The Japanese Import Demand Behavior : Stationarity and Cointegration**

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CONTENTS

Ⅰ. Introduction

In the area of international trade, there have been a number of studies designed to explain the import demand behavior. Previous studies of the demand for imports have generally used the traditional form of the estimating equation, i. e. one regressing the quantity of imports to the level of real income and the relative prices.The research on the import demand uniformly presumes that there exists an underlying equilbrium relationship between the quantity of imports and a given set of variables that explain it. However, if this assumption is

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1) Khan and Ross(1985)summarizes the literature

false, attempting to estimate the traditional formulation is invalid.

My purpose in this paper is to determine whether in fact the variables concerned with the Japanese import demand are (trend)stationary and there exists a cointegrating relationship. Though Khan and Ross (1975)thought of isolating the effects of the cyclical factors from those of the secular factors, they did not raise the issue in the context of non-stationarity and cointegration. Acknowledging the effect of the structural break such as the oil shock, my paper explores Perron(1989) type detrending method in the analysis of the import demand behavior. The resulting equation is estimated for Japan. The period of observation is 1974 : 1 to 1990 : 2 on a biannual basis.[2] Considering the exchange rate regime change and the 1st oil price shock, we excluded the period before 1974.

The plan of paper is the follwing. Section2 provides the results of the traditional estimation and the root tests. Scetion3 reports the results of the detrending method incorporating the structural shift and the implications of these results. Some conclusions are in section4.

II. The Demand for Imports and the Stationarity

The traditional formulation of an import demand equation relates the quantity of import demanded to domestic real income and the ratio of import prices to domestic prices. In log-linear terms the estimated result on the Japanese traditional import demand function is :

$$\log M(t) = -5.106 + .781 \log Y(t) - .456 \log P(t) \qquad (1)$$
$$(.749)\ (.061)\quad (.080)$$
$$R^2 = .908,\ D.W. = .512$$

, where M is quantity of import, Y is domestic real GNP, P = PM/PD where PM is import price index and PD is domestic wholesale price index. The values in the parentheses below the coefficient represent the standard errors. The data are taken from the International Financial Statistics.

2) Since the import demand is known to adjust to a shock in the right hand side variavbles with ususally more than a quarter, the biannual rather than the quarterly basis is chosen here.

The estimated result shows that both income and price elasticities show the expected signs. However, the above result following the traditional import demand literature just assumes the existence of an underlying equilibrium relationship between the quantity of imports and the right hand side variables that explain it. Recent advances in time series methodology allow researchers to test for the presence of equilibrium relationship between economic variables. Engle and Granger(1987)show that even if economic series may wander through time, there may exist some linear combination of the variables that converges to an equilibrium, i. e. the series are cointegrated.

We check the stationarity by the Dickey and Fuller(1979)normalized bias, $n(\alpha-1)$, and Phillips' (1987) $Z(\alpha)$ statistic which allows serial correlation as well as heteroscedasticity. (Table 1)reports the calculated statistics for each of the different series. The hypothesis of a unit root in the individual series is not rejected at the 5% level of significance for each concerned series.

To check the cointegration, Dickey-Fuller normalized bias and Phillips' $Z(\alpha)$test are applied to the residuals. The calculated values are -9.273 and -12.301, respectively. Therefore, we can't reject the null hypothesis of no cointegration at even 10% level of significance. This suggests that the traditional estimation results of the Japanese import demand equation such as Kreinin(1969), Goldstein and Khan(1976), Wilson and Takacs(1979)and so on are not reliable.

III. Results on the Detrending Method with Structural Break

When the concerned series are revealed to have unit roots, Perron(1989) suggested checking structural breaks and using the detrending method. Though Deyak, Sawyer and Sprinkle(1989) and Mah(1992, 1993) among others checked structural breaks in the import demand function, they did not relate the issue with the non-stationarity problem and the detrending method.

In this paper to overcome the non-stationarity problem found in the previous section, we use Perron(1989) type detrending method incorporating a structural break. Since plot-

ting, not reported here to save the space, shows breaks at the first half of 1986(1986 : 1), i. e. right after the Plaza meeting and the consequent yen appreciation, Perron(1989) type detrending method which acknowledges change in the form of level as well as the growth rate is adopted. Denote t : time trend, DU : dummy tating a value of one when t is greater than or equal to 1986 : 1 and zero otherwise, and DT : t −1985 : 2 and when t is greater than or equal to 1986 : 1and zero otherwise.

