USING STRUCTURAL BREAK TESTS TO EVALUATE POLICY CHANGE: THE IMPACT OF U.S.-JAPAN TRADE AGREEMENTS

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Empirical evaluations of trade agreements often rely on descriptive statistics or univariate time series methods to detect subsequent changes in trade flows. We conduct a more satisfactory test by evaluating an agreement in the context of a structural econometric model. Consistent with trade theory, import demand is modeled as a cointegrating relationship with income and relative price variables, where trade agreements may cause structural changes in cointegrating vectors. This approach is applied to study the effect of several U.S.-Japan market-opening trade agreements; in three of seven industries we find evidence of structural change that may be related to trade agreements.

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

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I. 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 recent measures 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 has entered into a number of bilateral trade agreements designed to increase import access. Among the most visible of the market opening policies have been the so-called Voluntary Import Expansion policies, Japanese commitments (or targets, depending on whom you ask) to raise foreign market share in key Japanese sectors. These include the U.S.-Japan semiconductor accord of 1986 and the auto parts agreements reached in 1992 and 1995. But, they also include licensing concessions, agreements on standards, and changes in government procurement designed to raise foreign import access.[2] In this paper we evaluate the impact of several U.S.-Japan trade agreements on targeted Japanese industries within a structural econometric model of trade.

While the wide range of import expansion measures looks impressive on paper, it is less clear how big an impact they have had on trade patterns. 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 a subset of targeted industries over the 1980-2000 period.

As we discuss below, there are a number of 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 limited econometric studies. In this paper, we look for evidence that trade agreements altered import patterns using a system-based time series methodology. Our approach permits us to test for changes in import behavior within the context of a structural econometric trade model. Our methodology, based on Hansen (2001), estimates a cointegrating relationship describing the long-run evolution of imports in each sector, and tests for a structural break in that relationship. An indication of a structural break in close proximity to an import expansion agreement is interpreted as evidence that the agreement significantly altered import behavior relative to the path predicted by the evolution of economic fundamentals.

The paper is organized as follows. In section 2, we outline 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 the standard empirical trade model. In section 4, we describe our methodology for evaluating structural change within a behavioral trade model. Section 5 presents results, and concluding observations are given in section 6.

II. 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 recent years. We will examine some of the more notable and well-defined 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. The following is a brief description of each agreement.

Medical and Pharmaceutical Products

As an outgrowth of the Market-Oriented, Sector-Specific (MOSS) talks, this 1986 agreement introduced more transparency and reduced regulations in the approval process for foreign drugs as well as eased approval and licensing for medical products. In most cases, data from foreign clinical trials could now be used, eliminating the need for costly and redundant domestic trials. A second agreement, in 1994, established more transparent procedures for Japanese government procurement of medical products and services. ACCJ members initially gave the 1986 agreements high marks for improving market access, although the year 2000 report raised concerns about remaining regulatory hurdles. Public procurement was viewed as much improved following the 1994 accord.

Semiconductors

The Semiconductor Trade Agreement (STA) consisted of two five-year agreements (1986, 1991) which had two goals: (1) to institute Fair Market Values (FMVs) for Japanese exports, essentially creating a price floor; and (2) to increase foreign market share of semiconductors from around 8 percent 20 percent. (A 1996 follow-on agreement is not considered in our analysis.) Some observers (Tyson, 1992, for example) have hailed the agreement as a very successful strategic trade policy. The ACCJ (1997) survey found the STA contributed to increased access for U.S. chips makers. The theoretical literature (Greaney, 1996; Nakamura, 1995; Parsons, 2000) generally finds such a policy distortionary and welfare reducing, however.

Tobacco

This 1986 agreement eliminated high tariffs on tobacco and additional retail taxes that were imposed on foreign tobacco products. In a separate understanding, foreign cigarettes were given equal treatment to domestic brands in the distribution channels managed by Japan Tobacco Inc., the government tobacco monopoly. ACCJ representatives viewed these agreements as very successful, and import market share increased substantially in succeeding years.

Paper and Paper Products

The 1992 measures to increase market access for paper products neither reduced formal trade barriers nor set numerical targets as in the case of semiconductors. Nonetheless, efforts were made by Japanese authorities to promote foreign paper products in Japan in part by providing market information and in some cases offering low-interest loans to foreign suppliers. U.S. authorities also attempted to promote U.S. exports to Japan by encouraging U.S. firms to respond to requirements of Japanese producers such as quality, metric sizing and delivery. Perhaps unsurprisingly, this policy was ranked as very ineffective according to the ACCJ, and little nominal growth in paper product imports occurred.

Wood Products

Also, as part of the 1986 MOSS talks, Japan agreed to tariff reductions on plywood and veneers, some wood products and lumber. They also began to make concessions on building codes and standards to permit increased use of foreign wood products in construction. More progress on building standards was made after the 1990 agreement, influenced in part by a U.S. Super 301 Trade Action that identified wood product imports as facing unreasonable restrictions. Tariffs were further reduced, and subsidies were also pared. Industry observers gave the 1990 agreement high marks but reported only limited progress in raising import volume, perhaps in part because of the weak conditions prevailing during the 1990s in the economy generally and in the construction industry specifically. (A 1997 agreement on graded lumber was not included in our analysis.)

Autos and Auto Parts

Three agreements (1987, 1992, and 1995) cover auto imports and procurement of American auto parts by Japanese companies. The most significant agreement, in 1992, laid out numerical targets believed by both sides to be attainable. In particular, there was the expectation that Japanese imports of U.S. auto parts would double to $4 billion by 1994, and that overall procurement of American parts by Japanese companies would rise by $10 billion during the same time frame. The 1992 agreement also included commitments by Japanese auto manufacturers to "design in" more American parts and U.S. and Japanese efforts to promote U.S. auto exports. The numerical goals of the agreement were substantially met. Though a relatively small measure quantitatively, the 1992 auto accord represents (with the STA) one of the few examples of results-oriented trade policies (ROTPs) between the U.S. and Japan.

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. Generally the agreements governing these sectors are not amenable to analysis using the structural econometric techniques of this paper either because data at a suitable level of disaggregation is not available, or because the industry was subject to additional quantitative restrictions that may obscure detection of policy effects. See ACCJ (1997, 2000) for a complete description of U.S.-Japan trade agreements.

There are a number of existing empirical studies of the impact of market-opening Japanese trade agreements. These studies include business surveys and descriptive analyses of trade patterns, as well as more formal econometric tests of structural change either in a single equation or multivariate setting.

