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## Documento de Trabajo

**Trade balances:  
Do exchange rates matter?**

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**TRADE BALANCES:**

**DO EXCHANGE RATES MATTER?**

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**ABSTRACT**

The concern about the persistence of the US trade deficit has generated a lively debate over its causes and the role of exchange rates in restoring external balance. While some authors argue that the traditional adjustment process linking the behavior of the trade balance to movements in the real exchange rate and the domestic and foreign income levels has worked, recent research has found no empirical evidence in support of a stable long-run relationship between these variables. In this paper I use the approach in Johansen (1988, 1991) to test for cointegration between US bilateral trade balances, exchange rates and incomes during the floating period. I also apply the testing procedure developed by Hansen (1992) to detect potential structural breaks in a cointegrating relationship in the bilateral trade balance equations. The empirical results indicate the following: (1) the presence of stable long-run trade balance equations in all six models, (2) no significant evidence of structural breaks, (3) the nominal exchange rate has better explanatory power than the real rate, (4) the Marshall-Lerner condition is supported by the data, (5) the exchange rate is weakly exogenous in the trade balance relationship and, (6) the traditional belief that trade balances adjust slowly to exogenous shocks is confirmed; however, the speed of adjustment varies significantly across bilateral models.

**RESUMEN**

La preocupación acerca del déficit comercial Norteamericano ha suscitado un animado debate sobre sus causas y el posible papel de una política cambiaria para restaurar el equilibrio externo. Mientras algunos autores mantienen que el proceso de ajuste tradicional que liga el comportamiento de la balanza comercial a las fluctuaciones en el tipo real de cambio y a los niveles de renta domésticos y extranjeros ha operado, recientes investigaciones empíricas no han podido encontrar evidencia de tal relación ni siquiera en el largo plazo. En el presente trabajo utilizo el enfoque de Johansen (1988, 1992) para contrastar la presencia de cointegración entre balanzas comerciales bilaterales Norteamericanas, el tipo de cambio y rentas durante el periodo de tipos de cambio flexibles. Utilizo también los contrastes de Hansen (1992) para detectar posibles cambios estructurales en las ecuaciones de cointegración de las balanzas comerciales. Los resultados empíricos indican lo siguiente: (1) la existencia de relaciones estables a largo plazo en las seis balanzas bilaterales analizadas, (2) la ausencia de cambios estructurales significativos, (3) el tipo de cambio nominal tiene mayor impacto en las balanzas comerciales que el tipo real, (4) la condición de Marshall-Lerner encuentra apoyo en los datos, (5) el tipo de cambio es exógeno en sentido débil en la ecuación de la balanza comercial y, (6) la creencia aceptada de que las balanzas comerciales responden lentamente a los shocks exógenos se ve confirmada, sin embargo, la velocidad de ajuste varía de forma significativa entre modelos.

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## I.- INTRODUCTION

The purpose of this paper is to investigate whether or not a stable long run relationship between U.S. bilateral trade balances, exchange rates and domestic and foreign incomes has held during the floating period. The role of exchange rates in trade balance behavior is of primary interest. The paper studies how important exchange rate fluctuations are in explaining the long run behavior of the trade balance and whether or not exchange rates can be considered exogenous in such a relationship. Both issues are fundamental in assessing the impact of exchange rate policies in the management of trade imbalances.

Over the years an extensive empirical literature has developed attempting to model trade balance behavior. Exchange rates and real incomes have traditionally been considered the essential determinants in the trade balance literature. In particular, the estimation of income and price elasticities in the demand for imports and exports has been a primary objective. Such information is crucial in determining the effectiveness of exchange rate policies to correct trade imbalances or to better understand the role of the goods market in the international transmission of disturbances. Recently, the concern about the persistence of the U.S. trade deficit has generated a renewed interest in this literature. The debate has centered on the causes of the deficit and the role of exchange rates in restoring external balance. Some authors, e.g. Krugman (1991) and Lawrence (1990), maintain that the adjustment process described in traditional trade balance models has worked and that the apparent lack of response of the U.S. trade balance to the drastic devaluation of the dollar since 1985 has been caused by the long lags in the adjustment process. Other authors, e.g. Baldwin (1988) and Baldwin and Krugman (1989), appeal instead to the notion of hysteresis to explain the persistence of the U.S. trade deficit. According to the hysteresis argument the large exchange rate

fluctuations in the last two decades may have caused permanent changes in the structure of world trade. Finally, there are researchers who defend that there is no theoretical foundation for the existence of a stable relationship between the exchange rate and the trade balance. For example, proponents of the equilibrium approach to exchange rates, e.g. Stockman (1987, 1990), argue that it is the nature of shocks that determines whether a positive, a negative or independent relation will be observed between the exchange rate and trade balance movements. The different explanations suggest substantially different testable implications concerning the long run behavior of the trade balance. The first argument implies the existence of a stable long run trade balance equation. On the contrary, the third argument would tend to reject the stability of traditionally specified trade balance models. Finally, the hysteresis hypothesis implies that while traditional models may work structural breaks in the trade balance relationship should be detected over extended periods of time.

Current econometric knowledge suggests that the techniques used in past research may not be appropriate for the analysis of long run trade balance behavior. On the one hand, exchange rates and incomes are usually assumed exogenous with respect to the trade balance, a tenuous assumption in a general equilibrium framework.<sup>1</sup> In addition, previous empirical analysis has often been performed using the levels of the relevant variables as regressors, even though these variables are typically nonstationary.<sup>2</sup> The potential lack of exogeneity and the nonstationarity of regressors are likely to induce substantial biases in the estimates of both the parameter values and the standard errors obtained with traditional econometric methods.

It is only in the last decade that appropriate econometric techniques to deal with nonstationary environments have been developed. Most useful among these

are the different methods to estimate long run relationships, such as cointegrating relationships, and their appropriate short run dynamic representation, the error correction model (ECM). The new techniques have been recently applied to the study of trade balance behavior; however, the results are still inconclusive. Clarida (1991), for example, finds evidence of stable behavior in U.S. aggregate imports over the last two decades. Similarly, Hickok, Hung and Wulfekuhler (1991), using rather large specifications to accommodate demand and supply side factors, find stable U.S. import and export equations for the same period. On the contrary, Yellen and Rose (1989) and Rose (1991) find no significant evidence of stable long run bilateral and aggregate trade balance equations for the U.S.. Furthermore, they find no short-run and long-run exchange rate influences on the trade balance. These results are somewhat surprising in view of the apparent positive co-movements between bilateral trade balances and exchange rates during the floating period as seen in figures 1-6. The failure to find stable long run trade balance equations may be explained in part by the choice of techniques in testing for cointegration. For example, Yellen and Rose (1989) and Rose (1991) employ the Granger and Engle (1987) two step OLS procedure and the Stock and Watson (1988) principal components procedure. However, as described by Gonzalo (1991), these two approaches are among the least powerful to detect cointegration.<sup>3</sup> A second potential problem in both of these papers is the choice of variables. For example, the GNP deflator, which relies heavily on finished goods and includes a large proportion of nontradable services, is used to deflate nominal trade balances and to compute the real exchange rate. Finally, the sample period considered 1960-1985 may have witnessed structural changes in the trade balance equation not only due to the change in exchange rate regime in 1973 but also due to the potential hysteresis in trade associated with the large exchange rate fluctuations during the floating period.<sup>4</sup>

potential problems caused by third country effects this type of analysis requires the computation of composite (effective) exchange rates and world output measures, a procedure that is likely to introduce large measurement errors. Comparatively, bilateral trade data in developed countries is likely to contain less measurement errors.<sup>6</sup> The use of bilateral data also avoids the problems associated with trade in commodities such as oil and agricultural products.

The empirical analysis yields the following results. First, I find a stable long-run specification of the trade balance equation in all six bilateral models. In the long-run specifications the nominal exchange rate exhibits better explanatory power than the real rate. The Marshall-Lerner condition is upheld by the data, with exchange rate elasticities in the neighborhood of unity. In addition, the exchange rate appears to be weakly exogenous in most bilateral relationships. Income elasticities tend to be higher than previously reported and often the equality between the domestic and foreign income elasticities cannot be rejected. Finally, the traditional belief that trade balances adjust slowly to exogenous shocks is confirmed; however, the speed of adjustment varies significantly across bilateral models.

The paper is organized as follows. In section 2, alternative theoretical foundations for the trade balance equation used in the empirical analysis are presented. The choice of variables is also discussed in this section. In section 3, I present a brief description of the econometric techniques used in the study. Section 4 begins by describing the data and the strategy used in the search for an appropriate specification of the long run behavior of the trade balance. The main results are then presented followed by an analysis of structural breaks. Finally, section 5 discusses the results, formulates some recommendations and describes future areas of research.

## II.- THE TRADE BALANCE EQUATION

The benchmark trade balance equation estimated in the paper is a linear version of the general expression,

$$TB = TB(R, Y, Y^*), \quad (1)$$

where  $TB$  is a measure of the trade balance in real terms,  $R$  is some measure of relative prices (e.g., the real exchange rate), and  $Y$  ( $Y^*$ ) is a measure of domestic (foreign) real income. An equation like (1) is at the core of almost all empirical models of the trade balance.<sup>7</sup> It is often derived from a two country version of the imperfect substitutes model in which the demand and supply for domestic exports and imports are assumed to be a function of relative prices and incomes. More recently, Clarida (1991) has shown that under certain assumptions a similar trade balance equation can be derived in the context of a two (nondurable) goods version of a representative agent, permanent income model. Consequently, equation (1) can be interpreted as an equilibrium (structural) relationship. However, it is not a reduced form equation since as indicated earlier relative prices and incomes are not necessarily exogenous. If the Marshall-Lerner condition holds, the long run price elasticity of the trade balance is expected to be positive. The signs of the domestic and foreign income elasticities, on the contrary, are less clearly defined. If income changes affect primarily the demand side the domestic income elasticity of the trade balance is expected to be negative while the foreign income elasticity is expected to be positive. Conversely, if income changes reflect supply side effects the signs may be reversed as in the perfect substitutes version of the model.

