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Are the U.S. Exports to and Imports from Japan Cointegrated?

Yangru Wu
Rutgers University

Junxi Zhang
University of Hong Kong

Abstract

The size and duration of the U.S. bilateral trade deficit with Japan has raised concern from both politicians and the general public. This paper seeks to investigate the behavior of this deficit by conducting stationarity tests on the deficit and tests for long-run relationships between U.S. exports to and imports from Japan. We show that, if an endogenously searched break is properly accounted for, exports and imports are cointegrated with a coefficient of one, and the deficit appears to be stationary. Thus, in contrast to the public's perception, we conclude that the U.S.-Japan trade deficit may not be "too large." (JEL Classifications: F14, C22)

I. Introduction

One main issue that has affected the U.S.-Japan relations in recent years is the huge U.S. bilateral trade deficit with Japan. During the 1980s the Unit-

* Correspondence Address: Yangru Wu, Department of Finance and Economics, Faculty of Management, Rutgers University, Newark, NJ 07102-1820, U.S.A., (Tel) 973-353-1146, (Fax) 973-353-1233, (E-mail) yangruwu@gsmack.rutgers.edu; We are very grateful to two anonymous referees for helpful comments. The usual disclaimer applies.

ed States developed an excess of imports over exports in its trade with Japan and this deficit rose to over US$50 billion in many years. This trade gap does not seem to be shrinking in the 1990s, with the imbalance over US$60 billion in 1997, and is expected to grow even wider in the future. The trade statistics between the two countries from 1980 are given in Table 1. A casual observation reveals that the bilateral trade deficit has widened rapidly over the last decade or so.

In terms of its absolute magnitude, the current size of the trade deficit $60 billion is more than one third of the U.S. overall trade deficit. When compared to the nearest competitor – China, such an amount is more than doubled. On the other hand, in terms of its relative magnitude, this size approximately represents a non-trivial figure of 1% of the U.S. GDP. As a result, the size and duration of this deficit has alarmed both politicians and the general public in the United States, and many people simply conclude that it is "too large."

Table 1

US-Japan Trade Statistics

(in US$ million)

Year	Imports from Japan	Exports to Japan	Trade Deficit
1980	30,866	20,790	10,076
1981	37,655	21,823	15,831
1982	37,743	20,966	16,777
1983	41,183	21,894	19,289
1984	57,135	23,574	33,560
1985	68,782	22,630	46,152
1986	81,911	26,881	55,029
1987	84,574	28,248	56,326
1988	89,802	37,731	52,070
1989	93,585	44,583	49,002
1990	89,655	48,584	41,070
1991	91,582	48,146	43,436
1992	96,461	47,763	48,697
1993	107,267	47,949	59,318

Data Source: OECD's Monthly Statistics of Foreign Trade: Series A.

In the last few years, the U.S. government has strongly pressured Japan to open up its markets to American goods and threatened to use numeric targets to monitor Tokyo's progress. While the Japanese have responded with promises to reduce import barriers and make firm progress in domestic deregulation, the trade imbalance has not responded well. Consequently, the trade friction between the two countries has grown and produced much controversy and acrimony as well as intense public debate. As the American public increasingly bashes Japan for "unfair" trade practices and non-tariff measures that have helped shut out American goods, the Clinton administration is compelled to slap trade sanctions on Japan – putting pressure on the latter to relent or face a possible trade war.

The recent car trade dispute highlights the increasing trade friction between the two biggest economies of the world. The U.S. has long claimed that Japan excludes U.S. cars and auto parts with non-tariff barriers, chiefly in distribution and sales. Autos and auto parts account for three-fifths of the U.S. trade deficit with Japan. Although the dispute seems to be temporarily over, further disputes could take place unless Japan's surplus with the U.S. were trimmed to a level that would be widely accepted in America.