The estimated results on the so − called 'crash and changing growth model' are reported in (Table 2). Looking at the estimation results, since the coefficients of DU except that for the import quantity are significantly different form zero at any reasonable level of significance, we can suggest that the levels of log Y and log (PM/PD) variables changed at around 1985 : 2 − 1986 : 1, the period of the Plaza meeting when the appreciation of Japanese yen was decided. The slope of the import quantity (the relative price) against trend changed (did not change) at the 5% level of significance. Whether that of the real income change or not is difficult to decide.

The estimation result from the crash and changing growth type detrended series is shown in eq. (2) :

$$\log M(t) = .000 + 2.448 \log Y(t) - .247 \log P(t) \qquad (2)$$
$$(.007)\ (.562)\quad (.074)$$
$$R^2 = .552,\ D.W. = 1.326$$

Comparing eq. (2) with eq. (1), first, the Durbin−Watson statistic ameliorates presumably due to the acknowledgement of the structural break. Second, the coefficient of the real income increases and the absolute value of the coefficient of the relative price is small in absolute value, whose low value is generally in accordance with the current knowledge concerning the Japanese import demand.[3)]

Though we can infer the meaning of the elasticities of income and price from eq. (2) or compare eq. (2)with eq. (1), the non−stationarity issue still remains. The unit root test results on the detrended series are shown in (Table 3). The null hypothesis of the existence of a unit root is not rejected for each of the variables at any reasonable level of significance. To

3) See Goldstein and Khan (1976) as an example.

check the cointegration of the model, Dickey−Fuller normalized bias and Phillips' $Z(\alpha)$ test are applied to the residual of the cointegrating regression. The calculated values are −8.451 and 15.524, respectively. We cannot reject the null hypothesis of no cointegration at any reasonable level of significance. The reported results suggest that, even if we acknowledge the structural break following the Plaza agreement on the appreciation of Japanese yen and detrend the data, the estimation result of the traditional Japanese import demand equation is not reliable.

Since the changing growth type detrending method may be more valid with respect to the import quantity term and the crash type may be more relevant for the relative price term, several combinations of detrended series are tried. However, all these trials yielded poorer performances than that reported in eq. (2) and also showed non−stationarity and non−cointegration.

In addition to the import demand function formulation used previously, we also used log Y(t), log PM(t) and log PD(t) as the appropriate right hand side explanatory variables, acknowledging the possible different responses of the import quantity with respect to the import price and the domestic price. The evidences on the unit root tests and the cointegration test appear in (Table 4). Even though Perron (1989) type detrending method is adopted with respect to the regression based on the 3 right hand side variables, as (Table 5) shows, the results show that all concerned detrended variables are also non−stationary. The cointegrating regression shows no rejection of cointegration at the 5% level of significance in terms of Dickey−Fuller normalized bias. However, Phillips' $Z(\alpha)$ test shows non−cointegration at any reasonable level of significance.[4] Furthermore, the estimated coefficients of the import price and the domestic poice were not consistent with a priori expectation. The interested reader can get more results upon request from the author.

The above−mentioned results showed that all the relevant series are still non−stationary despite the use of Perron type detrending method, which are quite in contrast with Perron's (1989) findings with respect to most U. S. macroeconomic variables. Furthermore, the model is shown as not cointegrated. If the time series involved are cointegrared, then the

4) Phillips and Ouliaris(1990) suggested $Z(\alpha)$ test in the sense of power and supported setting a null hypothesis of no cointegration. Their suggestion is adopted here.

regression parameters can be precisely estimated, because the estimated values under cointegration converge to their true values very guickly. In the absence of cointegration the steady state nature of the underlying model is doubtful, because it has no inherent tendency of returning to the equilibrium path.