A potential problem in some recent studies is the choice of variables to represent the trade balance and relative prices. Two different real measures of the trade balance are considered in the present study. The first one is the nominal trade bal-

ance deflated by the U.S. Wholesale or Producer Price Index. Following common practice the nominal trade balance is measured as the difference between bilateral domestic exports and imports.<sup>8</sup> Ideally, under the imperfect substitutes assumption, separate import and export price indexes should be used to compute trade volumes. However, very few countries have reliable import-export price data.<sup>9</sup> For example, there is good information about Japan and Germany's export prices, although in the U.S. good quality data has been available only since the mid-eighties (BLS series). Wholesale prices, on the other hand, are collected universally with much more accuracy. This index reflects more precisely changes in finished as well as intermediate industrial national products. Consequently it constitutes a better deflator than either, GNP deflators or the CPI. In fact, for those developed countries where both export and wholesale price data exists, differences between the two tend to be small.<sup>10</sup> Finally, whether one deflates imports with the local wholesale price index or with the foreign wholesale price index (after appropriate conversion of import values to foreign currency units) makes no difference for the results. The second definition of the trade balance volume is the logarithm of the export-import ratio. Following Clarida (1991), the equation regressing log imports against log income and log relative prices can be interpreted as a structural import demand equation. Therefore, subtracting the import demand equations for the two countries leads to a trade balance equation in which the dependent variable is our variable of choice. Furthermore, this specification of the trade gap has the advantage of allowing for the estimation of long-run income and price elasticities as well as a unit free measure of the trade balance speed of adjustment. Wholesale prices are also used in the computation of the real exchange rate. For future reference, the real exchange rate is defined as  $R = SP^*/P$ , where  $S$  is the bilateral nominal exchange rate and  $P^*/P$  is the ratio of the foreign to domestic wholesale price indices, hereafter referred to as the relative price index.

Equation (1) is a static representation of the trade balance. In the presence of market imperfections and/or adjustment costs such as transportation costs, inventory costs, and information and contractual cost typical in international business transactions, the trade balance is likely to adjust slowly to relative price and income changes.<sup>11</sup> For example, the adjustment process in the trade balance following a significant exchange rate change, the J-curve effect, has received substantial attention in the literature and has been referred to as a potential explanation for the persistence of the U.S. trade deficit. Consequently, the behavior of the trade balance is best captured in a dynamic setting by incorporating lagged values of all the relevant variables into equation (1). Under the assumption that the explanatory variables can be represented by a finite autoregressive process, the joint dynamic behavior of the system will be appropriately captured by a VAR process.

### III.- ECONOMETRIC METHODOLOGY

Let  $H_t$  be a  $(k \times 1)$  vector of  $I(1)$  variables whose dynamic behavior is captured by the following autoregressive model:

$$H_t = \Pi_1 H_{t-1} + \Pi_2 H_{t-2} + \dots + \Pi_p H_{t-p} + \epsilon_t, \quad (2)$$

where the  $\epsilon_t$ 's are  $IIN_k(0, \Sigma)$ . The previous system can be rewritten in first differenced form as:

$$\Delta H_t = \Pi \Delta H_{t-1} + \Gamma_1 \Delta H_{t-1} + \Gamma_2 \Delta H_{t-2} + \dots + \Gamma_{p-1} \Delta H_{t-p+1} + \epsilon_t, \quad (3)$$

where

$$\Gamma_i = -(\Pi_{i+1} + \Pi_{i+2} + \dots + \Pi_p), \quad (i = 1, \dots, p-1),$$

and

$$\Pi = \Pi_1 + \Pi_2 + \dots + \Pi_p - I.$$

The variables in  $\mathbf{H}$  are said to be cointegrated if there is a linear combination of them,  $\beta\mathbf{H}$ , that is  $I(0)$ .  $\beta$  is known as the cointegrating vector. If the system  $\mathbf{H}$  is cointegrated, then the  $\text{rank}(\mathbf{\Pi})=q < k$ , and there exist  $(k \times q)$  matrices  $\alpha$  and  $\beta$  such that  $\mathbf{\Pi} = \alpha\beta'$ .  $q$  is the dimension of the space of cointegrating vectors  $\beta$ , and the  $\alpha$ 's are the vectors of adjustment coefficients. In most cases  $q$  will be 1 and there is a unique cointegrating vector representing a stable long run relationship between the variables in  $\mathbf{H}$ . Johansen (1988) has developed a maximum likelihood procedure based on canonical correlation theory to estimate these coefficients. He also provides the mathematical foundation to conduct statistical tests in this nonstationary environment. The estimation method is as follows:

- 1.- Estimate (3) by OLS to determine  $p$  using a multivariate lag selection criteria.
- 2.- Regress  $\Delta\mathbf{H}_t$  on  $\Delta\mathbf{H}_{t-1}, \Delta\mathbf{H}_{t-2}, \dots, \Delta\mathbf{H}_{t-p+1}$  and save the residuals  $\mathbf{R}_{0t}$ .
- 3.- Regress  $\mathbf{H}_{t-1}$  on  $\Delta\mathbf{H}_{t-1}, \Delta\mathbf{H}_{t-2}, \dots, \Delta\mathbf{H}_{t-p+1}$  and save the residuals  $\mathbf{R}_{kt}$ .
- 4.- The estimate of  $\beta$  is the vector corresponding to the first canonical variate of  $\mathbf{R}_{kt}$  with respect to  $\mathbf{R}_{0t}$ .

The estimation procedure is simple to implement. First, one solves the equation

$$[\lambda\mathbf{S}_{kk} - \mathbf{S}_{k0}\mathbf{S}_{00}^{-1}\mathbf{S}_{0k}] = 0,$$

where  $\mathbf{S}_{ij} = \mathbf{T}^{-1} \sum_{t=1}^T \mathbf{R}_{it}\mathbf{R}_{jt}'$ ,  $(i, j = 0, k)$ .<sup>12</sup> The estimate of  $\beta$  is the eigenvector  $\mathbf{e}$  associated with the largest eigenvalue, appropriately normalized so that  $\mathbf{e}'\mathbf{S}_{kk}\mathbf{e} = 1$ . The adjustment coefficient is then estimated as  $\alpha = \mathbf{S}_{0k}\beta$ . The ordered eigenvalues  $\lambda_1 > \dots > \lambda_k$  play an important role in the testing for cointegration as well as in the testing of parameter restrictions. Two statistics are available to test for cointegration. They both help determine the rank of  $\mathbf{\Pi}$ . The first statistic (MAXEIG) tests the unconditional significance of individual eigenvalues. The second statistic (TRACE) tests the conditional significance of the ordered eigenvalues, for example,  $\lambda_n > 0$  given  $\lambda_{n+1} = \lambda_{n+2} = \dots = \lambda_k = 0$ .

Asymptotic critical values for these statistics are available in Osterwald-Lenum (1990); however, because of the small sample sizes used in this empirical study and the potential size distortions in the cointegration statistics arising from the lag length choice in the ECM, appropriately adjusted critical values have been computed. The relevant adjusted critical values are presented in tables 7 and 8. They were obtained from 5000 replications under the null of no cointegration. I used samples with 75, 50 and 30 observations and the EC models were estimated with 0, 1 and 4 lags. Restrictions on the  $\alpha$ 's and  $\beta$ 's can also be easily tested comparing the eigenvalues from the constrained and unconstrained models through a likelihood ratio test.<sup>13</sup> Furthermore, under the null of cointegration these statistics are distributed asymptotically as a  $\chi^2$ .

It is common practice in many studies to deal summarily with the dynamic specification of the error correction model. There are several reasons why a correct specification of the ECM is important. First, if the ECM lag length is underestimated the parameter estimates will be inconsistent and the behavior of the cointegration statistics will be unreliable. Second, a parsimonious representation of the dynamic behavior of the system may reduce potential multicollinearity problems that usually diminish the significance of the error correction term  $\beta\mathbf{H}_{t-1}$ . Finally, overparametrized EC models, such as those intended to capture seasonal components, may produce severe size distortions in the cointegration statistics in small samples. For example, in a simulation exercise using a sample size of 75 observations, roughly the typical sample size in this empirical analysis, the critical values associated with the 95% quantiles in the asymptotic distributions tabulated in the literature, e.g. Osterwald-Lenum (1990), induced a rejection of the null of no cointegration more than 20% of the time when the ECM was estimated with four lags. The size distortion increased with the dimension of the system and with smaller sample sizes. To address these concerns I have used a seasonal

dummy rather than a seasonal lag and the critical values used in the paper are sample size and ECM lag length adjusted. Finally, the lag specification is further checked using a statistic proposed by Hosking (1980) (HSK). The Hosking statistic is a multivariate extension of the better known Ljung-Box statistic applied to univariate time series. It is distributed asymptotically as a  $\chi^2$ , with the degrees of freedom being determined by the dimension of the system,  $k$ , the number of lags estimated in the error correction model,  $p$ , as well as the number of lagged correlation matrices used to compute the statistic (8 in the paper).