This issue has also generated interest among economists who look at it from a different perspective and seek alternative solutions. Unlike the public, who emphasizes the sheer size of the deficit and favors clashing policies such as imposing quantitative trade and surplus-reduction targets, economists are more concerned with the causes and sustainability of the deficit, as well as the likely consequences of a trade war on their respective economies and on the world trading system.[1] Interestingly, to date formal econometric tests on the time series properties of the deficit and the long-run tendency of the bilateral trade balance are relatively scarce admittedly tests for the overall U.S. current account deficit do exist, see Husted [1992] . This paper seeks to investigate the behavior of the deficit by focusing on the above two aspects, and the salient empirical findings are expect-

1. One example is that in July 1993 more than 50 economists, led by former GATT adviser Jagdish Bhagwati and including five Nobel Laureates, sent a letter to President Clinton and former Prime Minister Hosokawa by rejecting the U.S. demands for managed trade with Japan and urging both Japan and the U.S. to cooperatively task for their trade policies.

ed to shed some light on the on-going debate.

We conduct stationarity tests on the bilateral deficit and tests for cointegrating relations between U.S. exports to and imports from Japan. The cointegration approaches used here follow those of Park [1992] and Stock and Watson [1993]. We find that without a structural break, the deficit series appears to be non-stationary; however, if a structural break is included in the time series and cointegrating regressions, exports and imports are cointegrated with a coefficient of one, and the deficit is stationary. This contradicts the notion that Japan's trade surplus with America is "too large." To determine the crucial break date, we apply Zivot and Andrews' [1992] endogenous searching procedure as opposed to that suggested by Perron [1989], based on prior information.

The remainder of the paper is organized as follows. Section II proposes some simple theoretical motivations for the empirical tests employed in this paper, while Section III briefly describes the econometric procedure. Our empirical findings are given in Section IV. Finally, Section V concludes the paper.

II. Some Simple Theoretical Motivations

Our objective is to test for the presence (absence) of a long-run relationship between the time series of U.S. exports to and imports from Japan. As is well known, the presence of such a relationship suggests that the two series would never drift "too far" apart. Before describing the econometric procedure, we first explore some theoretical reasons which might imply such a relationship.

In order to derive a testable relationship between import and export, following Husted [1992], we write out a country's dynamic budget constraint which include activities in international markets as follows:

$$\Delta b_t = \frac{r_t - \quad_t}{1 + \quad_t} b_{t-1} + m_t - x_t , \qquad (1)$$

where r_t is the market (real) interest rate; m_t, x_t, and b_t denote ratios of import, export and international borrowings to output, respectively; γ_t, the growth rate of output at time t, is introduced to allow for possible different period-to-period growth.

Note that from (1), if the debt-to-GDP ratio, b_t, is difference stationary, then it is clear that long-run (cointegrating) relationships exist between exports, imports and debt servicing component. Furthermore, if b_t is stationary in levels, then exports and imports (excluding net interest payments) must be cointegrated.

Assuming that the world real interest rate and the growth rate are stationary with respective unconditional means of r and γ, eq.(1) then becomes,[2]

$$z_t + \frac{1+r}{1+\gamma} b_{t-1} = x_t + b_t, \tag{2}$$

where $z_t = m_t + [(1+r_t)/(1+\gamma_t) - (1+r)/(1+\gamma)]b_{t-1}$. Solving eq.(2) forward, as in Husted [1992], we get

$$m_t + \frac{r_t - \gamma_t}{1+\gamma_t} b_{t-1} = x_t + \sum_{j=0}^{\infty} \left(\frac{1+\gamma}{1+r}\right)^{j-1} (\Delta x_{t+j} - \Delta z_{t+j}) + \lim_{j\to\infty} \left(\frac{1+\gamma}{1+r}\right)^{j+1} b_{t+j}, \tag{3}$$

It is well known in the growth literature that in equilibrium the interest rate is higher than the growth rate (the difference equals the discount rate when the utility function is logarithmic). Thus, the factor $(1+\gamma)/(1+r)$ is smaller than one. Note that the LHS of (3) denotes the country's spending on imports as well as interest payments on debt.

