part of U.S.-Japan trade policy in the 1980s and 1990s. They are likely to play an important role in other bilateral relationships in coming years. Recently, the U.S. auto industry has begun to agitate for actions to open the Korean automobile market to American cars and parts, and the growing Chinese current account surplus promises to create pressure for market opening there as well. Before additional political capital is spent negotiating such agreements, one would like better evidence on their effectiveness. Our analysis argues for a model-based approach to gathering such evidence, but also highlights the challenges involved in doing so.

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Figure 1. Real Imports
(Vertical line marks date of U.S.-Japan Trade Agreement)

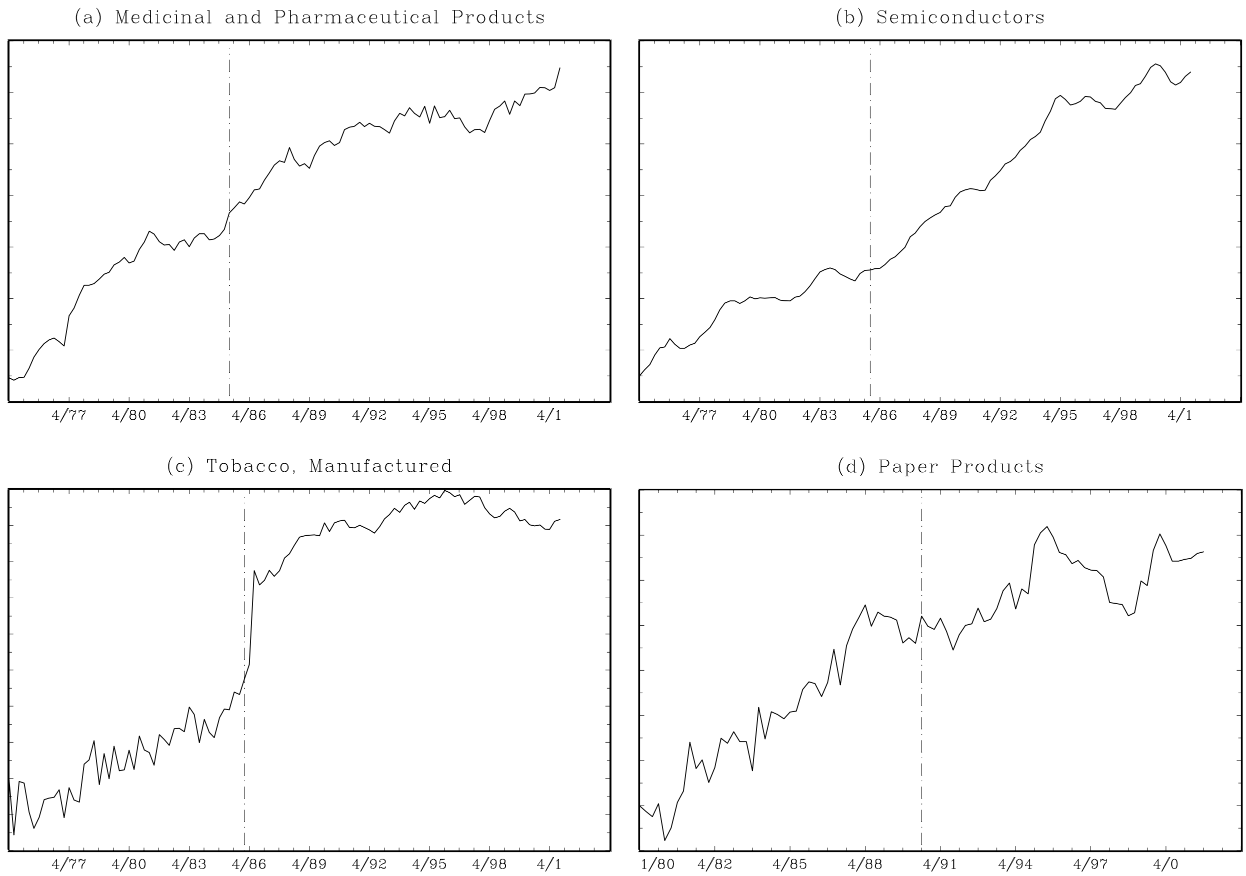

Figure 1. Real Imports (continued)

