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Bilateral J-Curve between U.S. and Her Trading Partners

By

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I. Introduction

Many studies that have tested the J-Curve phenomenon have employed aggregate trade data. The list includes Bahmani-Oskooee (1985), Felmingham and Divisekera (1986), Felmingham (1988), Rosensweig and Koch (1988), Himarios (1989), Bahmani-Oskooee and Malixi (1992) and Bahmani-Oskooee and Alse (1994). Many of these studies also employed the effective exchange rate. A problem with this approach is that a country's currency could appreciate against one currency and simultaneously depreciate against another currency. The weighted averaging will therefore smooth out the effective exchange rate fluctuations, yielding an insignificant link between the effective exchange rate and the total trade balance. Furthermore, as Rose and Yellen (1989) argue, when estimating a trade balance model using aggregate data one needs to construct a proxy for the rest-of-the-world income. This construct is ad hoc at best and at worst misleading. These problems can be avoided altogether by employing disaggregated data.

Two studies – Rose and Yellen (1989) and Marwah and Klein (1996) – have employed disaggregated data in testing the J-Curve phenomenon. Rose and Yellen did not find a long-run effect nor any evidence supporting the J-Curve phenomenon between the U.S. and her major trading partners. Such negative findings could be due to several deficiencies. First, they define the real trade balance to be the "difference

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Remark: We would like to thank an anonymous referee for valuable comments. Of course, we alone are responsible for any possible error.

between merchandise exports and imports, measured in current U.S. dollars, deflated by the American GNP deflator." (p. 58). The evidence in Miles (1979) versus Himarios (1985) suggests that the results are sensitive to the units of measurement. Second, their method was based on Engle-Granger cointegration analysis which uses the DF or ADF tests. Since no evidence was found in favor of cointegration, the short-run analysis was based on simple autoregressive analysis, rather than an error-correction modeling. However, as Kremers et al. (1992) demonstrate, when using the Engle and Granger (1987) method, the DF test may reject cointegration due to its low power. At the same time the coefficient on the error-correction term in the corresponding dynamic model could be highly significant supporting cointegration. They argue that "The error-correction-based test is preferable because it uses available information more efficiently than the Dickey-Fuller test (Kremers et al. 1992: 325)." Finally, in estimating their simple VAR model, no attempt was made to use an objective criterion when selecting the lag length of each variable.

The second study, Marwah and Klein (1996), also employs bilateral data between U.S. and her five largest trading partners and between Canada and her five largest trading partners with mixed results. One major deficiency in this later study is the use of nonstationary data. Since the model is estimated using the level of each time-series variable without checking for unit roots, the estimates can suffer from the "spurious regression" problem. Thus, the main purpose of this paper is to expand the literature on the short-run and the long-run relationship between the trade balance and the exchange rate on a bilateral basis, after correcting the shortcomings of previous research. To this end, we introduce a reduced-form trade balance model in Section II. Section III explains a relatively new method and our empirical results. Section IV concludes. Data definition and sources are cited in an Appendix.

II. The Trade Balance Model

The model employed here will be similar to that of Rose and Yellen (1989).[1] However, the trade balance is measured as the ratio of U.S. imports from trading partner *i* over her exports to the same trading partner rather than the difference between imports and exports. Bahmani-Oskooee (1991) has argued that this measure is not sensitive to the units

[1] For theoretical derivation of the reduced-form model, see Rose and Yellen (1989: 54–55).