Let $mb_t = m_t + [(r_t - \gamma_t)/(1+\gamma_t)]b_{t-1}$. Assuming that both x_t and z_t are difference-stationary processes and the limit term in (3) equals zero, it is straightforward to derive the following standard regression equation:

$$mb_t = a_0 + a_1 x_t + v_t \tag{4}$$

Under the null hypothesis that the economy's intertemporal budget constraint holds, we would expect that $a_1 = 1$ and v_t is stationary.

To conclude this section, we briefly discuss the testable implication of the theoretical motivation. As suggested above, if the country's intertemporal

2. It should be noted that both assumptions are empirically plausible. For the first assumption, Wu and Zhang [1996] document that interest rates are stationary for a panel of OECD countries, including the United States and Japan, the two countries under investigation in this study. Furthermore, Wu and Zhang [1997] find mean reversion in U.S. Treasury bill yields of many different maturities. As for the second assumption, it is widely believed among economists that per capita output can be classified either as a trend stationary or as a difference-stationary process, implying that output growth is a stationary series (see, for example, Christiano [1992]).

budget constraint holds, the amount that the economy borrows should be financed by the present value of the future trade surpluses; otherwise, the country bubbly finances its external debt.[3] Though it is economically possible for a country to run a permanent trade deficit against another country, the volume of trade deficit cannot grow without bound. In order words, in the long run, imports and exports should be cointegrated. Furthermore, politically, when the size of the trade deficit exceeds a threshold level, actions may have to be undertaken to deal with the situation. These considerations lend to our subsequent investigation of the long-run time series behavior of the bilateral trade flows between the U.S. and Japan.

III. Econometric Methodology

The principal statistical analysis pursued in this study concerns testing for the stationarity of the U.S. trade deficit with Japan. Throughout the paper, we focus on the material trade deficit rather the current account deficit for two reasons. First, the statistics of the bilateral current account between the two countries are not available. Second, it is widely known, see, for example, Obstfeld and Rogoff [1995], that from the end of World War I to fairly recent, for most countries, including the U.S. and Japan, the holdings of foreign assets have been limited both in quantity and scope. Therefore, compared to the quantities of commodity trade, the net capital gains on existing foreign assets are small and of secondary importance towards the calculation of current account balance. With this justification, the variable mb_t shall be replaced by m_t in our empirical specifications below. If the deficit contains a unit root, then it does not have a long-run mean. Innovations to the series can have a permanent effect and the series can potentially become arbitrarily large as time passes. On the other hand, if the deficit is stationary, one would expect its innovations to be of short-run consequences, and as time elapses there is a tendency for the series to revert to its mean. So, in this case the deficit will not be "too large."

The standard techniques to test for the stationarity of a time series are the augmented Dickey-Fuller (ADF) and the Phillips-Perron (PP) tests. As

3. For an alternative test of bubbles in international financial markets, see, *e.g.*, Wu [1995].

is well known, these tests have low power against local stationary alternatives in small samples, and hence failure to reject the null hypothesis of a unit root may not give one much information regarding whether the series is indeed stationary. Perron [1989] points out that these tests perform especially poorly when there is a break in the deterministic trend, and derives the asymptotic distribution of the test statistic which incorporates such a breaking trend. Perron's method, however, has received some criticism because the proposed break is chosen based on pretest examination of the data which leads his procedure to overstate the likelihood of the trend-break alternative. Zivot and Andrews [1992], among others, have introduced a method to endogenously search for a breakpoint and test for the presence of a unit root when the time series possesses a breaking trend, and it is this method which will be adopted in our study.[4] The procedure is briefly reviewed as follows.