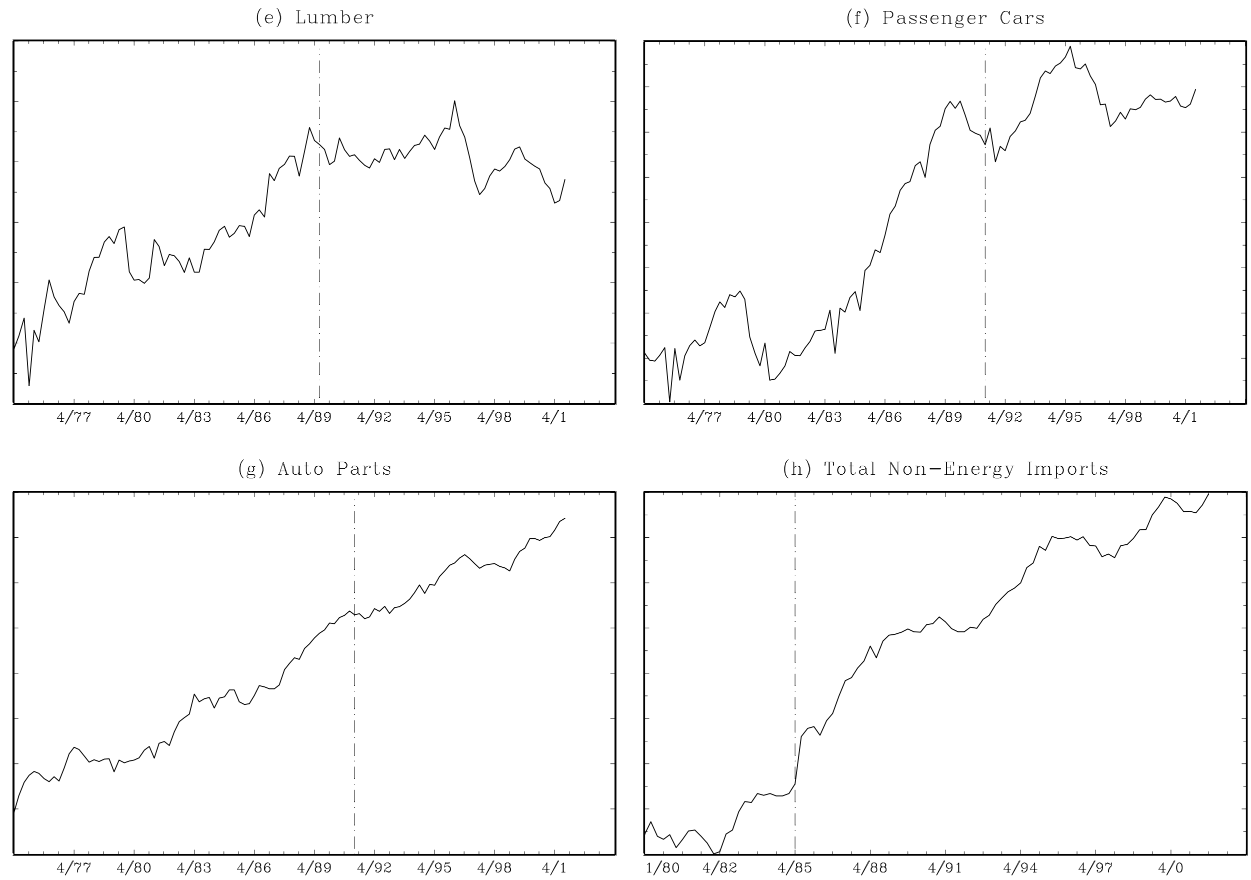

Table 1. A Timeline of Selected U.S.-Japan Trade Agreements, 1986-1992

Date	Policy	Summary
January 9, 1986	Report on Medical and Pharmaceutical MOSS Discussions	Reduced regulatory red-tape for foreign medical/pharmaceutical products and devices.
Announced July 31; signed September 2, 1986	Semiconductor Trade Agreement	Targeted increase of market share in Japan from 8% to 20%. Also implemented FMVs for Japanese exports.
October 6, 1986	Tobacco Trade Agreement	Eliminated import duties on foreign tobacco; loosened restrictive distribution system.
June 15, 1990	Wood Products Agreement	Additional tariff reductions; building standards changes to permit wood use
January 9, 1992	Auto and Auto Parts Plan	To double imports of U.S. autos parts by 1994; increase purchase of auto parts by Japanese affiliates.
April 5, 1992	Measures to Increase Market Access for Paper Products	To promote private sector purchases and overall market promotion of foreign paper.