The evidences presented in this paper suggest that the conventionally used market force variables such as the real income and the relative prices cannot explain the Japanese import demand behavior well.[5] Looking at the year of, for example, 1983 (Nogues, Olechowski, and Winters (1985)), the official Japanese non−tariff barrier coverage ratio is fairly low (11.9%) compared with that of all industrial country markets (27.1%). Consequently, in addition to the market force terms, the unofficial private barrier such as the peculiar distribution system not related with the Japanese official liberalization measure might be an additional effective variable in explaining the Japanese import demand behavior.[6]

Ⅳ. Conclusion

Adopting the elasticity approach, the import demand equation is analyzed with respect to Japan using biannual data over the officially well liberalized period. The unit root tests showed that all concerned variables are non−stationary and not cointegrated as well. Though Perron (1989) type detrending method acknowledging a structural break is adopted to overcome the problem, the result also shows that the detrended series are non−stationary and, furthermore, not cointegrated. It implies that the traditionally used market force variables cannot explain the Japanese import demand well and, the closed minds rather than the closed market or the private barrier such as the complex domestic distribution system might be the peculiar aspect of the Japanese import demand.

5) For a general interpretation of the non−cointegration, see Granger (1986).

6) Balassa (1986). Lawrence (1987) and others also concluded in this way, though not using rigorous econometric techniques. See Saxonhouse (1982) for the opposite side.

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Table 1 : Test Results for Unit Roots and Cointegration

Statistic	log M	log Y	log(PM/PD)	E
Dickey−Fuller n(α−1)	.149	.063	1.096	−9.273
Phillips'Z(α)	.144	.063	.661	−12.301

Notes : a) E denotes the residual term in the cointegrating regression.

b) The reported results in all tables are based on 4 lag truncation numbers. Though several other truncation numbers were also tried, the test results did not change at all qualitatively.

Table 2 : The Estimation Results on the Crash and Changing Growth Model

Dependent variable	constant	trend	DU	DT	R^2
log M	4.441	.010	.042	.041	.947
	(.024)	(.002)	(.045)	(.007)	
log Y	12.029	.024	−.038	.003	.997
	(.005)	(.000)	(.010)	(.001)	
log (PM/PD)	−.228	.010	−.431	−.007	.729
	(.040)	(.003)	(.076)	(.011)	

Note : The values in the parentheses below the coefficient denote the standard errors of the coefficients. DU = 1 if t is greater than or equal to 1986 : 1 and = 0 otherwise. DT = t − 1985 : 2 if t is greater then or equal to 1986 : 1 and = 0 otherwise. DU and DT check the crash (the level change) and the change in the growth rate against trend, respectively.

Table 3 : Test Results for Unit Roots and Cointegration : the Detrended Series Case

Statistic	log M	log Y	log(PM/PD)	E
Dickey − Fuller n(α−1)	−15.024	−9.763	−12.590	−8.451
Phillips' Z(α)	−3.928	−7.333	−2.999	15.524

Notes : See Perron (1989) for the critical values.

Table 4 : Test Results for Unit Roots and Cointegration : 3 Right Hand Side Variables Case

Statistic	log PM	log PD	R

Dickey − Fuller n(α−1)	.012	0.047	−11.626
Phillips' Z(α)	.006	.046	−4.010

Note : R denotes the residual term in the cointegrating regression in case of using 3 right hand side variables : log PM, log PD, and log Y.

Table 5 : Test Results for Unit Roots and Cointegration : 3 Detrended Right Hand Side Variables Case

Statistic	log PM	log PD	R
Dickey − Fuller n (α−1)	−12.720	−12.720	−29.764
Phillips' Z(α)	−3.045	−4.847	−.745

국 문 요 약

1985년 후반기에 발생한 구조변화효과를 감안하여, 일본수입수요와 유관한 변수들에 대하여 페론타입의 추세제거방법이 적용된다. 추세가 제거된 변수들은 불안정이고 공적분이 이루어지지 않기 때문에 기존의 연구결과들은 모두 무효가 될 가능성이 있는 것으로 보인다. 따라서 공식적으로 개방의 폭이 큰 일본수입수요를 설명함에 있어서 전통적으로 사용되는 시장변수들은 불충분한 것으로 판단된다. 닫힌 마음, 혹은 일본고유의 국내유통구조와 같은 사적 장벽이 일본의 수입수요를 설명하는데 주요한 변수가 아닌가 생각된다.