Consider the null hypothesis:

$$d_t = \mu_0 + d_{t-1} + u_{d,t}, \tag{5}$$

where $\{d_t\}$ represents a measure of the U.S. trade deficit with Japan. The selection of the possible break point is viewed as the outcome of an estimation procedure designed to fit $\{d_t\}$ by a trend-stationary process with a one-time break in the trend occurring at an unknown point in time. The idea is to search for the break that gives the most weight to the trend-stationary alternative. Let the possible break be T_B, $1 < T_B < T$, the regressions are specified as follows:

$$\text{Model A}: \Delta d_t = {}^{A}_{0} + {}^{A}_{1} d_{t-1} + {}^{A}_{2} t + {}^{A}_{3} DU_t + \sum_{j=1}^{k} {}^{A}_{j} \Delta d_{t-j} + u_{d,t}, \tag{6}$$

$$\text{Model B}: \Delta d_t = {}^{B}_{0} + {}^{B}_{1} d_{t-1} + {}^{B}_{2} t + {}^{B}_{4} DT_t + \sum_{j=1}^{k} {}^{B}_{j} \Delta d_{t-j} + u_{d,t}, \tag{7}$$

$$\text{Model C}: \Delta d_t = {}^{C}_{0} + {}^{C}_{1} d_{t-1} + {}^{C}_{2} t + {}^{C}_{3} DU_t + {}^{C}_{4} DT_t + \sum_{j=1}^{k} {}^{C}_{j} \Delta d_{t-j} + u_{d,t}, \tag{8}$$

4. Other procedures have also been proposed to test for a unit root with an unknown changing point; see, *e.g.*, Christiano [1992]. These methods are based on similar principles to search for a breaking trend. It is not the intention of this paper to exhaust all these tests.

where DU_t=1 if $t > T_B$ and 0 otherwise; and $DT_t = t$ if $t > T_B$ and 0 otherwise. Models A and B allow for a break in the intercept and in the time trend, respectively, while model C includes the hybrid of the two. In each model, the k extra regressors are intended to eliminate possible nuisance-parameter dependencies in the asymptotic distributions of the test statistics caused by serial correlation in the error terms. To determine the break and compute the test statistic for a unit root, an OLS regression is run in each model with a break point at T_B, where T_B ranges from 1 to T−2. Then the t-statistic for testing whether the first-lag coefficient is zero, *i.e.*, μ^i=0, ($i = A$, B or C) is computed. The test statistic for a unit root is the minimum t-statistic over all T−2 regressions and the break point is the time corresponding to such a statistic.

Given the break T_B, we further test whether imports and exports are cointegrated with a cointegrating vector of one when it is properly incorporated. The cointegrating regression is specified as follows:

$$m_t = b_0 + b_1 DU_t + b_2 x_t + u_t,$$

where once again $DU_t = 1$ if $t > T_B$ and 0 otherwise. Tests will determine whether $b_2 = 1$.

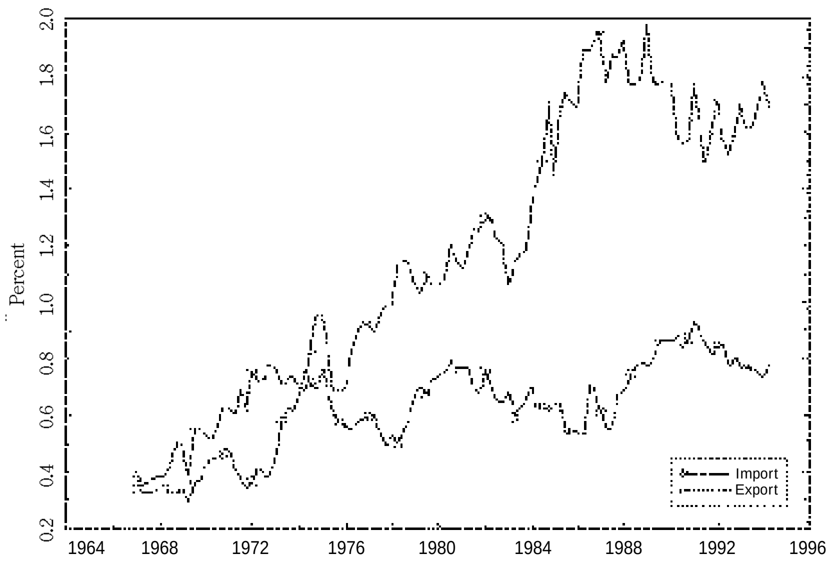
IV. Empirical Results

This section describes the data and presents our empirical findings, namely, results from stationarity and cointegrating analyses. As mentioned in the previous section, classical unit root tests are prone to structural breaks when policy alters. Therefore, this section documents both the results with and without such breaks.