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 US 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; Nakamura, 1995; 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.

We are aware of very few econometric studies of U.S.-Japan trade agreements. Greaney (2001) constructs Chow tests for structural breaks in import time series and surprisingly finds more evidence that U.S.-Japan agreements raised Japan's imports from third countries than from the U.S. 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 products exports to Japan.

A problem with non-structural analyses of trade agreements is that it is difficult to interpret a detected break in the time series. As Greaney (2001) acknowledges, finding a structural 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 structural model of trade is needed to capture these effects.

III. Structural Import Models

Typically import demand functions are derived from micro-based consumer optimization theory where it is assumed that imported and domestic goods are imperfect substitutes.[3] Two factors drive the demand for imports: income and relative prices. Gross Domestic Product (Expenditure) is often used as a proxy for income or domestic demand and a ratio of import prices to some domestic price index (usually the wholesale price index) is used as a measure of relative prices. A log-linear form is used both for its convenience in calculating elasticities and because the log transformation yields smoother data for estimation purposes.[4]

A typical import demand equation takes the form:

$$\ln M_t = a_0 + \sum_{i=1}^{k_1} \beta_{1i} \ln M_{t-i} + \sum_{i=0}^{k_2} \beta_{2i} \ln Y_{t-i} + \sum_{i=0}^{k_3} \beta_{3i} \ln \left(\frac{P_{mt}}{p_{dt}} \right)_{t-i} + \beta_4 T_t + u_t \tag{1}$$

where M is real imports (deflated by the import price index), Y is real income and P_m and P_d are import and domestic price indexes, T is a time trend, and u_t reflects measurement error. Distributed lags of the dependent and independent variables are included to capture dynamic effects and correct for serial correlation of the residuals. A common form of (1) includes a single lagged dependent variable and a one-year lag of the relative price term. In the estimation below, the order of the lag polynomials is chosen empirically to be sufficiently long to yield approximately white noise errors.

As model (1) is typically estimated using time-series data, OLS estimation in levels may yield a spurious regression.[5] In some cases, regressions with differenced data have been used to eliminate non-stationary regressors, although this approach will often neglect important long-run relationships.

In the 1990s, a number of economists began to estimate traditional import demand models using the cointegration techniques that emerged in the wake of Engle and Granger (1987) and others. Both single-equation cointegration techniques as well as system approaches to cointegration were used. An early example is Asseery and Peel (1991) where they implement an error-correction mechanism (ECM) approach and compare their results with more traditional econometric estimates of import demand. Rose (1991) and Rose and Yellen (1989) test for cointegration among the trade balance, income and relative prices. Asseery and Peel find imports to be much less elastic with respect to income and prices than traditional long-run estimates. Rose and Rose and Yellen find that the effect of the exchange rate on the trade balance is insignificant. 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. They find positive income elasticities—typically greater than unity—and modest negative relative import price elasticities.

IV. Detecting Structural Breaks in an Import Model

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.

Our approach is to estimate a traditional import demand function, relating real import volumes to real gross domestic expenditure and the relative import price, while allowing for a structural break near the time of the trade agreement. If a structural break is found in the vicinity of a policy change, then we will interpret the result as suggestive that the trade policy is 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 a number of previous papers that look for evidence of structural change in Japanese import relationships, including Corker (1989), Ochi and Utsunomiya (1993), Moriguchi (1993), 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.

Cointegration and Structural Change

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 an economic interpretation as well, namely, that the cointegrating relationship itself may describe the long-run relationship between two or more variables, and that although the variables may drift apart, over the long run they will tend to move together. This also complements the equilibrium concept in economics where, although there may be short-run dynamic fluctuations, market forces should bring variables “back-in-line” with each other in the long run.[6]

Important changes in policy or the economic regime may lead to changes in the long-run relationship. We can therefore look for structural breaks in cointegrating relationships as evidence of the impact of a particular policy. 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. 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.) In addition, single equation estimation methods may be inferior to systems methods for addressing problems of simultaneity and for identifying structural relationships.

Vector autoregressive (VAR) approaches to testing for structural change implement Maximum Likelihood Estimation (MLE) to test for either a change in the rank or a change in parameters of the cointegrating vector(s). The literature in this area is very limited. Hansen and Johansen (1993) develop tests for a change in the rank of the system, that is, the number of independent cointegrating relationships. Their method involves splitting the sample into two sub-samples and then testing whether or not the number of cointegrating vectors suggested by trace and/or eigenvalue tests are different. Seo (1998) uses a Lagrange multiplier test for a single structural change in the cointegrating and/or adjustment vectors at an unknown time period, against the alternative of constant parameters.

In this paper, we apply a framework due to Hansen (2001), which permits likelihood ratio tests for structural change at (possibly multiple) hypothesized break points. Hansen's technique is particularly appropriate in this setting, because market-opening trade agreements, if effective, would be expected to change parameters of long-run demand relationships—either price and income elasticities or deterministic components—at or near the time of the agreement. Hansen's framework also allows for the possibility of non-constant cointegration rank and hypothesis testing under the assumption of structural change.[7]

Hansen considers a p-dimensional time series process X_t generated by the vector error correction model (ECM),

$$\Delta X_t = \alpha\beta' X^*_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \Phi D_t + \varepsilon_t, \qquad t = 1,...,T, \tag{2}$$

where $\varepsilon \sim iid\, N(0,\Omega)$, α is a $p \times r$ matrix of adjustment coefficients and β is a $p_1 \times r$ matrix of the system's r cointegrating vectors. The vector X^*_{t-1} consists of lagged levels of X and possibly other restricted deterministic variables. In our case, a deterministic trend is included. The q-dimensional vector D_t contains deterministic terms that are not restricted to the cointegrating space, such as constant, trend and seasonal dummies. The summation term contains dynamic elements of the ECM, with maximum lag length, k. Note that the r cointegrating relationships in (2) take the form

$$W_t = X_t + \beta' X_t, \tag{3}$$

which are stationary linear combinations of the X_t variables under the null of cointegration.