While the theoretical foundations for the analysis of structural breaks in stationary environments have long been established, the theoretical foundations required to analyze structural breaks in non-stationary environments are still in their early stages. Just recently, the *Journal of Business & Economic Statistics* devoted its entire July 1992 issue to this topic. I follow Hansen's (1992) article in that volume to test for structural breaks in a cointegrated relationship. Hansen's method also provides an estimate of the potential break point. Hansen derives the asymptotic distribution of four different Lagrange multiplier tests for parameter instability. The first statistic,  $F_{\tau t}$ , tests the null of no structural breaks against the alternative of a known break point in the same spirit as the traditional Chow test. The second statistic,  $F_{sup}$ , is the maximum of the  $F_{\tau t}$  statistics over all possible break points ( $\tau \in (0, 1)$ ). In practice the range of potential break points is limited to a subset, e.g. [.15,.85]. Therefore, the  $F_{sup}$  statistic is designed to have power against the alternative of a unknown single break point. The other two statistics,  $F_{mean}$  (the sample mean of the  $F_{\tau t}$  values over all possible break points) and  $L_c$ , are designed to have power against the alternative that one or several parameters in the cointegrating vector follow a martingale process. This alternative more accurately captures parameter changes that take place over time rather than instantaneous changes. The interested reader is directed to Hansen's paper for

a detailed derivation of the statistics and a sample of their tabulated asymptotic distributions. As with any novel econometric technique, the finite sample performance of the statistics is still limited and the results should be interpreted with caution. Gregory and Nason (1991), for example, have studied their size and power in a cointegrated linear quadratic adjustment model. The simulation experiments indicate a reasonable performance.<sup>14</sup>

#### IV.- EMPIRICAL ANALYSIS

##### IV.1- Data Sources and Description

Since the focus of the paper is on long run relationship I have used quarterly data. The sample extends from 1973:1 to 1991:4 for a total of 76 observations. Nominal U.S. imports and exports were obtained from the Citibase data tapes. The nominal exchange rate is the end of the first month observation in each quarter, taken from the International Financial Statistics (IFS).<sup>15</sup> Wholesale prices and real GDP figures were also obtained from the IFS. The price variable for France is the GDP deflator since no comprehensive measure of wholesale prices is available for the complete sample.

##### IV.2- General Methodology

A unified approach is used in the analysis of all six bilateral models. First, I estimate the benchmark trade balance equation (1) that includes the natural logarithms of the real exchange rate (RXRT), the domestic real income (DRY) and the foreign real income (FRY) as explanatory variables. The U.S is the domestic country. However, in only two of the bilateral models are the variables in the benchmark specification cointegrated and with the proper parameter signs. The lack of cointegration appears to contradict clear graphical evidence of a strong

correlation between the trade balance and some of the individual explanatory variables, especially the nominal exchange rate. The graphical evidence suggests that the assumption of zero degree homogeneity of the trade balance with respect to the individual components of the real exchange rate may not be appropriate. In addition, there is evidence in the benchmark regressions that the domestic and foreign income trade balance elasticities are either statistically insignificant, e.g. the variables do not belong in the cointegrating relationship, or they have approximately the same value. Consequently, alternative specifications of the trade balance equation were estimated. The alternative specifications are obtained from the benchmark model by a) decomposing the real exchange rate into two separate variables, the nominal exchange rate (NXRT) and the relative price index (DWP/FWP), b) combining the domestic and foreign income levels into a single variable, the relative income level (FRY/DRY), whenever the equality between domestic and foreign income elasticities cannot be rejected and, c) dropping from the trade balance equation all the variables that are not fundamental to the cointegrated relationship. The objective is to derive the most parsimonious representation of a stable long run trade balance equation. This specification will be referred to as the preferred trade balance equation. There are several reasons why a reduction in the dimensionality of the EC model may be desirable. First, for a given sample size, the power of the cointegration test will decrease with the dimension of the system. For example, if both income levels belong in the long run trade balance equation and the income elasticities are equal, the rejection of the null of no cointegration will be stronger (higher significance level) in the reduced system that includes the relative income levels than in the system that includes the income levels separately. Second, even though asymptotically the addition of irrelevant (nonstationary) variables to a cointegrated system will not affect the performance of the cointegration tests, as these redundant variables will be assigned a zero weight in the cointegrating vector, in finite samples the power of the

tests may be seriously reduced due to the nonstationary noise of the redundant variables.

To find the appropriate specification for the long run trade balance equation I follow an incremental approach. In the first step, the potential cointegrating relationship between the trade balance and the real exchange rate is tested. If cointegration is found and the cointegrating vector has the appropriate signs the search stops as we have found a specification that explains the long run behavior of the trade balance that is consistent with the theory. Otherwise, we continue the specification search by decomposing the real exchange rate into the nominal exchange rate and the price index ratio and examining their individual and joint long run effects on the trade balance. In the potential case that both variables are separately cointegrated with the trade balance the choice between the two is determined by the following criteria.<sup>16</sup> First, choose the variable that is exogenous to the system (a positive or an insignificant adjustment coefficient). If both variables are either exogenous or endogenous choose the one that provides the best fit as measured by an  $R^2$ .<sup>17</sup> If no reasonable cointegrating vector is found then real incomes are added under two alternative settings. First, under the assumption that income elasticities have the same magnitude. Second, under the assumption that they are different. In the latter case the equality restriction on the elasticities is formally tested via a likelihood ratio test, *LRT*.

Because of their similar cointegrating behavior only results for the second version of the model are presented. The specification of the trade balance in the second version offers a more straightforward interpretation of the parameter estimates as long run elasticities. All the bilateral models are well characterized by an ECM of order 1 with the addition of seasonal dummies. The specification is further checked with a Hosking statistic (HSK) of order 8. Because of the complex-

ities in computing the degrees of freedom only the percentile of the appropriate  $\chi^2$  distribution is presented. As a measure of goodness of fit an  $R^2$  is also provided. Note that since Johansen's approach does not require an ex-ante normalization of the cointegrating vector to estimate the ECM the  $R^2$  can be negative. This is usually an indication that the variable either does not belong or is exogenous in the cointegrating relationship. Tables 1-6, provide estimates for the cointegrating vector (normalized by the trade balance coefficient) and the coefficients of adjustment,  $\alpha$ . The adjustment coefficients are normalized as to capture the per period percentage adjustment in each variable to the deviation from its long run value. The coefficient should be negative. A positive or insignificant adjustment coefficient indicates that the variable is weakly exogenous in the context of the trade balance relationship. The tables also present the values for Johansen's MAXEIG and TRACE statistics to test for cointegration. An indication is made when more than one cointegrating vector is found. In addition, various likelihood ratio type statistics are computed to test for zero restrictions on the individual alphas,  $LRT_\alpha$ , and betas,  $LRT_\beta$ . A zero value for a particular coefficient in the cointegrating vector indicates that the corresponding variable does not belong in the cointegrating relationship. A zero value for a particular adjustment coefficient indicates that the corresponding variable is not responsive to the temporary deviations from the equilibrium relationship. Under the null that the restriction is true and that a cointegrating relation exists between all or some of the variables in the system the statistics will be distributed asymptotically as  $\chi^2_{(1)}$ .

Following the cointegration analysis of the six bilateral models, I study the possible presence of structural breaks in the trade balance relationship using the approach developed by Hansen (1992). The statistics are computed for two different specifications: the traditional trade balance equation (1) and the preferred equation selected in the specification search. The information provided includes

the values of the  $F_{sup}$ , the  $F_{mean}$  and the  $L_c$  statistics and the break point estimate from  $F_{sup}$ . I also provide a plot of the  $F$  values for all possible break points in the [.15,.85] trimmed sample. The break point estimate is used to divide the full sample into two subsamples. The cointegration analysis is repeated for each subsample, provided they contain 30 or more observations.

### IV.3- Bilateral Results

#### Canada

Table 1 presents the results for the Canadian bilateral model. The benchmark trade balance equation (1) shows no evidence of cointegration. The TRACE and MAXEIG statistics are respectively 32.27 and 18.71, both of which are below the 50% significance level. Furthermore, the  $R^2$  for equation (1) is negative. Consequently, the analysis proceeds according to the incremental approach suggested earlier. The trade balance and the real exchange rate also fail to be cointegrated. However, both the nominal exchange rate and the relative price index are separately cointegrated with the trade balance although the relative price coefficient has the wrong sign. This simultaneous cointegration result is not surprising as the the nominal exchange rate and the relative price index are themselves cointegrated at approximately the 95% level of significance with the nominal exchange rate appearing to be exogenous in that relationship. In addition, the  $R^2$  in the model with the nominal rate is .45, significantly larger than in the model with the relative price index where it is only .09. Consequently, the nominal exchange rate is the variable incorporated into the trade balance equation. The introduction of the real incomes into the equation brings the significance level of the cointegration statistics below the 90 percentile partly due to the loss of power associated with the increase in the dimensionality of the system. In that equation, the equality

restriction between the domestic and income elasticities is rejected. Furthermore, when the restriction is imposed the coefficient in the cointegrating vector associated with the relative income variable is not statistically significant. Therefore, the preferred trade balance equation includes only the nominal exchange rate as an explanatory variable.