A. Data and Preliminary Diagnosis

The data used in this study are quarterly flows of U.S. imports from and exports to Japan. The sample, dictated by data availability, covers the period from the fourth quarter of 1966 to the first quarter of 1994. The nominal imports and exports observations are obtained from the OECD's *Monthly Statistics of Foreign Trade: Series A*, while the U.S. GDP observations are

Figure 1
US-Japan Trade: Var/GDP Ratio



from IMF's International Financial Statistics (IFS). Our primary interest focuses on the trade deficit to GDP ratio, as this measure is more pertinent for a growing economy. Figure 1 presents the time-series plots of these two series. In addition to this share measure, we also consider two real measures of the trade flows in levels. For the first measure, the U.S. consumer price indices (CPI) are used to deflate both nominal imports and exports to convert them to real; while for the second measure, the U.S. import prices are used to deflate imports and export prices to deflate exports. The price series are taken from the IFS as well.

As a preliminary diagnosis of the data, we conduct ADF and PP tests for a unit root in each series under investigation. Table 2 reports the results. It is clear that for all six series examined, the null hypothesis of a unit root cannot be rejected even at the 10% level by any test statistic.[5] On the other hand, when the tests are applied to the first differences of these series, the null hypothesis can all be rejected at the 5% level or better. These results are

5. Note that the results are qualitatively the same even when a break point is incorporated. They are not reported here to conserve space but are available from the authors upon request.

Table 2
ADF and PP Tests for a Unit Root in Imports and Exports

		Var	*T*	*k*	*ADF*	*PP*
(1) Var/GDP Ratios	Without Time Trend	Import	110	8	-1.220	-1.257
		Export	110	10	-1.847	-2.196
	With Time Trend	Import	110	9	-2.441	-2.983
		Export	110	10	-2.643	-2.901
(2) CPI as the Index	Without Time Trend	Import	110	8	-0.642	-0.618
		Export	110	10	-1.027	-0.958
	With Time Trend	Import	110	9	-3.148	-2.979
		Export	110	10	-2.711	-2.929
(3) Import & Export Prices as the Indices	Without Time Trend	Import	110	10	-0.001	-0.003
		Export	110	10	-0.224	0.035
	With Time Trend	Import	110	4	-1.982	-2.183
		Export	110	4	-2.233	-2.305

Notes: 1. The optimum lag length, k, is selected as suggested by Campbell and Perron [1991].
2. Critical values, which are computed by using MacKinnon's [1990] method, are:

	10%	5%	1%
Without Time Trend	-2.58	-2.92	-3.44
With Time Trend	-3.16	-3.46	-4.01

again not reported to economize on space. Therefore, both import and export series can be characterized as difference-stationary series.

B. The Stationarity Results Without a Structural Break

In the absence of a structural break, we first examine the behavior of the trade deficit series. Standard ADF and PP tests are applied to all the three measures of the trade deficit series, and the results are presented in Table 3. These results feature regressions without and with time trend. Three comments are in order. First, in all cases, estimates of the first-lag coefficient in the deficit series are negative, suggesting that these coefficients are

Table 3
ADF and PP Tests for a Unit Root in the Trade Deficit

		T	*k*	*ADF*	*PP*
(1) Var/GDP Ratios	Without Time Trend	110	8	-1.28	-1.45
	With Time Trend	110	8	-2.60	-2.67
(2) CPI as the Index	Without Time Trend	110	10	-0.92	-1.12
	With Time Trend	110	9	-3.34	-2.84
(3) Import & Export Prices as the Indices	Without Time Trend	110	7	-0.59	-0.87
	With Time Trend	110	7	-2.30	-2.40

Notes: see Table 2.

smaller than one. Second, in all cases, the t statistics are unable to reject the null hypothesis of a unit root, even at the 10% significance level. Combining the two observations, it is clear that although the coefficients are smaller than one, they are not statistically significant. Third, these results are robust for both the ADF and PP tests over all the three measures of deficit.