Hansen incorporates structural change by allowing parameters to change at given change points: $T_1, \ldots, T_{m-1}$, where $0 < T_1 < \ldots < T_{m-1} < T$. This generalizes (2) to:

$$\Delta X_t = \alpha(t)\beta(t)' X^*_{t-1} + \sum_{i=1}^{k-1} \Gamma_i(t)\Delta X_{t-i} + \Phi(t)D_t + \varepsilon_t, \qquad t = 1,\ldots,T, \tag{4}$$

where the time-varying parameters are piecewise constant and given by

$$\alpha(t)\beta(t)' = \alpha_1\beta_1{}' 1_{1t} + \ldots + \alpha_m\beta_m{}' 1_{mt,} \tag{5}$$

with a similar piecewise form for $\Gamma_1(t)$, $\Phi_1(t)$, and $\Omega(t)$ parameters. Here, Hansen has defined an indicator function to identify the *m* sub-samples,

$$1_{jt} \equiv 1(T_{j-1} + 1 \le t \le T_j), \qquad j = 1,\ldots,m, \tag{6}$$

where $T_0 \equiv 0$ and $T_m \equiv 0$.

Notice that in this specification, the cointegrating space for a sub-sample *j* is defined by the vector β_j, and two sub-samples β_i and β_j are not required to have the same number of columns. This permits in principle for changes in cointegrating rank across sub-samples, although in our application a constant cointegrating rank is assumed for the complete sample.

Hansen shows that the model defined by (2) through (6) can be rewritten as a regression equation with constant parameters. (See Hansen, pp. 4-5, for details.) Importantly, the resulting equation takes the form of a reduced rank regression, and so properties of the model can be derived from known results on such models (see Anderson, 1984, and Johansen, 1995). Unlike a standard reduced rank regression equation, the sub-period structure of Hansen's model yields a particular block-diagonal structure, which can be imposed using a set of linear restrictions. The model also has a non-constant covariance matrix. Hansen derives the maximum likelihood estimators for the parameters of this "generalized reduced rank regression."

Tests for structural change can be conducted by testing for parameter constancy across sub-periods using the linear restriction matrices described above. Hansen shows that for the appropriately restricted model, and assuming constant cointegration rank, the likelihood ratio test of a restricted model with q fewer parameters against the more general model is distributed as χ^2 with q degrees of freedom.[8] (See Theorem 14, p. 17.) While in principle stability tests can be conducted on all model parameters, in practice testing is restricted to a subset of parameters so that the number of free parameters is small relative to the sample size. In this paper, we restrict our attention to possible changes in the β vector that defines the long-run cointegrating relationship.

An outline of the estimation procedure is as follows. First, the system is estimated assuming constant parameter vectors. Trace tests are used to indicate the cointegration rank of the system.[9] Once a rank for the system has been selected, we estimate the unrestricted system that permits a structural break in the cointegrating vector β at the time of the trade agreement, and we estimate the system with the restriction that the parameter vector is constant across pre- and post-agreement time periods. We then calculate a chi-square test for the restriction in question and compare it to the appropriate critical value. For industries subject to more than one agreement, separate tests of parameter constancy are conducted at each agreement date.

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 been derived (Hansen, 2001, p. 21).

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. Because of the uncertainty about when the effects of an agreement might first be felt, we have tested for breaks within a window of several quarters following the agreement. This also sheds some light on the robustness of estimation results.

V. 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, time series of real imports and an industry-specific relative price were constructed. 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 on Trade of Japan*. We chose to work with quarterly, rather than monthly time series to remove some of the short-term volatility of the underlying series. Real imports were calculated by deflating the nominal series by the appropriate import price index, as outlined below.

As discussed above, the choice of categories to study was based in part on 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 to construct an appropriate series.

Obviously these decisions involve judgment. Other factors also influenced the selection of categories, as described above.

Import and domestic wholesale prices were constructed from Bank of Japan price series published in *Price Indexes Annual* and available for recent years on the web. 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 long time 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 estimated time series of tariff rates for that product, based on tariff line data from the Custom Tariff Schedules of Japan.[10] Tariff rates are given in Appendix Table A1. While for many categories the decline in tariffs is fairly small, for tobacco tariffs fall from 90% to 35% in a single fiscal year.

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). All of the series (except

GDE) have been seasonally adjusted using the Census X-11 multiplicative procedure, and log transformations were taken.[11]

The time series are shown in Figures 1.1 to 1.9. For each industry, we have graphed real imports against the corresponding relative price series. In some cases, it is clear that changes in relative prices, especially those associated with the yen appreciation after 1995, may help to explain movements in import volumes. (Note, for example, how the surge in medicine imports in Figure 1.1 corresponds with the fall in relative import prices.) The structural framework we adopt in this paper is intended to capture such fundamental determinants of trade to permit a precise evaluation of trade policy effects.

Unit root tests

Augmented Dickey-Fuller and Phillips-Perron tests for nonstationarity were performed on each data series. Results are reported in Appendix Table A2. The tests show strong evidence of unit roots in all of the series with or without trends at standard significance levels.

A number of the series also appear to have deterministic trends. Therefore, in the cointegration analysis below we allowed for a possible trend in the DGP.

Cointegration and Structural Change

Table 3 reports the results of trace tests for the rank of the cointegrating space for each of the models.[12] For four of the seven industries and for aggregate non-energy imports we can reject the null of no cointegrating vector at the 95% confidence level. In the case of passenger cars the test statistic is very near the 5% critical value for rejection. For the two remaining series (tobacco and paper) evidence of cointegration is weaker.

There are no cases where a rank of two is indicated, although test statistics for semiconductors and lumber are close to the 5% critical values.

There are difficulties interpreting cointegration test results in our setting. First, the existence of a structural break might lead to incorrect inference.[13] 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 analysis that follows, we assume a rank of one for all models, and we proceed to test for structural change in the single cointegrating vector associated with the largest eigenvalue of the system.

In Table 4, we report the result of tests of structural stability in the cointegrating vector near the date of the relevant U.S.-Japan trade agreement(s). In each case, we also report normalized cointegrating vectors under the assumption that parameters are constant over the entire sample, and under the alternative hypothesis of a structural break. If the rank of the system is one, it may be reasonable to interpret the single cointegrating vector as an import demand function, so we have normalized the vector accordingly.[14] In this case, the coefficients represent long-run elasticities of real imports with respect to income (real GDP) and relative price.