The value of the TRACE statistic in the preferred model is 18.04, very close to the 95% adjusted critical value of 18.67. The MAXEIG statistic is 16.57, which is larger than the 95% adjusted critical value of 15.71. The exchange rate elasticity of the trade balance is positive at 0.96 and statistically different from zero. The  $R^2$  for the trade balance equation is .45 and the Hosking statistic (HSK) corresponds to the 40<sup>th</sup> percentile of the appropriate  $\chi^2$  distribution. The most surprising result is that both adjustment coefficients are negative and statistically significant, indicating that both the trade balance and the nominal exchange rate adjust to the temporary deviations from the long run relationship. The speed of adjustment in the exchange rate is quite low with only 9.58% of the deviation adjusted each quarter. The trade balance seems more responsive in its adjustment but still a mere 17.21% of the gap is adjusted each quarter. The results are consistent with the observed lagged response of the trade balance to exogenous exchange rate changes.<sup>18</sup>

#### France

Table 2 presents the cointegration results for the U.S.-France bilateral model. The TRACE and MAXEIG statistics for the benchmark equation are 63.64 and 34.42 respectively, both well above the 95% adjusted critical value. The elasticities are all significant and their signs are consistent with the predictions of the imperfect substitutes model. The domestic income elasticity is larger than that of the foreign country. Furthermore, the likelihood ratio test  $LRT_\gamma$  strongly rejects the

equality of the two coefficients. The  $R^2$  for the estimated long run relationship is .69 and the Hosking statistic, HSK, is at the 46<sup>th</sup> percentile of the corresponding  $\chi^2$  distribution. Even though the results are satisfactory it is still necessary to test the zero homogeneity of the real exchange rate components. When the real exchange rate is decomposed into the nominal rate and the relative price index the strong signs of cointegration are still present; however, two new features appear. First, the relative price variable can be excluded from the equation since the  $LRT_\beta$  is not statistically significant. Second, the equality between the foreign and domestic income elasticities can no longer be rejected. Therefore, the preferred specification includes the nominal exchange rate and the relative income levels as explanatory variables. The TRACE and MAXEIG statistics in the preferred model are 49.90 and 28.76, well above the 99% adjusted critical values of 40.33 and 28.03.<sup>19</sup> The exchange rate elasticity is .67 and the relative income elasticity is 3.62, both values significantly different from zero. Despite the reduction in the dimension of the system the  $R^2$  increases to .74 while the Hosking statistic, HSK, is now only at the 22<sup>nd</sup> percentile. The trade balance is the only apparent endogenous variable in the system with a highly significant adjustment coefficient of .57; that is, more than half of the deviation from the equilibrium trade value is adjusted each quarter.

#### Germany

The results for the U.S.-Germany relationship are presented in Table 3. The benchmark model shows no signs of cointegration; nevertheless, this specification has a good deal of explanatory power with an  $R^2$  of .66. When the specification incremental approach is implemented neither the real exchange rate nor its separate components seem to exhibit any explanatory power in the trade balance equation as every  $R^2$  is negative. That leaves only the two real incomes as po-

tential explanatory variables. The significance level of the cointegration statistics increases when the trade balance equation only includes the income variables. Furthermore, the equality of the income elasticities cannot be rejected as in the benchmark model. Therefore, the preferred specification has as sole explanatory variable the relative income level. The TRACE and MAXEIG statistics in the preferred model, 15.01 and 13.07 respectively, are slightly below the 90% adjusted critical value. The  $R^2$  is still high at .59 and the Hosking statistic is at the 54<sup>th</sup> percentile.<sup>20</sup> The relative income elasticity, 7.81, is large but consistent with the imperfect substitutes model. An interesting feature is that while the equation has substantial explanatory power the trade balance seems to be exogenous or has a very small speed of adjustment. The coefficient of adjustment is negative at -.095, but is not statistically significant. Only relative incomes seem to respond to exogenous trade balance shocks, in particular, the German income variable seems to be the most sensitive of the two.

#### Italy

The results for this bilateral model, in Table 4, are very similar to those in the US-France relationship. The TRACE and MAXEIG statistics for the benchmark model are 54.10 and 28.94 respectively. The first statistic is above the 95% adjusted critical value while the second is significant at the 90% level. The elasticities are all significant and their signs are consistent with the imperfect substitutes model. The domestic income elasticity is larger than that of the foreign country. Furthermore, the  $LRT$  statistic strongly rejects the equality of these coefficients. The  $R^2$  for the estimated long run relationship is .73 and the Hosking statistic is at the 88<sup>th</sup> percentile. As in the US-France model it is the nominal exchange rate variable that is able to better explain the long run behavior of the trade balance. When the real rate is replaced by the nominal rate the equality between the foreign and domestic

income elasticities can no longer be rejected. Therefore, the preferred specification includes the nominal exchange rate and the relative income levels as explanatory variables. The selected model retains its strong support for cointegration with TRACE and MAXEIG statistic values of 49.43 and 28.29 well above the 99% adjusted critical values.<sup>21</sup> The exchange rate elasticity is 1.84 and the relative income elasticity is 10.78 both significantly different from zero. The  $R^2$  increases to .77 while the Hosking statistic is now only at the 97<sup>th</sup> percentile.<sup>22</sup> The trade balance is the only apparent endogenous variable in the system with a highly significant adjustment coefficient of .44; that is, close to a half of the deviation from the equilibrium trade balance level is adjusted each quarter.

#### Japan

The cointegration results for the Japanese model are presented in Table 5. The variables in the benchmark equation are not cointegrated. Furthermore, the  $R^2$  in this model is negative. The trade balance is not cointegrated with either the nominal or the real exchange rate. In addition, the exchange rate elasticity has the wrong sign.<sup>23</sup> Only when the real exchange rate is decomposed into the nominal rate and the relative price do the elasticity signs become consistent with the theory. However, the system still shows no signs of cointegration. Finally, the relative income level is added to the former specification. The TRACE and MAXEIG statistics for the model that includes the nominal exchange rate, the relative price index and the relative income level are 50.27 and 27.15. The first value is larger than the 90% adjusted critical value while the second is only significant at the 80% level. The elasticity estimates are all statistically significant, have reasonable values and the proper signs. The income elasticity is 5.25 and the exchange rate elasticity is 1.44. Furthermore, the  $R^2$  is .47 and the Hosking statistic corresponds to the 17<sup>th</sup> percentile of the appropriate  $\chi^2$  distribution. A striking feature of the

model is the seemingly exogenous behavior of the trade balance. The coefficient of adjustment for the trade balance is -.06, negative but statistically insignificant, indicating a very slow trade balance adjustment. Only the relative price index seems to be responsive to the deviations from its long run values.

#### United Kingdom

Table 6 presents the results for the U.S.-U.K. bilateral model. The elasticity estimates in the benchmark model are satisfactory and have the proper signs. The explanatory power is moderate with an  $R^2$  of .29. However, the TRACE and MAXEIG statistics fail to reject the null of no cointegration. Of all the trade balance specifications attempted the bivariate model involving the trade balance and the nominal rate is the only one where the cointegration results are marginally satisfactory.<sup>24</sup> The TRACE and MAXEIG statistics for the preferred trade balance equation are 15.51 and 13.68 respectively. The first value is not far from 16.30, the 90% adjusted critical value, while the second is significant at that percentile. The exchange rate elasticity is 1.10, is statistically significant and has the appropriate sign. The  $R^2$  is moderate at .19 and the Hosking statistic validates the chosen ECM specification at the 53<sup>rd</sup> percentile. Both the trade balance and the exchange rate appear to be endogenous with significant adjustment coefficients. The trade balance adjustment coefficient is .19 and the exchange rate adjustment coefficient is .14. Therefore, trade balance and exchange rate adjustments are relatively slow.

#### Summary of Results

The previous analysis indicates that except for France and Italy the benchmark equation is not an appropriate characterization of the long run behavior of the trade balance behavior since the real trade balance is not cointegrated with the real exchange rate and the domestic and foreign real incomes. However, when

the real exchange rate is divided into its two integral components, the nominal exchange rate and the relative price index, then the trade balance is cointegrated with a subset of the explanatory variables in all six instances. Despite the relative low power of the tests the null of no cointegration is rejected at the 5% level of significance in three cases (Canada, France and Italy) and is rejected at the 10%-15% level of significance in the remaining bilateral models. The estimated long run elasticities in the preferred specification have signs that are always consistent with the imperfect substitutes model.

Except for the bilateral model with Germany, the exchange rate always contributes significantly to the long run movements of the trade balance. Furthermore, it is not the real but the nominal exchange rate that best explains the long run behavior. The exchange rate elasticities are always positive and their values oscillate between 0.67 and 1.44, the latter value corresponding to the Japanese bilateral model. In four of the six preferred trade balance specifications (France, Germany, Italy and Japan) the income levels are part of the equation. In fact, in all four models the equality of the domestic and foreign income elasticities is never rejected. The income elasticity values oscillate between 3.6 and 7.8, and are generally above previously reported elasticities values.

The traditional assumption that trade balances respond slowly to exogenous shocks is confirmed. In all the bilateral models the trade balance coefficient of adjustment ( $\alpha$ ) is negative. For two of the models, Germany and Japan, the coefficient is so small that it is statistically insignificant. In the Canada and U.K. models the speed of adjustment is less than 20% per quarter. However, in the bilateral models with France and Italy the speed of adjustment increases to a 50% quarterly rate. Note that these two models also exhibit the strongest cointegration signs. The nominal exchange rate appears to be adjusting to deviations from its

long run value in two cases only, Canada and U.K., but at a fairly slow pace. Relative income levels seem to exhibit an exogenous behavior in the trade balance equation, with the exception of the German model where the German income level appears to be the variable responsible for that adjustment.

#### IV.4- Analysis of Structural Breaks

Table 7 presents the results from the analysis of structural breaks. The statistics  $F_{sup}$ ,  $F_{mean}$ ,  $L_c$ , and the break point estimate are computed for both the benchmark and the preferred specification of the trade balance equation. The asymptotic critical values tabulated in Hansen (1992) are also presented at the bottom of the table. In addition, Figures 7-9 present the plots of the F-values corresponding to each possible break point in the [.15,.85] trimmed sample for all six bilateral models.

The different tests seem to find little evidence of structural breaks in the U.S.-Canada data. Even though the values are well below the appropriate significance levels they appear to indicate a potential change around 1982. Consequently, the analysis was performed for both the pre and post 82:1 subsamples. The explanatory power of the trade balance equation is greater in the second subsample where in addition to the nominal exchange rate, income levels form part of the selected specification. A peculiar feature of the bilateral model with Canada was the endogeneity of the exchange rate in the trade balance equation. To investigate the stability of this feature the preferred trade balance equation was estimated using a rolling sample of 8 years. While the coefficient of adjustment for the exchange rate changes throughout the sample, it is always significantly different from zero. The consistency in the endogenous behavior of the exchange rate throughout the sample may be a reflection of the sustained management of the exchange rate on

the part of the Central Bank of Canada.