of measurement and it reflects the movement of the trade balance in real or nominal terms. Furthermore, it allows us to specify the model in Log form such that the first-differenced variables measure the rate of change. Thus, the model takes the following form:

$$\ln TB_{jt} = a + b \ln Y_{U.S.t} + c \ln Y_{jt} + d \ln EX_{jt} + \varepsilon_t , \qquad (1)$$

where TB_j is the ratio of U.S. imports from country j to her exports to the same country; $Y_{U.S.}$ is the U.S. real GDP set in index to make it unit-free; Y_j is the index of real GDP of trading partner j and EX_j is the real bilateral exchange rate defined in a way that a decrease reflects a real depreciation of the U.S. dollar against the currency of trading partner j. As far as the expected signs are concerned, under normal condition we would expect U.S. income to carry a positive coefficient. As $Y_{U.S.}$ rises, the U.S. will import more causing the *TB* variable to rise. However, if the increase in U.S. income is due to an increase in the production of import-substitute goods, imports may actually decline, yielding a negative estimate for b. By the same token, country j's income could carry a negative or positive coefficient as well. If real depreciation of the dollar, i.e., a decrease in *EX* is to discourage U.S. imports and encourage her exports (thus, improve the trade balance), we would expect the estimate of d to be positive. Equation (1) along with its short-run dynamic adjustment, to be explained in the next section, is the basis of our empirical analysis to which we turn next.

III. The Method and the Results

Since our interest is to detect the short-run as well as the long-run response of the bilateral trade balance to real bilateral exchange rate changes, the appropriate method is to employ error-correction modeling and cointegration techniques. The first step in applying such techniques is to identify the size and location of the autoregressive roots. To determine whether the variables in the model are characterized by unit roots requires the application of one of many possible tests. However, the existing tests for unit roots can at times yield different outcomes (Bahmani-Oskooee 1998). Due to this uncertainty Pesaran and Shin (1995) and Pesaran et al. (1996) introduce yet another method of testing for cointegration. The approach known as the Autoregressive Distributed Lag (ARDL) approach has the advantage of avoiding the classification of variables into I(1) or I(0), so unlike standard cointegration tests, there is no need for unit root pre-testing. The error-correction version of the ARDL model pertaining to the variables in (1) is as follows:

$$\begin{aligned}\Delta \ln TB_{j,t} = {} & a_0 + \textstyle\sum_{i=1}^{n} b_i \,\Delta \ln TB_{t-i} + \sum_{i=1}^{n} nc_i \,\Delta \ln Y_{U.S.,t-i} \\ & + \textstyle\sum_{i=1}^{n} d_i \,\Delta \ln Y_{j,t-i} + \sum_{i=1}^{n} f_i \,\Delta \ln EX_{j,t-i} \\ & + \delta_1 \ln TB_{t-1} + \delta_2 \ln Y_{U.S.,t-1} + \delta_3 \ln Y_{j,t-1} \\ & + \delta_4 \ln EX_{j,t-1} + \varepsilon_t .\end{aligned} \tag{2}$$

The ARDL procedure then involves two stages. In the first stage, the null hypothesis of "non-existence of the long-run relationship" defined by H_0: $\delta_1 = \delta_2 = \delta_3 = \delta_4 = 0$ is tested against the alternative of H_1: $\delta_1 \neq 0$, $\delta_2 \neq 0$, $\delta_3 \neq 0$, $\delta_4 \neq 0$. The relevant statistic to test the null is the familiar F-statistic. However, the asymptotic distribution of this F-statistic is non-standard irrespective of whether the variables are I(0) or I(1). Pesaran et al. (1996) have tabulated two sets of appropriate critical values. One set assumes all variables are I(1) and another assumes that they are all I(0). This provides a band covering all possible classifications of the variables into I(1) and I(0) or even fractionally integrated. If the calculated F-statistic lies above the upper level of the band, the null is rejected, indicating cointegration. If the calculated F-statistic falls below the lower level of the band, the null cannot be rejected, supporting lack of cointegration. If, however, it falls within the band, the result is inconclusive. In such an inconclusive case, following Kremers et al. (1992), the error-correction term will be a useful way of establishing cointegration. Since data are quarterly, we impose four lags on each first-differenced variable in (2) and provide the result of F-test for cointegration in Table 1.[2]

Recall that a significant F-statistic which tests the joint significance of the lagged level of the variables in (2) will be an indication of cointegration among the variables involved. It is clear from Table 1 that in the results for France, Germany, Italy, and Japan, the calculated F-statistic is greater than or close to the upper-bound critical value, rejecting the null of no cointegration. However, in the case of the U.K., the null cannot be rejected due to the fact that our calculated F-statistic is less than the lower-bound critical value. Finally, in the case of Canada we have an inconclusive outcome because the calculated F-statistic is less than the upper-bound critical value but greater than the lower bound. As indicated above, the results in Table 1 were obtained after imposing

[2] All calculations are carried out by MFIT4.0, a statistical package by Pesaran and Pesaran (1997).