Looking first at coefficients of the constant parameter case, we note that elasticity estimates do not uniformly conform to theory. Income elasticities are generally positive and between 0.8 and 1.4, consistent with the range of estimates reported in the literature, with the notable exceptions of passenger cars and tobacco.[15] Relative price elasticities show more variation, with several displaying a perverse positive effect. The latter may reflect lack of price sensitivity of demand, the influence of supply in this underidentified system (see footnote 14), or biased estimation due to structural change. For aggregate non-energy imports, relative price also enters with a small perverse positive sign. The deterministic trend coefficients are all positive, consistent with persistent increase in import penetration and product variety over this period.[16]

Turning to the stability test results, note that for each industry there is at least one agreement in the vicinity of which we can reject parameter constancy. Because of uncertainty about when the effects of a new agreement might be felt, we report tests for a four-quarter window that includes the agreement quarter itself and three succeeding quarters. In the case of the 1990 Wood Products Agreement, a significant break is detected only after three quarters have elapsed.

While these break test results, taken as a whole, appear to support a significant impact of trade agreements on import relationships, inspection of pre- and post-break cointegrating relationships in some cases casts doubt on this interpretation. For passenger cars, pre-break parameter estimates frequently have perverse signs, and post-break elasticities are unbelievably large. Parameter signs for tobacco and the pre-break sample for paper are also perverse. For autos and tobacco, the estimation results at adjacent candidate break points demonstrate considerable instability. (This is true to a lesser extent for auto parts and lumber.) While one might take such sensitivity as evidence that only break points in certain periods are "correct," we are concerned that it may also signal a fundamental difficulty fitting a robust import model over the relatively short available samples.

In addition to the sensitivity of structural parameters to particular break points, the direction of change of parameters between pre- and post-break samples does not always appear consistent with a positive impact of the agreement on import volumes. In the case of the 1990 Wood Products Agreement, for example, both the income elasticity and deterministic trend fall after the trade agreement.

Perhaps not surprisingly, the industry where the clearest support is evident for a positive trade agreement effect is semiconductors.[17] At the time of the first semiconductor trade agreement in 1986, there is a significant break in the parameter vector. Estimation under the assumption of a break yields larger income and price elasticities for both pre-and post-break samples than emerges under the null of constant parameters and a post-break deterministic trend growth rate that is 2% higher than in the pre-agreement period. A similar result occurs at the time of the second STA, where the difference between pre- and post-agreement trends is 1.6%, although the structural break is not significant at the 95% level. Elasticity estimates are fairly stable within the four-quarter window following the agreements, except for the break test at 1992Q2.

Auto parts and medicine results also provide some support for a positive policy impact. Both income and relative price elasticities rise in the period following the 1992 Auto and Auto Parts Plan (although the deterministic trend parameter declines slightly). While there is also a significant structural break at the time of the 1987 auto agreement, parameter estimates for subsamples vary erratically. Both U.S.-Japan agreements affecting medical and pharmaceutical products appear to raised income and price sensitivity of these imports, although again there was a small accompanying decline in the deterministic trend.

For comparison purposes, we also analyze structural change in a quarterly index of aggregate non-energy imports. Testing for breaks at the time of the Plaza and Louvre accords, we find that the latter is significant at the 95% level, while the former is very nearly so. Admission of a break yields elasticity estimates that more closely conform to expectations, with income elasticity well above unity, and a significant negative price elasticity for both pre- and post-break samples. Post-break import price sensitivity is lower, perhaps reflecting changing behavior in the face of the extraordinary price changes that occurred during this period. This is consistent with the argument by Greaney (2000), Lincoln (1999), and others that large relative price changes were needed to break old buying patterns. The deterministic trend component of import growth increased by about 1% in the post-break period. An arbitrary mid-sample break test does not reject the null of constant parameters, perhaps dispelling concern that the break tests are prone to detecting spurious breaks.

VI. Conclusions

We have evaluated the effectiveness of 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 image of the effectiveness of trade agreements. We do find evidence that the Semiconductor Trade Agreement (especially the first one) caused a trend increase in semiconductor imports. This is consistent with the predominant view by both researchers and industry experts that the agreement was a success, at least on purely mercantilist grounds. Imports of auto parts appear to have become more income and price elastic after the 1992 agreement, and medical products following both 1986 and 1994 accords. In both cases, however, there were slight declines in the deterministic trend component. Models for paper and lumber, while exhibiting breaks, did not support net positive impacts from the trade agreements.

In other cases, including the effect of the highly touted 1992 auto accord on assembled passenger cars, while break tests reject parameter constancy, the structural trade models are not well behaved. Estimated parameters for pre- and post-break samples differ dramatically when the hypothesized break point is moved by one quarter. We fear that this may reflect difficulty reliably estimating the standard empirical trade model when the available data samples are fairly short.

This is an important issue. 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. Relatively short samples may be at the heart of the problem, or it may be that the standard trade model is not well suited to modeling disaggregated Japanese imports. Further research in this area is needed.

It is important to note that statistical structural break analysis of the kind applied here will always be subject to a difficulty of attribution. While a trade agreement or other policy shift could cause a structural 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 structural approach to gathering such evidence, but also highlights the challenges involved in doing so.

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Figure 1.1. Medicinal and Pharmaceutical Products