The  $F_{mean}$  and  $L_c$  statistics in the French bilateral model suggest the presence of some form of change. The estimated break point is somewhere in the 86-87 period. However, since the  $F_{sup}$  statistic is below the appropriate critical value, the change appears to be progressive rather than drastic. In fact, the plot of F-values suggests 1981 as the possible beginning of the change. Several equations were estimated for the pre and post 81:1 subsamples. As in the previous case, the explanatory power of the trade balance equation is greater in the second subsample where, unlike the full sample results, the equality restriction between income elasticities can now be rejected. The foreign income elasticity is slightly larger than the domestic income elasticity, both values increasing over the full sample estimates. The exchange rate elasticity, on the contrary, decreases slightly.

In the German bilateral model all the statistics fail to reject the null of no structural breaks. The break point estimate is 88:1 which may be capturing the potential effects of the German unification in 1989. When the last three years of the sample are dropped results barely change from the full sample case. However, unlike the full sample results the real exchange rate is no longer irrelevant in the explanation of the long run behavior of the trade balance. The real exchange rate elasticity in the 73:1-88:1 subsample is .519 and the income elasticity is 7.840. The TRACE and MAXEIG statistics are 39.003 and 24.463 respectively, both significant at the 5% percent level. The  $R^2$  of the regression is .762. Both the trade balance and the relative incomes are endogenous with a speed of adjustment coefficient of .329 and .306 respectively. There is no evidence of structural breaks in the U.S.-Italy model. The plot with the F-values suggests the presence of two potential break points, 81:4 and 86:2. The model was reestimated for two overlapping subsamples, 73:4-86:2 and 81:4-91:4. The estimates do not differ

across subsamples and are similar to the full sample results. The significance of the cointegration statistics decreases marginally, leaving only one cointegration vector.

The analysis of structural breaks shows a very significant rejection of the null in the Japanese model. However, this result has to be interpreted with caution as the estimated break point is at the end of the sample. The preferred trade balance equation was reestimated for the 73:1-88:4 sample. The elasticities and the cointegration statistics increase in value. The TRACE and MAXEIG statistics are able to reject the null of cointegration at the 10% and 5% significance levels respectively. However, the  $R^2$  falls to .323. The analysis failed to produce any substantive result in the British model. However, unlike the Canadian case, the endogeneity of the exchange rate unveiled in the full sample results seem to be more the influence of the first part of the sample. The rolling regressions show an exogenous exchange rate behavior after 1985.

## V.- CONCLUSIONS

The persistence of the U.S. trade balance deficit in spite of the drastic dollar devaluation after 1985 has generated a lively debate over its causes and cures. The persistence of the deficit has led numerous authors to argue that the traditional determinants of trade balance behavior, exchange rates and incomes, are no longer sufficient to explain recent events. Authors such as Krugman (1991) and Lawrence (1990) maintain instead that the adjustment process embodied in traditional trade balance models has worked. However, there is little empirical evidence to support that claim. Recent work by Yellen and Rose (1989) and Rose (1991) has failed to find a stable long run relationship between the trade balance and its traditional determinants. Furthermore, their results also indicate that the exchange rate plays

no significant role in explaining trade balance behavior.

In the present paper I have reexamined the empirical evidence in support for a stable long run relationship between U.S. bilateral trade balances, exchange rates and incomes during the floating period. I have also examined the relevance of exchange rate fluctuations in explaining the long run behavior of the trade balance and whether or not exchange rates can be considered exogenous in such a relationship. Both issues are fundamental in assessing the impact of exchange rate policies in the management of trade imbalances. The paper differs from previous research in several aspects. First, I use the maximum likelihood approach suggested by Johansen (1988, 1991) to test for cointegration. Johansen's approach also provides estimates of the long run elasticities and the trade balance speed of adjustment. Second, I conduct a specification search to determine an appropriate characterization of the long run behavior of the trade balance. Finally, I use Hansen's (1992) approach to test for parameter stability in a cointegrated relationship and to estimate potential break points. The six bilateral relationships studied in the paper involve the U.S. and its major trade partners: Canada, France, Germany, Italy, Japan and the United Kingdom (the G-7 group).

The empirical results indicate the presence of stable long run trade balance equations in all six models, however, cointegration results are weakest for the U.S.-Japan model. In most bilateral models, the specification that best explains the long run behavior of the trade balance includes the exchange rate and the income levels as explanatory variables. The empirical analysis also shows no significant evidence of structural breaks in the selected trade balance equations. There is some indication of slow changes in the U.S.-France bilateral relationship during the early eighties. In addition, the cointegration results for the German and Japanese bilateral models improve marginally when the last three years of the sample are

dropped. In general, I find that the nominal exchange rate has better long run explanatory power than the real exchange rate. This result is not surprising given the strong correlation between the two rates and the difficulty to obtain appropriate import-export price measures. The long run exchange rate elasticity is in the neighborhood of one, while the income elasticities, whose signs are always consistent with the imperfect substitutes trade model, vary across countries and are usually larger than previously reported estimates. In addition, the equality between the domestic and foreign income elasticities is seldom rejected. The exchange rate and income levels can be generally assumed exogenous in most trade balance equations. Finally, trade balances are found to adjust slowly to exogenous shocks, where the speed of adjustment varies significantly across bilateral models.

The results in this paper suggest that the position advocated by Krugman and Lawrence has a solid empirical foundation. Traditional trade balance equations such as those derived from the imperfect substitutes model of trade are sufficient to explain the long run behavior of the trade balance. The fact that the exchange rate can be considered exogenous with respect to the trade balance and that the Marshall-Lerner condition holds in the long run indicates that economic policies designed to moderate exchange rate fluctuations will be successful in managing external imbalances between industrialized countries. However, the lags in trade balance adjustment require between one and two years before the full effects of a devaluation may be felt. The results also indicate that phenomena such as hysteresis in trade or "pass-through" effects seem to play only a small role in the explanation of the long run behavior of the trade balance as the estimated equations are stable throughout the sample.

This study has concentrated on U.S. bilateral relationships with other industrialized countries. It was argued that the bilateral data is less likely to be

polluted with measurement errors and it avoids the idiosyncrasies of trade in commodities. While trade in manufactures has been considered the "primary interface between the U.S. and the world economy" (Lenz, (1992)), U.S. trade imbalances with developing countries have also contributed to the present deficit. Therefore, in future research I plan to analyze aggregate trade balance behavior. In addition, I also plan to consider the separate analysis of import and export behavior. It is expected that such an analysis will provide a better understanding of the role of supply and demand factors in trade as well as the apparent peculiarities of Japan's trade practices.

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## FOOTNOTES

<sup>1</sup> Even in partial equilibrium approaches such as the portfolio balance model of exchange rate determination or the imperfect substitutes model of trade the exchange rate cannot be taken as exogenous.

<sup>2</sup> Occasionally, detrended series are used in the analysis. However, this method of eliminating the nonstationarity of the series is likely to induce a spurious correlation between the detrended variables.

<sup>3</sup> In addition, as indicated in Phillips and Ouliaris (1990), these two techniques may suffer from a normalization problem as results may depend on the choice of dependent variable in the OLS regression.

<sup>4</sup> See, for example, Deyak et al. (1989) for some evidence on changes in trade balance behavior across exchange rate regimes.

<sup>5</sup> For example, Rose and Yellen (1989) found no additional explanatory power when the effective exchange rate was added to the trade balance equations. These results are consistent with earlier work by Rousslang and Parker (1984) who found very small estimates of cross-price elasticities in import demand equations.

<sup>6</sup> See, for example, Walter (1991) for an analysis of quality issues concerning the U.S. merchandise trade statistics.

<sup>7</sup> See Goldstein and Khan (1985) and Bryant et al. (1988) for comprehensive surveys of the empirical literature.

<sup>8</sup> An alternative definition involves the difference between foreign and domestic bilateral imports. Import data is generally considered of superior quality because of customs collection procedures. However, most countries usually list c.i.f. figures that include insurance and transportation costs. In addition, foreign data on domestic exports is likely to also reflect collection lags. Except for Canada, both measures of the trade balance are very similar. The differences between Canadian import data and U.S. export data may be caused by the large amount of small transactions across the border.

<sup>9</sup> The more commonly available unit value indexes are not satisfactory measures. See, for example, Alterman (1991) and references therein.

<sup>10</sup> For example, during the 1985–1991 period the inflation differential between the two series was a mere 1.7% in the U.S. and 0.1% for Germany. However, in

Japan that difference was a significantly larger 10.5%.

<sup>11</sup> See Gagnon (1989) for a recent discussion of this issue.

<sup>12</sup> The solution is easily obtained with the help of a standard eigenvalue routine by solving the equivalent problem  $|\lambda I - C^{-1}S_{k0}S_{00}^{-1}S_{0k}C^{-1}| = 0$ , with  $C$  being the Cholesky decomposition of  $S_{kk}$ , such that  $CC' = S_{kk}$ .

<sup>13</sup> Consider the restriction  $\beta = H\rho$ , where  $\rho$  is the cointegrating vector under restrictions and  $H$  is the appropriate transformation matrix. The eigenvalues of the system under restrictions can be easily calculated solving  $|\lambda H'S_{kk}H - H'S_{k0}S_{00}^{-1}S_{0k}H| = 0$ . In the presence of only one cointegrating vector, the likelihood ratio statistic for the hypothesis  $\beta = H\rho$  is of the form  $LRT = T \ln \frac{1-\lambda_1^r}{1-\lambda_1}$ , where  $\lambda_1^r$  and  $\lambda_1$  are the restricted and unrestricted largest eigenvalues respectively. A similar but slightly more complicated procedure can be used to test restrictions on the  $\alpha$ 's, such as non-zero restrictions.

<sup>14</sup> I would like to thank Gregory and Nason for providing me with the computer programs used in their paper.

<sup>15</sup> The analysis has been also performed with the quarterly average of daily rates. The parameter estimates are not affected by that choice. However, because of the lower variability in the average measure the significance of the cointegration statistics decreases marginally.