Table 2. The Data

Category	Series No./Description	Import Price	Dom. WS Price	Sample
Medicine	No. 507 Medicinal and Pharmaceutical Products	Medicines	Medicines	1975:1-2002:3
Semi-conductors	No. 70311 Thermionic…semicond. dev, I.C.s, etc.	Integrated Circuits	Integrated Circuits	1975:1-2002:3
Tobacco	No. 10303 Tobacco, manufactured	Tobacco/ Cigarettes	Tobacco/Cigarettes	1982:1-2002:3
Paper	No. 60701 Paper and Paperboard	Paper	Paper and Paperboard	1980:1-2002:3
Lumber	No. 2070703 Lumber	Lumber	Lumber	1975:1-2002:3
Passenger Cars	No. 7050101 Passenger motor cars	Passenger Cars	Passenger Cars	1975:1-2002:3
Auto Parts	No. 70503 Parts of road motor vehicles	Passenger Cars	Automobile Parts	1975:1-2002:3
Total Non-energy Imports	Total imports less mineral fuels	Non-oil Weighted Ave IPI	Non-oil Weighted Ave Dom WPI	1980:1-2002:3

Sources: quarterly trade values are simple aggregates of monthly data from the Japan Tariff Bureau, *Summary Report on Trade of Japan*, various issues; import and domestic prices are from Bank of Japan, *Price Indexes Annual*, various issues (data since 1995 from BoJ web site).

Table 3. Cointegration Trace Tests

	Rank = 0	Rank >= 1	Rank >= 2	Rank >=3
Medicine	48.8* (47.2)	28.2 (29.7)	14.1 (15.4)	3.2 (3.8)
Semiconductors	52.3* (47.2)	24.4 (29.7)	8.9 (15.4)	3.3 (3.8)
Tobacco	50.5* (47.2)	20.2 (29.7)	8.9 (15.4)	0.01 (3.8)
Paper	69.9* (47.2)	35.3* (29.7)	18.1* (15.4)	4.1* (3.8)
Lumber	37.4 (47.2)	25.0 (29.7)	14.2 (15.4)	4.5* (3.8)
Passenger Cars	69.4* (47.2)	36.8* (29.7)	13.5 (15.4)	5.2* (3.8)
Auto Parts	60.2* (47.2)	30.6* (29.7)	12.0 (15.4)	1.9 (3.8)
Total Non-Energy Imports	54.9* (47.2)	28.8 (29.7)	12.7 (15.4)	2.58 (3.8)

Note: Note that all tests were done on the full sample, including Tobacco. Figures in parentheses are 5% asymptotic critical values from Johansen (1995). * indicates significant at or above the 95% confidence level.

Table 4. Structural Break Tests

Category	Break Period	y	pm	pd	Dummy Variables: level-dummy	y-dummy	pm-dummy	pd-dummy
Medicine	1986 Q1	3.475 *	-0.508	0.789	5.034	-0.413	-0.314	1.046
		(1.92)	(-1.61)	(1.05)	(0.20)	(-0.22)	(-0.74)	(0.78)
Semiconductors	1986 Q3	4.545 **	-0.320	0.263	22.447	-1.624	-0.982 ***	-0.949
		(2.07)	(-1.15)	(0.37)	(0.70)	(-0.66)	(-3.05)	(-1.26)
Tobacco	1986 Q4	2.318 ***	0.605	4.291 ***	-18.696 *	1.581 **	-2.092 **	-2.758
		(3.30)	(0.69)	(2.67)	(-1.83)	(2.05)	(-2.38)	(-1.54)
Paper	1992 Q2	1.015	-0.083	-1.187	-20.856	1.592	-0.566	3.172 **
		(1.63)	(-0.13)	(-0.86)	(-1.56)	(1.56)	(-0.58)	(2.19)
Lumber	1990 Q2	2.147 ***	-0.464 *	1.335 ***	15.103	-1.165	0.316	0.791
		(14.66)	(-1.83)	(4.16)	(1.62)	(-1.64)	(0.84)	(1.05)
Total Non-Energy	1985 Q3	2.780	-1.120	2.956	5.868	-0.467	1.176	-3.986
		(0.42)	(-0.80)	(0.33)	(0.07)	(-0.07)	(0.80)	(-0.45)

Note: A constant and one lead and lag of DOLS difference terms were included in all cases; these coefficients are not reported. DOLS adjusted t-statistics are in parentheses. A single asterisk indicates significance at the 10% level; two asterisks 5%; three asterisks 1%. Equations for Autos and Auto Parts are not reported, as no satisfactory model could be established.