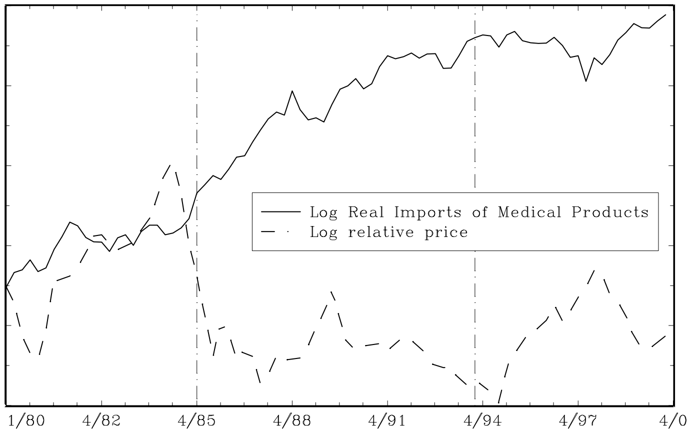

Figure 1.2. Semiconductors

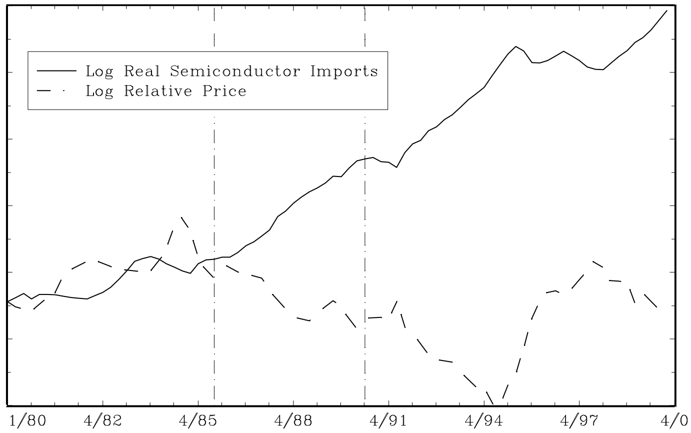

Figure 1.3. Tobacco, Manufactured

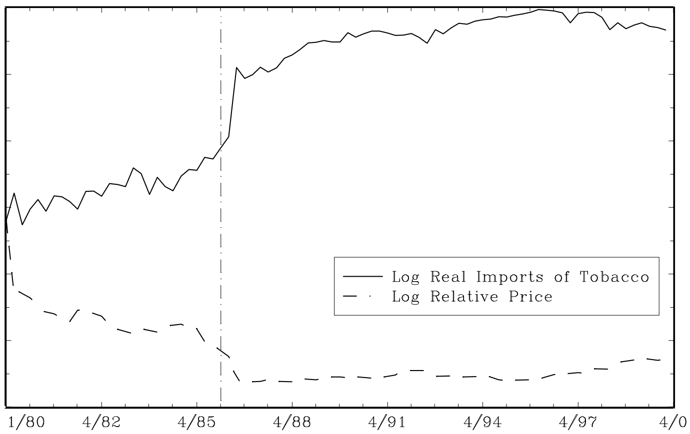

Figure 1.4 Paper Products

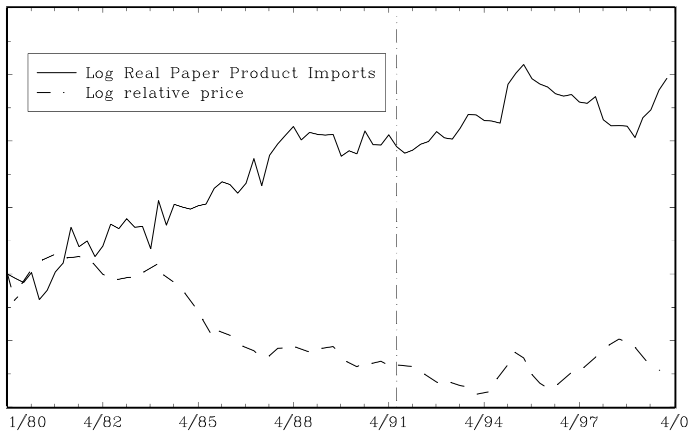

Figure 1.5. Lumber

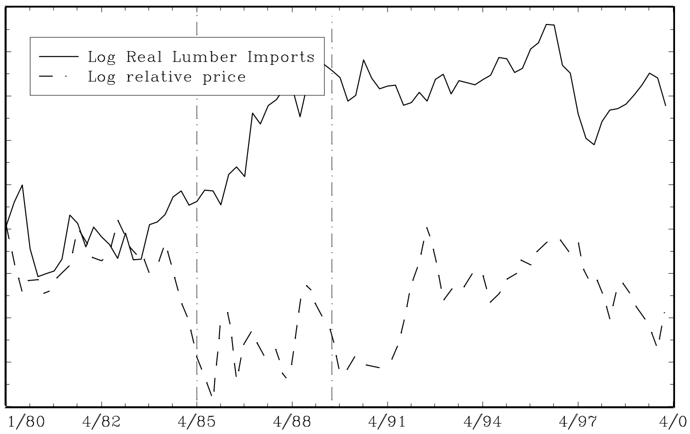

Figure 1.6. Passenger Cars

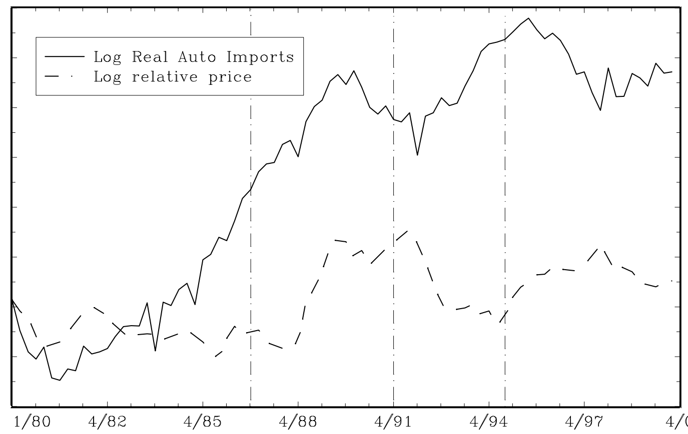

Figure 1.7. Autos Parts

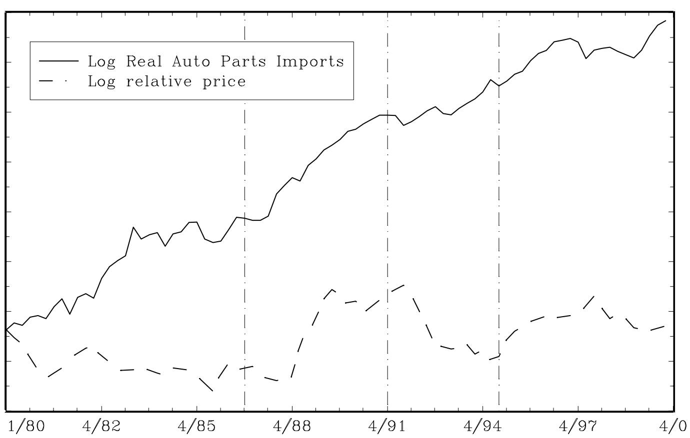

Figure 1.8. Total Non-Energy Imports

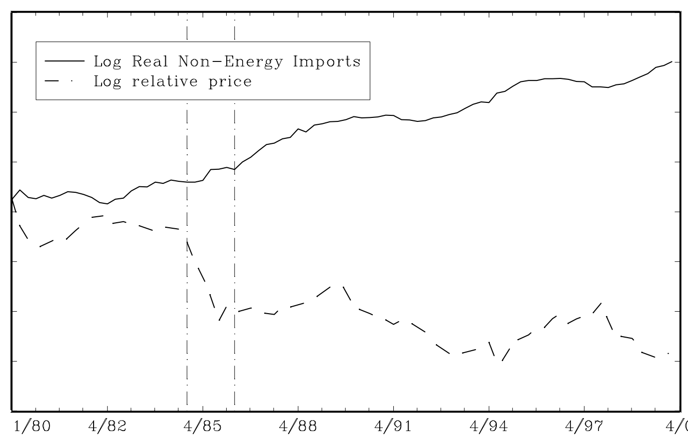

Figure 1.9 Gross Domestic Expenditure

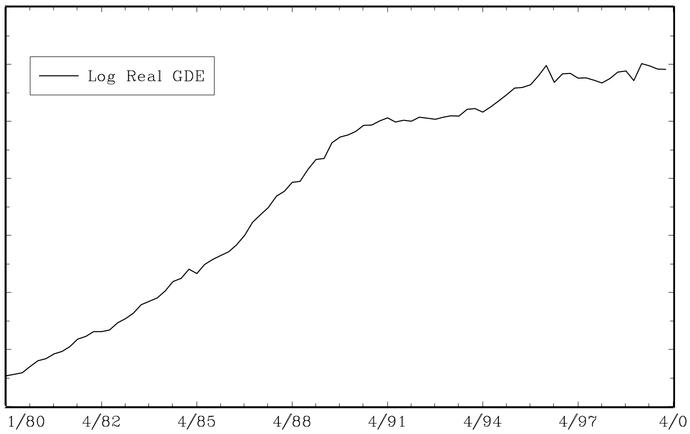

Table 1. A Timeline of Selected U.S.-Japan Trade Agreements

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.
January 10, 1986	Wood Products Agreement	Reduce tariffs on select lumber and wood products; consultations on building code
October 6, 1986	Tobacco Trade Agreement	Eliminated import duties on foreign tobacco; loosened restrictive distribution system.
August 18, 1987	MOSS Talks on Transportation Machinery	Talks on standards, testing, and cooperation; established acceptability of foreign parts for repairs.
June 15, 1990	Wood Products Agreement	Additional tariff reductions; building standards changes to permit wood use
June 11, 1991	Semiconductor Trade Agreement 2	Reconfirmed goal of "gradual and steady progress" toward 20% U.S. market share target; resolved dumping complaints.
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.
November 1, 1994	Measures Related to Japanese Public-Sector Procurement of Medical Technology	New measures and guidelines for public procurement of medical technology, products and services.
August 23, 1995	Japan Automotive Agreement	Increase dealerships in Japan; further increase in auto parts imports though no numerical targets were agreed upon