<sup>16</sup> For example, under certain conditions such as purchasing power parity the nominal exchange rate and relative price indexes may be cointegrated. Therefore if one of the variables is cointegrated with the trade balance the other is also likely to be.

<sup>17</sup> If both variables were to be included in the equation more than one cointegrating vector is likely to appear. Inference and testing in a multidimensional cointegrating space may be subject to an identification problem.

<sup>18</sup> An interesting feature in the results for the benchmark model is the sign of both income elasticities, a positive domestic income elasticity and a negative foreign income elasticity, a result more consistent with the perfect substitutes model than with the more traditional of imperfect substitutes.

<sup>19</sup> There are also strong signs of a second cointegrating vector. This may reflect the fact that the trade balance and the nominal exchange rate are independently cointegrated.

<sup>20</sup> A graphical analysis shows the two time series to be very strongly correlated. However, the co-movement seems to be weaker after 1987.

<sup>21</sup> As in the US-France model there are also strong signs of a second cointegrating vector.

<sup>22</sup> Several lag specifications were attempted with no success. However, the residuals from each individual equation are approximately white noise.

<sup>23</sup> This probably reflects the fact that despite the yen appreciation in the last few years the U.S. trade balance with Japan has worsened.

<sup>24</sup> A visual inspection of the time series graphs suggests a strong correlation between the two variables, especially after 1975.

TABLE 1

ECM : U.S.-CANADA  
73:1 - 91:4

Benchmark Trade Balance Equation

Variables : TB, RXRT, FRY, DRY  
 $R^2 = -0.124$ , %HSK = .764

Parameter Estimates

	TB	RXRT	FRY	DRY
$\beta$	1.000	-2.316	+5.537	-6.438
LRT $_{\beta}$	4.999	7.599	7.857	7.452
LRT $_{\gamma}$	4.508			
$\alpha$	-0.149	-0.063	-0.073	-0.093
LRT $_{\alpha}$	6.074	1.614	1.481	1.365

Cointegration Tests

q	4	3	2	1
TRACE	1.632	6.004	14.201	33.178
MAXEIG	1.632	4.371	8.197	18.977

Preferred Trade Balance Equation

Variables : TB, NXRT  
 $R^2 = 0.454$ , %HSK = .402

Parameter Estimates

	TB	NXRT
$\beta$	1.000	-0.963
LRT $_{\beta}$	14.259	12.378
$\alpha$	-0.172	-0.096
LRT $_{\alpha}$	5.157	10.941

Cointegration Tests

q	2	1
TRACE	1.478	18.046
MAXEIG	1.478	16.568

TABLE 2

ECM : U.S.-FRANCE  
73:1 - 91:4Benchmark Trade Balance EquationVariables : TB, RXRT, FRY, DRY  
 $R^2 = 0.688$ , %HSK = .464

## Parameter Estimates

	TB	RXRT	FRY	DRY
$\beta$	1.000	-0.806	-3.958	+4.712
LRT $_{\beta}$	13.312	13.005	6.404	8.139
LRT $_{\gamma}$	10.136			
$\alpha$	-0.456	+0.037	+0.031	-0.167
LRT $_{\alpha}$	9.158	0.386	0.779	6.966

## Cointegration Tests

q	4	3	2	1
TRACE	1.221	8.462	29.222	63.645
MAXEIG	1.221	7.241	20.760	34.423

Preferred Trade Balance EquationVariables : TB, NXRT, FRY/DRY  
 $R^2 = 0.739$ , %HSK = .244

## Parameter Estimates

	TB	NXRT	FRY/DRY
$\beta$	1.000	-0.671	-3.621
LRT $_{\beta}$	12.741	11.954	6.848
$\alpha$	-0.571	+0.035	-0.093
LRT $_{\alpha}$	9.622	0.451	1.813

## Cointegration Tests

q	3	2	1
TRACE	5.345	21.141	49.904
MAXEIG	5.345	15.796	28.763

TABLE 3

ECM : U.S.-GERMANY  
73:1 - 91:4Benchmark Trade Balance EquationVariables : TB, RXRT, FRY, DRY  
 $R^2 = 0.657$ , %HSK = .332

## Parameter Estimates

	TB	RXRT	FRY	DRY
$\beta$	1.000	-0.336	-8.249	+7.970
LRT $_{\beta}$	7.919	0.979	7.378	7.815
LRT $_{\gamma}$	0.500			
$\alpha$	-0.146	+0.016	-0.244	-0.066
LRT $_{\alpha}$	1.568	0.586	5.355	1.036

## Cointegration Tests

q	4	3	2	1
TRACE	0.268	8.174	21.897	44.160
MAXEIG	0.268	7.907	13.722	22.263

Preferred Trade Balance EquationVariables : TB, FRY/DRY  
 $R^2 = 0.586$ , %HSK = .542

## Parameter Estimates

	TB	FRY/DRY
$\beta$	1.000	-7.812
LRT $_{\beta}$	10.249	10.168
$\alpha$	-0.095	-0.217
LRT $_{\alpha}$	1.666	8.173

## Cointegration Tests

q	2	1
TRACE	1.938	15.008
MAXEIG	1.938	13.070

TABLE 4

ECM : U.S.-ITALY  
73:1 - 91:4Benchmark Trade Balance EquationVariables : TB, RXRT, FRY, DRY  
 $R^2 = 0.728$ , %HSK = .877

## Parameter Estimates

	TB	RXRT	FRY	DRY
$\beta$	1.000	-1.270	-3.998	+6.022
LRT $_{\beta}$	10.123	9.618	4.663	6.446
LRT $_{\gamma}$	10.148			
$\alpha$	-0.323	+0.028	-0.054	-0.057
LRT $_{\alpha}$	9.628	0.134	2.460	1.396

## Cointegration Tests

q	4	3	2	1
TRACE	0.824	6.548	25.153	54.096
MAXEIG	0.824	5.724	18.605	28.943

Preferred Trade Balance EquationVariables : TB, NXRT, FRY/DRY  
 $R^2 = 0.768$ , %HSK = .973

## Parameter Estimates

	TB	NXRT	FRY/DRY
$\beta$	1.000	-0.905	-4.282
LRT $_{\beta}$	12.841	11.371	6.507
$\alpha$	-0.438	+0.071	-0.034
LRT $_{\alpha}$	12.847	1.189	0.378

## Cointegration Tests

q	3	2	1
TRACE	5.774	21.136	49.430
MAXEIG	5.774	15.362	28.295

TABLE 5

ECM : U.S.-JAPAN  
73:1 - 91:4Benchmark Trade Balance EquationVariables : TB, RXRT, FRY, DRY  
 $R^2 = -4.227$ , %HSK = .115

## Parameter Estimates

	TB	RXRT	FRY	DRY
$\beta$	1.000	-3.402	+0.846	-2.126
LRT $_{\beta}$	0.364	1.353	0.016	0.023
LRT $_{\gamma}$	0.029			
$\alpha$	-0.053	-0.005	-0.004	-0.001
LRT $_{\alpha}$	3.377	0.011	5.103	0.072

## Cointegration Tests

q	4	3	2	1
TRACE	2.155	10.408	24.066	43.134
MAXEIG	2.155	8.253	13.658	19.068

Preferred Trade Balance EquationVariables : TB, NXRT, DWF/FWP, FRY/DRY  
 $R^2 = 0.465$ , %HSK = .165

## Parameter Estimates

	TB	NXRT	DWF/FWP	FRY/DRY
$\beta$	1.000	-1.443	+5.694	-5.255
LRT $_{\beta}$	12.509	6.250	12.512	9.589
$\alpha$	-0.060	+0.101	-0.283	-0.041
LRT $_{\alpha}$	0.474	3.599	12.835	0.089

## Cointegration Tests

q	4	3	2	1
TRACE	2.638	8.982	23.124	50.270
MAXEIG	2.638	6.344	14.143	27.146

**TABLE 6**

ECM : U.S.-U.K.  
73:1 - 91:4

**Benchmark Trade Balance Equation**  
Variables : TB, RXRT, FRY, DRY  
R<sup>2</sup> = 0.292, %HSK = .671

**Parameter Estimates**

	TB	RXRT	FRY	DRY
$\beta$	1.000	-0.860	-4.905	+4.623
LRT <sub><math>\beta</math></sub>	6.079	2.519	3.862	5.215
LRT <sub><math>\gamma</math></sub>	0.162			
$\alpha$	-0.318	-0.070	-0.008	-0.045
LRT <sub><math>\alpha</math></sub>	6.425	1.325	0.010	1.371

**Cointegration Tests**

q	4	3	2	1
TRACE	0.788	11.513	23.698	42.310
MAXEIG	0.788	10.725	12.185	18.612

**Preferred Trade Balance Equation**

Variables : TB, NXRT  
R<sup>2</sup> = 0.191, %HSK = .531

**Parameter Estimates**

	TB	NXRT
$\beta$	1.000	-1.100
LRT <sub><math>\beta</math></sub>	9.115	10.938
$\alpha$	-0.193	-0.142
LRT <sub><math>\alpha</math></sub>	5.438	8.880