To summarize, our findings imply that when a structural break is not taken into account, the trade deficit (as a share of the U.S. GDP) appears to be nonstationary and, in a sense, can become "too large." This might seem alarming for policy makers. Note that the nonstationarity result is rather robust, since it also applies to other measures of the deficit. It should be mentioned that we also have conducted tests without a break to see whether imports and exports are cointegrated, and the results are against the presumption of cointegration, which are not reported to conserve space.

C. The Stationarity Results with a Structural Break

As demonstrated by Perron [1989], the power for standard unit-root tests is rather low when the time series in fact possesses a break point. So our results in the preceding subsection may not be robust with respect to the choice of the date of the break. It is thus interesting to test for the existence of a possible break in the series of the real trade deficit, which proves to be essential to our results in this paper. The presence of a break date, possibly in the early 1980s, seems to be plausible and consistent with the simultaneous adoption of fiscal policy changes in the U.S. and Japan. The U.S. adopted very expansionary fiscal policy – the Reagan tax cuts and military spending increases, while Japan shifted to more expansionary policies. Such a view is popular among economists, see, *e.g.*, Bergsten [1991] and Masson *et al*. [1994].

The natural way to incorporate a break date is to rely on visual inspection of the data. However, it might be more preferable in many cases to use a formal statistical procedure to endogenously determine the break point. To this end, Zivot-Andrews' endogenously searching procedure has been utilized. As, *a priori*, no clue can be provided for whether the break is in the intercept, in the time trend or in both, we test all three cases as in eqs.(7)-(9), and the results are documented in Table 4. As for the procedure of determining the break and computing the test statistic, we run an OLS regression of each model with a break at T_B, where T_B spans from 1 to T-2. For each value of T_B, the number of extra regressors, k, is determined using the procedure suggested by Campbell and Perron [1991]. That is, start with a large $k_{\max}$ and then estimate the model with $k_{\max}$ lags. If the coefficient of the last included lag is significant at the 10% level, select $k = k_{\max}$. Otherwise, reduce the order of lags by one until the coefficient on the last included lag is significant.[6] Once the optimal lag length is chosen, the t-statistic for testing whether the first-lag coefficient is zero, *i.e.*, $\mu^i = 0$ ($i = A, B, C$) is computed. The test statistic for a unit root is the minimum t-statistic over all T-2 regressions and the break point is the time corresponding to such a statistic.

We have two main findings. First, in nearly all cases, the significant break is detected to be at the third quarter of 1983, which coincidentally is identical to that found using *ex ante* selection criterion by Husted [1992], who

6. It is worth mentioning that the choice of lag length k can affect the test results and other procedures to select the lag length k exist in the literature. In a recent paper, Ng and Perron [1995] demonstrate that an overly parsimonious model can have large size distortions, while an over-parameterized model may have low power. But the size problem is more severe than power loss. They show that methods based on sequential tests have an advantage over both the Said and Dickey's fixed rule and the information-based rules such as the Akaike information criterion and the Schwartz criterion because the former have less size distortions and have comparable power. The procedure adopted in this paper falls into this category of the general-to-specific sequential procedures.

Table 4
Zivot-Andrews Tests for a Unit Root in the Trade Deficit

1. Model A

	T	T_B	k	μ_0^A	μ_2^A	μ_3^A	μ_1^A
(1) Var/GDP Ratios	110	1983.III	9	0.00 (0.87)	0.00 (3.46)	0.00 (2.58)	-0.42 (-4.64*)
(2) CPI as the Index	110	1983.III	9	7.40 (0.59)	91.76 (4.15)	1.01 (2.83)	-0.37 (-5.47***)
(3) Import & Export Prices as the Indices	110	1983.III	9	23.37 (1.81)	93.08 (3.87)	0.47 (1.56)	-0.31 (-4.90***)