Source: American Chamber of Commerce in Japan (1997, 2000).

Table 2. The Data

Category	Series No./Description	Import Price	Dom. WS Price	Sample
Medicine	No. 507 Medicinal and Pharmaceutical Products	Medicines	Medicines	1980:1-2000:4
Semi-conductors	No. 70311 Thermionic…semicond. dev, I.C.s, etc.	Integrated Circuits	Integrated Circuits	1980:1-2000:4
Tobacco	No. 10303 Tobacco, manufactured	Tobacco/ Cigarettes	Tobacco/Cigarettes	1980:3-2000:4
Paper	No. 60701 Paper and Paperboard	Paper	Paper and Paperboard	1980:1-2000:4
Lumber	No. 2070703 Lumber	Lumber	Lumber	1980:1-2000:4
Passenger Cars	No. 7050101 Passenger motor cars	Passenger Cars	Passenger Cars	1980:1-2000:4
Auto Parts	No. 70503 Parts of road motor vehicles	Passenger Cars	Automobile Parts	1980:1-2000:4
Total Non-energy Imports	Total imports less mineral fuels	Non-oil Weighted Ave IPI	Non-oil Weighted Ave Dom WPI	1980:1-2000:4

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). The tobacco series was truncated to avoid a massive tariff change in the first quarter of 1980 (from 335% to 90%).

Table 3. Cointegration Trace Tests

	Rank = 0	Rank >= 1	Rank >= 2
Medicine	53.6* (42.2)	14.8 (25.5)	4.6 (12.4)
Semiconductors	50.7* (42.2)	23.3 (25.5)	8.5 (12.4)
Tobacco	34.4 (42.2)	11.7 (25.5)	3.8 (12.4)
Paper	36.5 (42.2)	13.2 (25.5)	3.4 (12.4)
Lumber	48.8* (42.2)	23.8 (25.5)	5.1 (12.4)
Passenger Cars	40.7 (42.2)	19.9 (25.5)	5.4 (12.4)
Auto Parts	42.4* (42.2)	21.1 (25.5)	6.5 (12.4)
Total Non-Energy Imports	49.5* (42.2)	18.3 (25.5)	3.4 (12.4)

Note: A VAR with four lags was used in each case. While trace test results are presented here, they are merely suggestive of possible cointegrating relationships among the variables. Formal tests for rank with structural changes are currently unavailable (Hansen, 2001). Figures in parentheses are 5% asymptotic critical values from Johansen (1995).