**Cointegration Tests**

q	2	1
TRACE	1.829	15.507
MAXEIG	1.829	13.878

**TABLE 7**

**MAXEIG STATISTIC  
CRITICAL VALUES**

NVAR	NOBS	LAGS	50%	80%	90%	95%	97.5%	99%
2	30	0	7.816	11.336	13.470	15.585	17.370	19.758
		1	8.365	11.983	14.401	16.556	18.594	20.945
		4	11.174	16.201	19.537	22.499	25.222	29.229
		0	7.718	10.927	13.126	15.246	17.235	19.627
	50	1	8.207	11.757	14.066	16.003	18.293	20.447
		4	9.251	13.397	16.016	18.571	21.037	23.707
		0	7.686	11.071	13.117	15.362	17.115	19.504
		1	7.930	11.306	13.544	15.711	17.929	20.806
	75	4	8.535	12.223	14.522	16.835	18.969	21.246
		0	7.527	10.768	12.912	14.900	17.065	19.194
		1	7.497	10.832	12.947	15.050	16.757	19.090
		4	7.694	10.996	12.992	14.869	16.821	18.822
3	30	0	13.586	17.792	20.335	22.797	25.543	28.495
		1	15.384	20.128	22.923	25.562	28.658	32.217
		4	27.298	36.671	42.496	47.686	52.494	60.493
		0	13.177	17.300	19.943	22.362	24.113	27.379
	50	1	14.174	18.739	21.612	24.303	26.435	29.057
		4	17.738	23.160	26.534	29.800	32.589	36.711
		0	13.096	17.124	19.668	21.730	23.713	26.567
		1	13.651	18.021	20.856	23.054	25.217	28.028
	75	4	15.689	20.421	23.542	26.129	28.867	32.808
		0	12.652	16.508	18.904	21.074	22.893	25.752
		1	12.955	16.936	19.275	21.545	23.620	26.135
		4	13.200	17.240	19.714	22.153	24.118	26.509
4	30	0	19.595	24.579	27.451	30.067	32.581	35.723
		1	23.477	29.571	33.241	36.659	39.697	43.150
		4	80.918	109.193	128.875	147.242	168.901	187.675
		0	18.836	23.472	26.173	28.526	30.966	33.632
	50	1	20.765	25.770	28.708	31.530	34.463	37.004
		4	29.682	37.145	41.815	45.871	49.439	53.448
		0	18.611	23.142	25.899	28.354	30.457	33.500
		1	19.752	24.758	27.838	30.428	32.946	35.412
	75	4	24.212	30.520	34.231	37.399	40.112	44.166
		0	17.828	22.160	24.783	27.136	29.155	32.135
		1	18.478	22.867	25.378	27.870	30.147	33.330
		4	19.008	23.580	26.531	29.111	31.515	34.566

NVAR: Number of Variables in the System  
NOBS: Sample Size  
LAGS: Number of Lags in the ECM

TABLE 8

TRACE STATISTIC  
CRITICAL VALUES

NVAR	NOBS	LAGS	50%	80%	90%	95%	97.5%	99%
2	30	0	9.716	13.805	16.316	18.835	21.073	23.770
		1	10.326	14.636	17.373	19.765	22.384	25.438
		4	13.515	19.497	23.505	27.113	29.907	34.479
	50	0	9.569	13.525	15.993	18.182	20.306	23.383
		1	10.140	14.406	16.952	19.373	21.214	24.356
		4	11.330	16.167	19.493	22.344	24.614	27.829
	75	0	9.516	13.506	16.028	18.369	20.744	23.331
		1	9.832	13.828	16.302	18.667	20.991	24.304
		4	10.499	14.834	17.629	20.015	22.524	25.249
	400	0	9.393	13.208	15.663	17.953	20.078	23.517
		1	9.363	13.228	15.690	18.072	20.080	22.997
		4	9.573	13.531	15.895	18.049	20.129	23.074
3	30	0	21.459	27.326	30.847	34.012	36.972	40.191
		1	24.440	30.552	34.334	37.783	41.168	45.869
		4	42.368	55.152	62.478	69.212	76.212	84.017
	50	0	21.045	26.594	29.904	32.881	35.707	39.425
		1	22.354	28.664	32.306	35.395	38.478	41.709
		4	27.850	35.527	40.568	44.387	47.961	52.535
	75	0	20.739	26.394	29.256	32.107	34.579	38.040
		1	21.622	27.532	30.995	33.785	36.442	40.330
		4	24.638	31.121	35.287	38.886	42.296	47.286
	400	0	20.109	25.389	28.709	31.525	34.479	37.216
		1	20.563	25.977	29.102	32.060	34.773	38.027
		4	20.897	26.537	29.501	32.597	35.285	38.703
4	30	0	38.312	45.702	49.957	53.827	57.315	60.871
		1	45.442	54.243	59.641	64.378	68.897	74.928
		4	145.823	184.694	207.786	228.522	246.846	273.499
	50	0	36.685	43.871	47.914	51.178	54.724	58.790
		1	40.279	47.894	52.428	56.441	59.990	63.633
		4	57.387	69.145	75.310	81.107	85.580	92.135
	75	0	36.262	43.229	47.255	51.174	54.212	58.350
		1	38.702	45.790	50.095	54.042	57.592	62.823
		4	46.959	56.282	61.420	66.247	71.293	77.266
	400	0	34.909	41.650	45.229	48.280	51.539	55.432
		1	35.971	42.721	46.751	50.079	53.024	56.878
		4	36.945	44.182	48.380	52.023	54.864	58.168

NVAR: Number of Variables in the System  
NOBS: Sample Size  
LAGS: Number of Lags in the ECM

TABLE 9

TESTS FOR STRUCTURAL BREAKS  
73:1 - 81:4

Benchmark Trade Balance Equation  
(k=4)

	F <sub>sup</sub>	Break Date	F <sub>mean</sub>	L <sub>c</sub>
CANADA	7.674	81 : 3	3.282	0.454
FRANCE	12.742	87 : 3	6.876	0.730
GERMANY	6.037	88 : 2	3.304	0.162
ITALY	9.366	81 : 4	5.692	0.295
JAPAN	40.362	88 : 4	5.059	0.279
U.K.	7.106	79 : 2	3.624	0.300

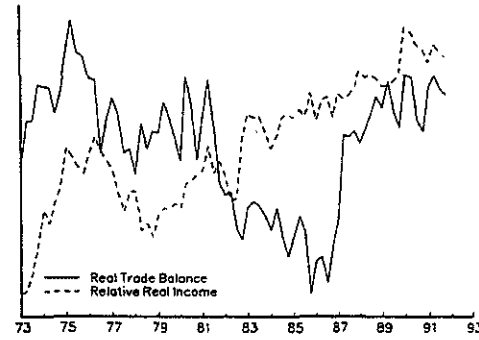
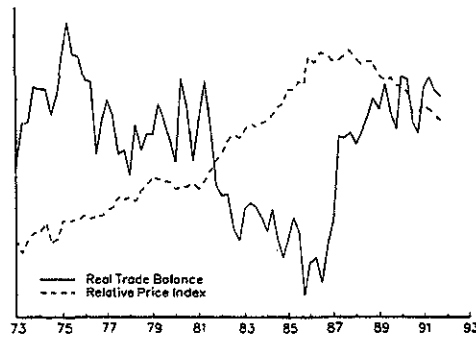
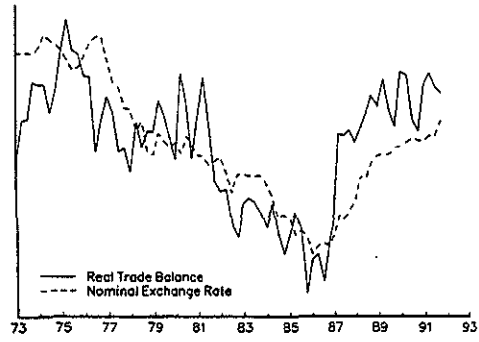
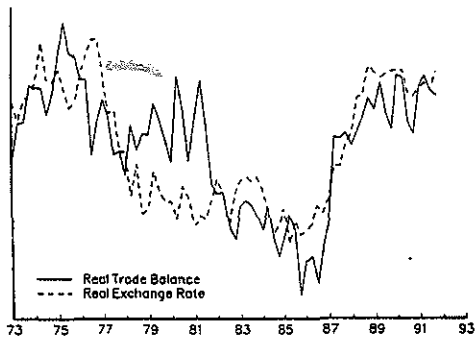
Preferred Trade Balance Equation

	(k)	F <sub>sup</sub>	Break Date	F <sub>mean</sub>	L <sub>c</sub>
CANADA	(2)	7.194	76 : 4	3.547	0.343
FRANCE	(3)	7.291	86 : 1	4.041	0.326
GERMANY	(2)	2.062	87 : 3	0.844	0.095
ITALY	(3)	8.214	86 : 2	4.276	0.418
JAPAN	(4)	10.272	77 : 4	5.127	0.376
U.K.	(2)	5.568	85 : 2	2.018	0.247

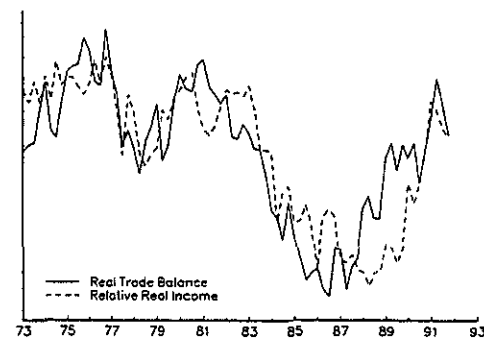
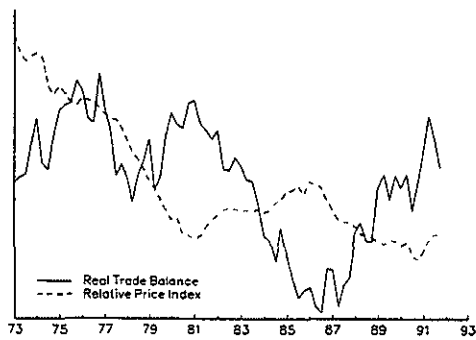
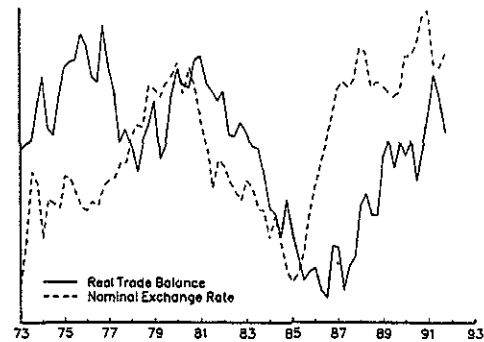
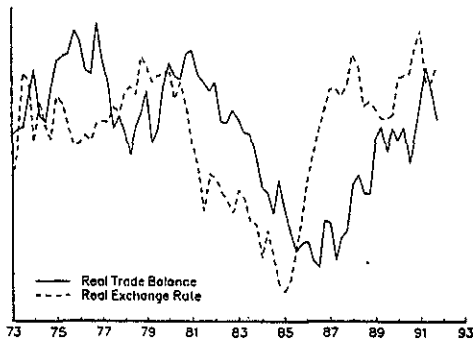
TESTS FOR STRUCTURAL BREAKS  
ASYMPTOTIC CRITICAL VALUES