2. Model B

	T	T_B	k	μ_0^B	μ_2^B	μ_4^B	μ_1^B
(1) Var/GDP Ratios	110	1988.IV	8	-0.00 (-1.50)	0.00 (3.37)	-0.00 (-2.38)	-0.21 (-3.44)
(2) CPI as the Index	110	1983.III	9	7.70 (0.55)	0.84 (2.18)	0.83 (3.07)	-0.31 (-4.64**)
(3) Import & Export Prices as the Indices	110	1983.III	9	33.33 (2.27)	0.29 (0.91)	1.22 (3.87)	-0.33 (-4.92**)

3. Model C

	T	T_B	k	μ_0^C	μ_2^C	μ_3^C	μ_4^C	μ_1^C
(1) Var/GDP Ratios	110	1983.III	9	0.00 (0.39)	0.01 (3.89)	0.00 (3.23)	-0.00 (-2.97)	-0.40 (-5.33**)
(2) CPI as the Index	110	1983.III	9	-2.80 (-0.21)	186.23 (3.20)	1.34 (3.35)	-1.19 (-1.75)	-0.39 (-5.68***)
(3) Import & Export Prices as the Indices	110	1983.III	9	30.35 (2.01)	49.45 (0.92)	0.36 (1.09)	0.64 (0.90)	-0.33 (-4.90*)

Notes: 1. T_B is the break point. The dummy variables are defined as follows: $DU_t = 1$ if $t > T_B$ and 0 otherwise; $DT_t = t$ if $t > T_B$ and 0 otherwise.
2. For each choice of the breaking point T_B, the optimum lag length, k, is selected as suggested by Campbell and Perron [1991].
3. Numbers inside parentheses are t-ratios, where superscripts *, ** and *** indicate a significant test statistic at the 10, 5 and 1 percent levels, respectively.
4. For each model specification, critical values, which are obtained from Zivot and Andrews [1992], are:

	10%	5%	1%
Model A	-4.58	-4.80	-5.34
Model B	-4.11	-4.42	-4.93
Model C	-4.82	-5.08	-5.57

studies the U.S. aggregate current account deficit. This apparently reflects the fiscal policy changes in both countries in early 1980s and suggests that the behavior of the bilateral deficit differs greatly before and after the third quarter of 1983.[7]

Second, and more importantly, our results in Table 4 show otherwise when the structural break is incorporated in the data. In sharp contrast to the standard unit-root test results of Table 3, we are able to reject the null hypothesis of a unit root for nearly all cases. For example, in model A where a break is in the intercept, the t statistics are −4.64 if deficits are expressed as a ratio of GDP, −5.47 if CPI is used as the index, and −4.90 if import/export prices are used as the indices. The first figure is statistically significant at the 10% level, while the latter two are significant at the 1% level. The only exception that fails to reject the null hypothesis of a unit root even at the 10% significance level is model B (where a break is in the time trend) if ratios are used. Therefore, our results show that when a break in the deficit process during the early 1980s is accounted for, the deficit series appears to be stationary, and the conditions for intertemporal budget constraint for the United States seem to be satisfied.

D. The Cointegration Results

The preceding findings suggest that bilateral trade deficit does not contain a unit root when a break is properly incorporated. The unit root test for trade deficits imposes the assumption that imports and exports are cointegrated with a unit cointegrating vector. As Hendry and Mizon [1978] and Kremers, Ericsson and Dolado [1992] argue, imposing an unrealistic restriction can reduce the power of the test. Therefore, as a further confirmation, we also test whether imports and exports are cointegrated without restriction and we test whether the cointegrating vector is equal one given the break. Standard ADF and PP tests are then applied to the residuals of

7. Note that imports and exports should depend on the relevant relative prices and incomes. Shifts in those possibly integrated variables could give rise to the need for the structural break dummy variable. Identifying the factors causing the structural break is beyond the scope of this paper, but is of independent interest for future research.

regression (11) for testing cointegration and these results are presented in the upper panel of Table 5. It is found that for both ADF and PP tests, the null hypothesis of no cointegration is rejected at the 5% level when CPI's are used, and at the 1% level when import and export prices are used as deflators; as for the ratio measure, ADF and PP tests reject the null at the 10% and 5% levels, respectively.