Table 4. Structural Break Tests

	Break Test				Normalized Cointegrating Vector								
					Constant Parameters			Pre-break sample Parameters			Post-break Sample Parameters		
Category	Break period		LR	P-val	Income	Rel Price	Trend	Income	Rel Price	Trend	Income	Rel Price	Trend
Tobacco	**1986 Q4**	**TTA**	20.768	0.000 *	-2.180	1.780	0.003	-1.321	1.361	-0.007	-2.316	1.787	0.010
	1987 Q1		22.553	0.000 *				-0.003	0.003	0.015	-0.061	0.178	0.000
	1987 Q2		40.899	0.000 *				0.253	-0.017	0.012	-0.233	0.356	0.001
	1987 Q3		26.673	0.000 *				-0.003	-1.100	0.012	0.770	-0.603	-0.004
Semiconductors	**1986 Q3**	**STA 1**	10.128	0.038 *	0.973	-0.142	0.044	1.416	-1.283	0.042	1.134	-0.812	0.061
	1986 Q4		2.905	0.574				1.198	-0.702	0.017	1.176	-0.981	0.065
	1987 Q1		4.374	0.358				1.343	-1.074	0.017	1.203	-1.092	0.068
	1987 Q2		5.492	0.240				1.634	-1.814	0.007	1.311	-1.560	0.080
	1991 Q3	**STA 2**	8.405	0.078	0.973	-0.142	0.044	1.253	-0.786	0.043	1.240	-1.071	0.059
	1991 Q4		8.386	0.078				1.372	-1.088	0.036	1.276	-1.244	0.064
	1992 Q1		7.185	0.126				1.009	-0.191	0.050	1.136	-0.748	0.054
	1992 Q2		8.187	0.085				0.723	0.538	0.062	0.872	0.004	0.042
Medicine	**1986 Q1**	**MOSS**	10.170	0.038 *	1.348	-0.956	0.004	1.434	-1.173	0.013	1.400	-1.076	0.003
	1986 Q2		10.509	0.033 *				1.333	-0.933	0.010	1.401	-1.082	0.004
	1986 Q3		12.455	0.014 *				1.321	-0.907	0.010	1.413	-1.111	0.004
	1986 Q4		12.309	0.015 *				1.291	-0.836	0.009	1.415	-1.117	0.004
	1994 Q4	**PSP**	17.565	0.002 *				1.213	-0.658	0.008	1.372	-1.033	0.005
	1995 Q1		17.123	0.002 *				1.216	-0.666	0.008	1.372	-1.033	0.005
	1995 Q2		16.717	0.002 *				1.221	-0.676	0.008	1.371	-1.033	0.005
	1995 Q3		13.926	0.008 *				1.224	-0.682	0.007	1.377	-1.046	0.005
Paper	**1992 Q2**	**PAP**	15.238	0.004 *	0.948	-0.250	0.008	0.593	0.539	0.018	0.989	-0.348	0.007
	1992 Q3		15.002	0.005 *				0.591	0.546	0.018	1.009	-0.404	0.007
	1992 Q4		14.749	0.005 *				0.582	0.563	0.018	0.946	-0.235	0.006
	1993 Q1		16.044	0.003 *				0.616	0.490	0.017	1.251	-1.054	0.010
Passenger Cars	**1987 Q3**	**MOSS**	7.076	0.132	-3.812	11.121	0.055	6.726	-14.278	-0.016	7.549	-15.113	-0.065
	1987 Q4		5.629	0.229				76.646	-183.821	-0.530	7.643	-15.349	-0.065
	1988 Q1		4.464	0.347				-8.839	23.487	0.098	8.367	-17.096	-0.068
	1988 Q2		5.490	0.241				-28.783	72.155	0.249	6.343	-12.510	-0.042
	1992 Q1	**A&APP**	13.120	0.011 *				35.253	-84.209	0.166	3.323	-5.590	-0.007
	1992 Q2		15.131	0.004 *				-5.836	16.161	0.023	4.058	-7.392	-0.008
	1992 Q3		14.636	0.006 *				-5.007	14.187	0.030	4.123	-7.579	-0.006
	1992 Q4		17.900	0.001 *				-8.774	23.470	0.012	4.269	-8.026	-0.001
	1995 Q3	**JAA**	14.905	0.005 *				0.169	1.495	0.050	-10.576	27.390	0.077
	1995 Q4		17.910	0.001 *				0.118	1.624	0.049	9.490	-19.528	-0.087
	1996 Q1		13.656	0.008 *				-0.218	2.450	0.048	5.635	-10.654	-0.047
	1996 Q2		12.229	0.016 *				-0.612	3.421	0.047	4.946	-9.147	-0.035

					Normalized Cointegrating Vector								
	Break Test				Constant Parameters			Pre-break sample Parameters			Post-break Sample Parameters		
Category	Break period		LR	P-val	Income	Rel Price	Trend	Income	Rel Price	Trend	Income	Rel Price	Trend
Auto Parts	**1987 Q3**	**MOSS**	9.554	0.049 *	1.443	-1.574	0.018	7.749	-16.973	-0.074	-2.893	8.458	0.061
	1987 Q4		9.214	0.056				4.436	-8.910	-0.020	2.705	-4.338	-0.006
	1988 Q1		8.823	0.066				3.655	-7.095	-0.016	-0.491	3.035	0.028
	1988 Q2		10.958	0.027 *				3.686	-7.160	-0.021	-1.001	4.253	0.031
	1992 Q1	**A&APP**	12.278	0.015 *				1.201	-2.268	0.033	1.956	-2.801	0.015
	1992 Q2		13.970	0.007 *				1.337	-1.346	0.033	2.131	-3.275	0.017
	1992 Q3		10.945	0.027 *				1.583	-1.957	0.032	2.670	-4.635	0.019
	1992 Q4		9.171	0.057				4.937	-10.067	0.052	-24.655	63.581	-0.029
	1995 Q3	**JAA**	5.743	0.219				1.395	-1.453	0.019	1.530	-1.701	0.011
	1995 Q4		5.972	0.201				1.314	-1.270	0.020	1.626	-1.932	0.010
	1996 Q1		5.101	0.277				1.116	-0.803	0.022	1.466	-1.533	0.010
	1996 Q2		4.726	0.317				1.117	-0.809	0.022	1.478	-1.566	0.010
Lumber	**1986 Q1**	**MOSS**	3.537	0.472	0.813	0.216	0.013	1.275	-0.773	0.003	0.960	0.028	0.003
	1986 Q2		3.939	0.414				1.424	-1.153	0.005	0.938	0.025	0.006
	1986 Q3		4.953	0.292				1.376	-1.029	0.004	0.996	-0.116	0.006
	1986 Q4		4.851	0.303				0.286	2.182	-0.056	-0.760	3.321	0.065
	1990 Q2	**WPA**	5.834	0.211				0.902	0.056	0.009	0.728	0.624	0.001
	1990 Q3		8.700	0.069				0.961	-0.088	0.005	0.716	0.659	0.001
	1990 Q4		12.017	0.017 *				1.100	-0.410	0.003	0.790	0.487	0.000
	1991 Q1		14.600	0.006 *				1.046	-0.289	0.003	0.806	0.458	-0.001
Total Non-Mineral	**1985Q3**	**Plaza Accord**	9.368	0.053	1.181	0.192	0.014	1.915	-1.390	-0.002	1.463	-0.484	0.011
	1987 Q1	**Louvre Accord**	11.896	0.018 *				1.588	-0.685	0.000	1.375	-0.281	0.013
	1990 Q2	Mid-Sample	6.596	0.159				1.251	0.029	0.006	1.274	-0.032	0.013

Note: All regressions were run with four lagged difference terms. Asterisk indicates a break is detected at the 5% significance level.