Appendix Table A1. Average Tariff (Percent) for Individual Import Categories

Fiscal Year	Medicine	Semiconductors	Tobacco	Paper	Lumber	Passenger Cars	Auto Parts	Non-Energy Total
1975	7.6	9.5	355	6.8	3.5	6.4	7.0	n/a
1976	7.6	9.5	355	6.8	3.5	6.4	7.0	n/a
1977	7.6	9.5	355	6.8	3.5	6.4	7.0	n/a
1978	6.6	9.5	355	6.5	3.5	0	7.0	n/a
1979	6.3	9.5	355	6.5	3.75	0	7.0	n/a
1980	7	7.1	90	6.5	3.75	0	6.6	2.1
1981	6.9	8.6	35	6.7	3.75	0	6.6	2.1
1982	6.4	8.9	35	5.9	3.39	0	6.0	2.2
1983	6.3	5.6	20	5.6	3.39	0	6.0	2.1
1984	6	4.2	19.7	5.2	2.72	0	5.8	2.1
1985	5.4	0	20.4	4.9	2.39	0	5.3	2.2
1986	4	0	23.9	3.8	2.28	0	0	2.7
1987	4	0	0	3.5	2.03	0	0	2.8
1988	3.2	0	0	3.7	2.11	0	0	3.0
1989	3.2	0	0	3.3	2.11	0	0	2.6
1990	2.3	0	0	3.1	2.0	0	0	2.4
1991	2.2	0	0	3.1	2.0	0	0	3.0
1992	2	0	0	3.1	2.0	0	0	3.1
1993	2	0	0	3.1	2.0	0	0	3.3
1994	2	0	0	3.1	2.0	0	0	3.1
1995	0	0	0	3	2.0	0	0	2.9
1996	0	0	0	2.8	2.0	0	0	2.6
1997	0	0	0	2.3	2.0	0	0	2.4
1998	0	0	0	2.1	2.0	0	0	2.5
1999	0	0	0	1.8	2.0	0	0	2.2
2000	0	0	0	1.4	2.0	0	0	2.4
2001	0	0	0	1.1	2.0	0	0	2.4
2002	0	0	0	0.7	2.0	0	0	2.4

Note: these rates are for each Japanese fiscal year, which begins on April 1st and ends on March 31st. Monthly price and trade data were adjusted accordingly. Data used for the first quarter of 1980 may have had differing rates, not reported here.

[*] This paper was completed while Byron Gangnes was Visiting Associate Professor at Singapore Management University, Term II 2004. We would like to thank Theresa Greaney and Sumner La Croix for very helpful comments. Related results were presented at the Western Economic Association International conference, San Francisco, July 2001.

[1] Japan has a long history of friction with trade partners. Over the years, foreign governments have responded with a range of policies to protect specific industries including orderly marketing agreements, the use and threatened use of retaliatory duties, voluntary export restraints (VERs), as well as pressure for tariff and quota concessions under the GATT/WTO. The shift in emphasis from domestic protection to opening Japanese markets in part reflects increased constraints—after a generation of multilateral trade liberalization—to the use of many of the traditional trade protection tools. For example, VERs, the mainstay of U.S. policy in the 1980s, became effectively illegal under the Uruguay Round GATT accord.

[2] Japan has also acted (ostensibly) unilaterally to raise imports, notably in an early-1990s policy package that included unilateral tariff reductions, changes in licensing procedures, infrastructure investments geared toward trade, and explicit tax incentives for raising imports. See MITI-JETRO, c1993.

[3] In some sectors (Medicine, Semiconductors, Wood and Autos) there was more than U.S.-Japan agreement during this time period. Because of the limited sample, we have chosen in these cases to focus on what appears to be the more significant agreement. For Medicine, ACCJ (1997, 2000) survey respondents rated the first agreement as much more effective than the second. For Wood, the first agreement was seen as largely ineffective and so the second agreement was selected for this study. Similar criteria were used for Semiconductors and Autos and Auto Parts (the latter of which were later dropped).

[4] Import demand functions of this form are consistent from micro-based consumer optimization theory where it is assumed that imported and domestic goods are imperfect substitutes. Industry imports are a function of nominal income, import and domestic prices in the industry and all other prices,

$$m_k = f\left(Y, P_k^m, P_k^d, P_{\neq k}\right).$$

Under homogeneity, assuming a log-linear form, and taking an aggregate deflator to capture all "other" prices, we arrive at equation (1). A more detailed discussion of this model can be found in Mutti (1977). Under the assumption of no cross-price effects, the model could be further simplified to one involving a single relative import price term. For more discussion of the micro foundations of import demand and a summary of empirical studies through the early 1980s, see Goldstein and Kahn (1985). The log-linear form is commonly used, both for its convenience in calculating elasticities and because the log transformation yields smoother data for estimation purposes, but it is not without detractors. See, for example, Marquez (1999).