Furthermore, we estimate the cointegration regression and test whether the cointegrating vector b_2 is equal to unity. Various estimators have been proposed in the literature. Because evidence is sketchy at this time as to which procedure has the best sampling properties, we consider two popular ones. These are Stock and Watson's [1993] dynamic OLS estimator and Park's [1992] canonical cointegrating regression estimator. Both estimators have the asymptotic standard normal distributions.[8] As shown in the lower panel of Table 5, this null hypothesis that $b_2 = 1$ cannot be rejected at 10% level by any of the test statistics. Therefore, our cointegration analysis provides further evidence that in the long run imports and exports under scrutiny tend to move together closely so that the trade deficit is indeed stationary.

Before concluding this section, we would like to make two important remarks. First, while we motivate the long-run relation between imports and exports, including debt servicing (see *eq*. (4)), on the two reasons explained in Section III, this third component is not included in the empirical analysis. The results reported in this section that imports and exports are cointegrated imply that the debt service account should be stationary. Furthermore, as our test is based on a bilateral basis, it is a stronger restriction than that obtained from the intertemporal budget constraint for the home country (see *eq*.(1)).

Second, we have focused our work on single-equation methods. Presumably, employing a multivariate method such as in Johansen [1988] to re-examine the hypotheses can be a useful investigation and is left for future research. Moreover, it is helpful to note that testing for cointegration is only one way to demonstrate that our satisfactory model specification has been obtained See the discussion in Hendry [1995] .

8. See these papers for the derivation of the asymptotic distributions.

Table 5
Cointegration Results

	Method	Var/GDP	CPI as the Index	Import & Export Prices as the Indices
Residual-Based Tests for Cointegration	ADF	-3.12*	-3.40**	-4.43***
	PP	-3.45**	-3.81**	-4.16***
Estimate of b_2	SW	1.272 (0.986)	1.281 (1.397)	0.973 (-0.123)
	CCR	1.117 (0.457)	1.198 (0.97)	0.897 (-0.59)

Notes: 1. The cointegration regression is as follows:
$m_t = b_0 + b_1 DU_t + b_2 x_t + v_t$,
where: $DU_t = 1$ for observations after 1983.III and 0 otherwise.
2. Critical values for test of cointegration, which are computed using MacKinnon's [1990] method, are: -3.09 for 10%, -3.40 for 5%, and -4.03 for 1%.
3. Numbers inside parentheses are the t-statistics for test of the hypothesis of b_2=1, where superscripts *, ** and *** indicate a significant test statistic at the 10, 5 and 1 percent level respectively. Under both the Stock-Watson (SW) and Park (CCR) methods, the t-statistic has an asymptotic standard normal distribution.

V. Conclusion

This paper has studied the emerging U.S. bilateral trade deficit with Japan by conducting stationarity and cointegration tests. We find that if a structural break, as detected to be present at the third quarter of 1983, is properly accounted for, the deficit is stationary. Our result is robust in that the nonstationary hypothesis has been rejected for a variety of trade measures over a number of model specifications.

In the wave of the current heated debate, our finding may be useful from a policy standpoint, albeit they must be necessarily tentative. As recognized by many economists, the public debate which centers on alleged "unfair trade policies" by Japan might be quite misleading *e.g.* Ethier [1988] . The result presented in this analysis provides some support for this view. It should be pointed out, however, that the finding that imports and exports are cointegrated does not necessarily represent, *prima facie*, evidence against "unfair trade policy." According to the Lerner symmetry argument, in general equilibrium, a change in import taxes acts like a change in export taxes, through its equivalent effects on relative prices. This implies that a more restrict trade regime in, say, Japan would act to lower both imports and exports.

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