	F <sub>sup</sub>			F <sub>mean</sub>			L <sub>c</sub>		
	k = 2	k = 3	k = 4	k = 2	k = 3	k = 4	k = 2	k = 3	k = 4
90%	10.6	13.0	15.3	3.73	5.18	6.66	.450	.556	.680
95%	12.4	14.8	17.2	4.57	6.17	7.68	.575	.690	.834

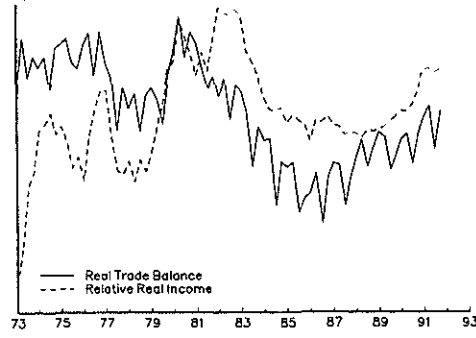
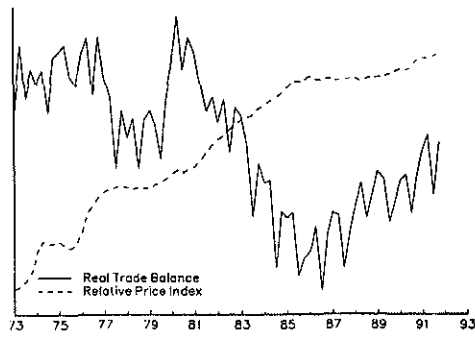
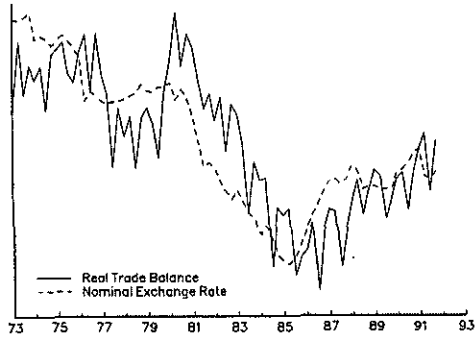
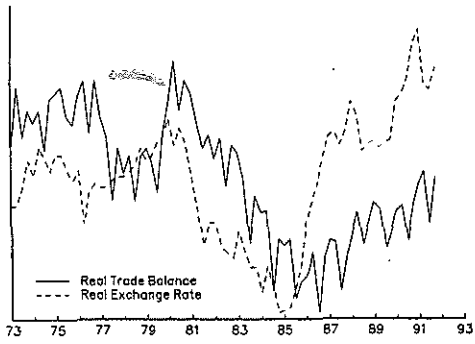
# U.S. - CANADA



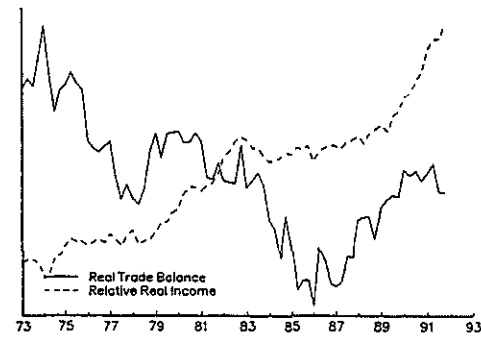
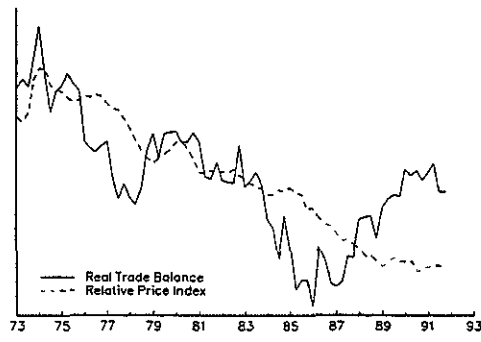
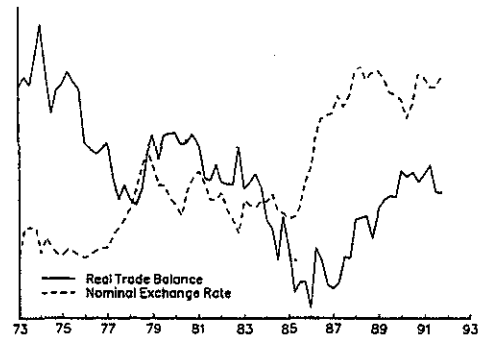
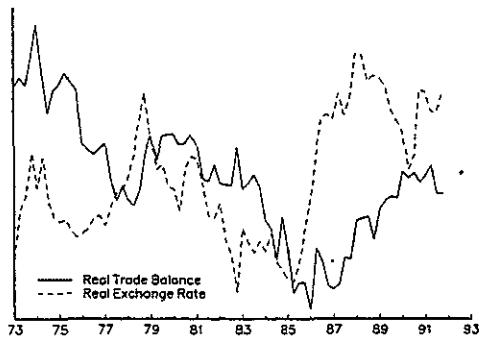
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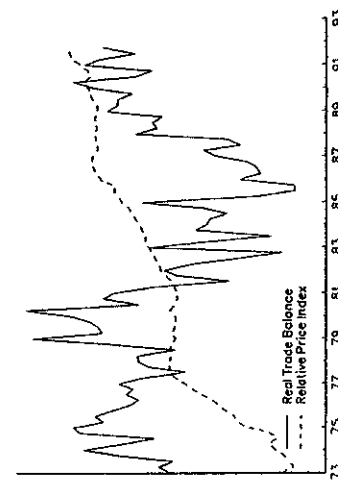
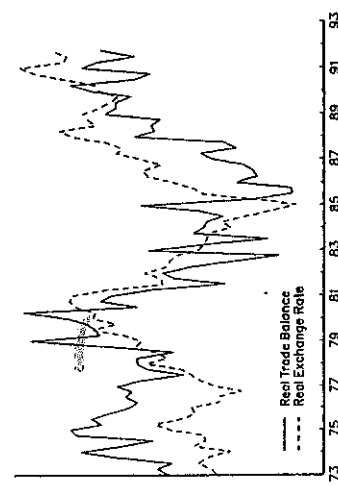
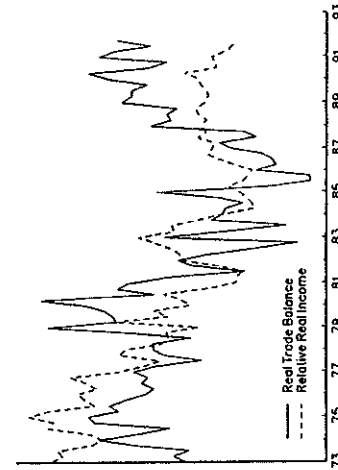
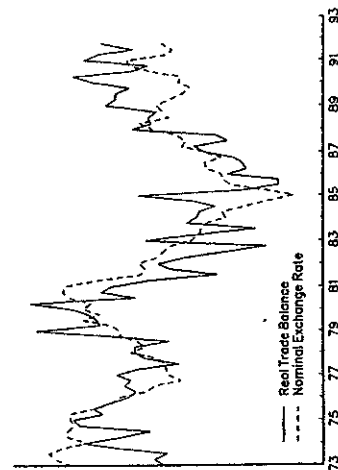
# U.S. - ITALY



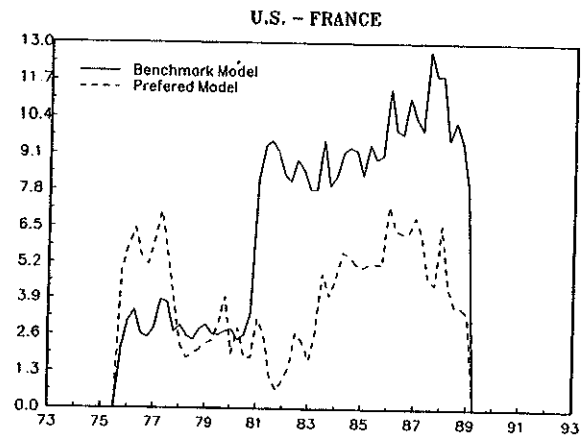
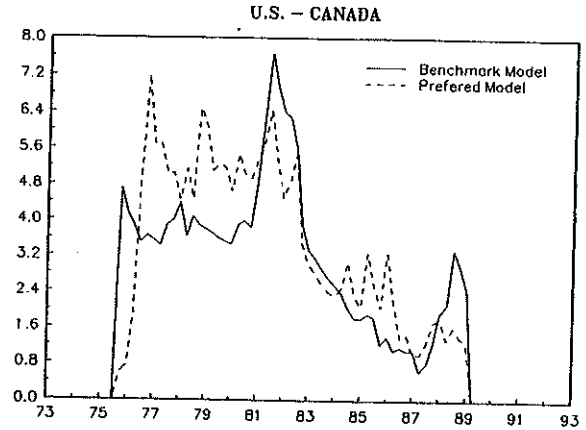
# U.S. - JAPAN



U.S. - U.K.



STRUCTURAL BREAKS  
Hansen F-Test



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