[5] In general, a deterministic time trend may also be included, but trend terms typically did not alter results significantly and were not retained in the models reported here.

[6] Cointegration takes note of the fact that two or more series that are I(1) processes, may, when regressed upon each other, result in some linear combination which is an I(0) process. Although cointegration is purely a statistical concept, it lends itself to the economic interpretation that there exists a long-run relationship between two or more variables. While the variables may drift apart in the short term, over the long run they will tend to move together. For a thorough introduction to the history of unit roots and cointegration as well as an excellent presentation of the recent advances, methodologies and future direction of cointegration, see Maddala and Kim (1999).

[7] Both single-equation cointegration techniques as well as system approaches to cointegration have been applied to trade modeling in recent years. An early example for aggregate demand is Asseery and Peel (1991) who implement an error-correction mechanism (ECM) approach. Rose and Yellen (1989) test for cointegration among the trade balance, income and relative prices. More recently, Caporale and Chui (1999) measure income and price elasticities for overall exports and imports for 21 countries including

Japan using ARDL and DOLS time series techniques. Recent disaggregated import demand estimates have used an error correction method. See, for example, Pattichis (1999) and Gallaway *et al.* (2003).

[8] A number of alternative tests for structural change under cointegration have been proposed. Quintos and Phillips (1993) develop tests for parameter constancy in cointegrating relations in a single-equation setting. Residual-based, single-equation methods have been developed by Gregory and Hansen (1996). However, like residual-based tests generally, they have low power because they tend to neglect model dynamics. (See Maddala and Kim, p. 203.) Full information maximum likelihood methods based on the multivariate Johansen (1995) procedure, such as Hansen (2003) may be superior to single equation methods for addressing problems of simultaneity, although performance is typically poor in small samples. Unfortunately, there appear to be no comprehensive Monte Carlo studies on this recent literature.

[9] The tobacco sample used for estimation starts in 1982 to avoid the massive reduction in tariffs that occurred between 1979 and 1981. Where tariff line detail could be matched closely with corresponding trade value data (and the number of tariff lines was not prohibitively large), weighted average tariff rates were constructed for the import category in question. This was true for tobacco, semiconductors, and autos. For paper, medicines, and auto parts, appropriate weights were not available, and a simple average of tariff lines was used. In each case, average tariffs rates for the commodity or category were calculated for each of the twenty-plus years, and these rates were used to 'mark-up' the corresponding series by that percentage.

For overall non-energy imports we took a different tack. Rather than attempting a simple or weighted average over thousands of categories, we followed Clemens and Williamson (2001) in deriving an implicit overall tariff rate by dividing the total tariff revenue by total import value. Interestingly, despite the conclusion of the Uruguay Round and other unilateral tariff reductions conducted by Japan over the past 20-plus years, the average tariff rate has declined very little.

[10] In all cases, four lagged difference terms were included in the vector error correction models with the exception of non-energy imports which used five. This is consistent with Caporale and Chui (1999) who find that for most countries' aggregate imports cointegration can be found with four lags, but for Japan (and three other countries) five lags are necessary.

[11] By using a single equation structure, we implicitly assume there is no important endogeneity that must be considered. That is very likely the case for income, since industry imports are negligibly small. It is somewhat harder to argue for the two industry relative price terms. In fact, because we have not formally identified structural import demand equations from the VAR system, the cointegrating relationship could represent a linear combination of both demand and supply effects. This difficulty is common in the applied trade literature.

[12] Hooper and Marquez (1995) report a mean import price elasticity for Japan of –0.97 from 13 studies dating to 1987. The range of estimates is large, however, with price elasticity estimates from –3.4 to –0.26. More recent estimates of Japanese income/price elasticities include Asseery and Peel (1991), 1.36/-0.64; Ceglowski (1996), 0.73/-0.67; and Caporale and Chui (1999), 1.33/-0.33. On a disaggregated level, typically price elasticities (not only for Japan) are found to be between (negative) 1 and 3, as in the early study by Stone (1979). However, the variation can often be large and often obtaining reliable long-run parameters is difficult as shown in Gallaway et al. (2003).

[13] This result differs somewhat from Parsons (2002) who found no structural break in either the long run cointegration parameters, nor in the adjustment coefficients. This may highlight the sensitivity of results to differing econometric methodologies, particularly in small samples.