Appendix Table A1. Average Tariff Rates for Individual Import Categories

Fiscal Year	Medicine	Semiconductors	Tobacco	Paper	Lumber	Passenger Cars	Auto Parts	Non-Energy Total
1980	7	7.1	90	6.5	3.75	0	5.3	2.1
1981	6.9	8.6	35	6.7	3.75	0	0	2.1
1982	6.4	8.9	35	5.9	3.39	0	0	2.2
1983	6.3	5.6	20	5.6	3.39	0	0	2.1
1984	6	4.2	19.7	5.2	2.72	0	0	2.1
1985	5.4	0	20.4	4.9	2.39	0	0	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	2.00	0	0	2.4
1991	2.2	0	0	3.1	2.00	0	0	3.0
1992	2	0	0	3.1	2.00	0	0	3.1
1993	2	0	0	3.1	2.00	0	0	3.3
1994	2	0	0	3.1	2.00	0	0	3.1
1995	0	0	0	3	2.00	0	0	2.9
1996	0	0	0	2.8	2.00	0	0	2.6
1997	0	0	0	2.3	2.00	0	0	2.4
1998	0	0	0	2.1	2.00	0	0	2.5
1999	0	0	0	1.8	2.00	0	0	2.2
2000	0	0	0	1.4	2.00	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, but are not reported in this table.

Appendix Table A2. Unit Root Tests

	ADF, no trend	ADF, trend	P-P, no trend	P-P, trend
Medicine Imports	-1.59	-1.21	-1.46	-1.66
Medicine RP	-1.84	-2.41	-2.05	-2.22
Semiconductor Imports	-0.19	-2.35	-0.21	-0.21
Semiconductor RP	-1.49	-1.81	-1.45	-1.42
Tobacco Imports	-1.62	-0.70	-1.28	-0.77
Tobacco RP	-1.87	-1.29	-2.53	-1.13
Paper Imports	-2.15	-2.17	-1.28	-2.74
Paper RP	-1.72	-1.24	-1.32	-1.66
Lumber Imports	-1.86	-1.38	-1.51	-1.90
Lumber RP	-2.01	-1.99	-2.97*	-2.93
Passenger Cars Imports	-1.67	-1.50	-0.85	-1.38
Passenger Cars RP	-1.83	-2.22	-1.85	-2.59
Auto Parts Imports	-1.16	-2.00	-0.96	-2.43
Auto Parts RP	-1.77	-2.25	-1.92	-2.69
Total Non-Energy Imports	-0.24	-2.61	-0.07	-2.12
Total Non-Energy RP	-1.12	-2.13	-1.80	-2.30
Real GDE	-1.77	-0.39	-2.07	0.60

Note: All tests shown are in levels. Augmented Dickey-Fuller (ADF) tests used four lags and Phillips-Perron (P-P) tests used a truncation lag of three. Lag length selection for the ADF tests was based on AIC and SC. For the P-P tests, the automatic truncation lag suggested by Newey and West (1994) based on the number of observations is employed. All t-stats (save one) fall well below MacKinnon critical values for 10%, 5% and 1%. Tests on first differences of all variables (not reported here) strongly rejected the unit root hypothesis at a 1% level.

* significant at 5% level.

Endnotes

[1] Japan has a long history of friction with foreign 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 recent 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] This is the so-called "Armington Assumption" commonly employed in estimation of trade equations, which first appeared in Armington (1969). 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).

[4] While commonplace, the log-linear specification is not without its detractors. See, for example, Marquez (1999).

[5] See Campbell and Perron (1991) for a summary of the importance of unit roots and cointegration in time series data.

[6] 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] The tests for parameter constancy presented here were done using the Maximum Likelihood Analysis of Cointegrated Processes (MLECO) GAUSS code obtained from Peter Hansen's web site at http://www.econ.brown.edu/fac/Peter_Hansen/. Other preliminary diagnostic tests were done in Eviews or PcGive.

[8] For a general discussion of likelihood ratio tests see Davidson and MacKinnon (1993).

[9] We have included four lags of each variable in the VAR when estimating the cointegrated systems. As is well known in the literature, incorrect lag specification can cause size distortion or loss of power (Maddala and Kim, 1999). Results for several common selection criteria (AIC, SC, and HQ) varied considerably, but generally ranged from three to five lags, with four or five lags most common. Given the relatively small sample size, we have decided to err on the side of parsimony. Full sample elasticities varied little whether four or five lags were used.

[10] 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 years

(1980-2000), 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 years, the average tariff rate has declined very little.

[11] There is some debate as to whether or not seasonally adjusting the series, aggregation, and log transformations are appropriate. For a discussion of the advantages and disadvantage of each, see Maddala and Kim (1999).

[12] In all cases, four lagged difference terms were included in the vector error correction models. See note 9, above.

[13] Cointegration rank tests under structural change do not exist in the literature (Hansen, 2001).

[14] This interpretation is supported by an inspection of adjustment (loading) coefficients in the three-variable vector error correction model. In all cases, the adjustment coefficient on the error correction term in the import equation has the expected negative sign. For the real GDP and relative price equations, in about half the cases adjustment coefficients are positive. In fact, we have not identified structural import demand equations from the VAR system, so interpretation of the cointegrating relationship could actually represent a linear combination of both demand and supply effects.

[15] Goldstein and Kahn (1985), surveying empirical research from 1973 to the early 1980s, report an average income elasticity for aggregate imports of 1.2 and an average price elasticity of –0.95 (for those studies with a significant negative price elasticity). Hooper and Marquez (1995) report a mean import price elasticity of –0.97 from 13 studies dating back to 1996. The range of estimates is very large, for example price elasticity estimates range from –3.4 to –0.26 in the studies reviewed by Hooper and Marquez. More recent estimates of Japanese income (price) elasticities include Assery and Peel (1991), 1.36 (-0.64); Ceglowski (1996), 0.73 (-0.67); and Caporale and Chui (1999), 1.33 (-0.33). Elasticity estimates also vary widely for different commodities. See, for example, Ceglowski (1996).

[16] Inclusion of a time trend reduces unrealistically large income elasticities in a number of sectors, suggesting the importance correctly identifying deterministic sources of import growth separate from the influences of income and relative prices.

[17] This result differs from that found in Parsons (2002) which found no structural break in either the long run cointegration parameters, nor in the adjustment coefficients. One could argue that the more focused exogenous testing procedure employed here is more powerful given the use of *a priori* information concerning the timing of the policy. Yet, it also highlights the sensitivity of results to differing econometric methodologies, particularly in small samples.