Table 1 – *The Result of F-Test for Cointegration Among the Variables of the Bilateral Trade Balance between the U.S. and Trading Partner J*

Trading partner *j*	Calculated F-statistic
Canada	2.57
France	4.39
Germany	6.97
Italy	3.56
Japan	7.31
U.K.	2.03

Note: At the 10 percent level of significance when there is an intercept but no trend in the error-correction model, the critical value bounds of the F-statistic are 2.42 and 3.57.

only four lags on each of the first-differenced term in (2). As a sensitivity analysis, we changed the lag order and repeated the analysis. The results were somewhat sensitive to the choice of the lag order. However, this should be no concern at this stage due to more efficient results of the second stage.

Once we have established the existence of cointegration, we move to the second stage of the procedure which involves estimating the error-correction model (2). The main aim here is to detect the short-run dynamics. If the variables were found to be cointegrated, the lagged level of the variables which jointly together form the lagged error-correction term must be retained. However, even if there is no cointegration, we still retain the lagged error term to determine its significance and thus the long-run relationship. As mentioned above, this is an alternative, but an efficient way of establishing cointegration in the sense of Engle and Granger (1987). In this stage, we employ the adjusted R^2 criterion to select the lag length of each variable. For brevity of presentation, we only report the coefficient estimates of exchange rate ($\Delta \ln REX_{t-i}$) and the lagged error-correction term denoted by EC_{t-1} in Table 2.[3]

As indicated before, the short-run effects of depreciation are reflected in the coefficient estimates obtained for the lagged value of the first-differenced exchange rate variable. Furthermore, negative coefficients

[3] Full information estimates of each model between U.S. and each trading partner which include coefficient estimates of all variables in (2) are available from the authors upon request. In these results we have also tried Akaike's Information Criterion to select the lag length which yielded similar results.

Table 2 – *Coefficient Estimates of Exchange Rate and Error-Correction Term*

	Trading partner					
	Canada	France	Germany	Italy	Japan	U.K.
$\Delta \ln REX_t$	0.26 (0.82)	0.22 (0.77)	–0.24 (0.96)	–0.29 (0.96)	–0.27 (1.31)	0.22 (0.76)
$\Delta \ln REX_{t-1}$	0.53 (1.71)	–0.15 (0.49)	–0.73 (2.33)	–0.83 (2.89)	–0.45 (1.86)	–0.29 (0.96)
$\Delta \ln REX_{t-2}$		–0.26 (0.88)	–0.87 (2.80)	–0.69 (2.44)	–0.35 (1.64)	0.01 (0.04)
$\Delta \ln REX_{t-3}$		–0.11 (0.35)	–1.54 (4.90)		–0.52 (2.50)	–0.25 (0.87)
$\Delta \ln REX_{t-4}$		–0.36 (1.14)	–0.70 (2.14)		–0.31 (1.43)	0.06 (0.23)
$\Delta \ln REX_{t-5}$		–0.05 (0.16)	–0.91 (3.38)		–0.58 (2.52)	0.96 (3.66)
$\Delta \ln REX_{t-6}$		–0.09 (0.36)	–0.33 (1.24)		–0.32 (1.33)	–0.42 (1.54)
$\Delta \ln REX_{t-7}$		–0.84 (3.09)	–0.66 (2.48)			
$\Delta \ln REX_{t-8}$			–0.46 (1.88)			
$\Delta \ln REX_{t-9}$			–0.61 (2.43)			
$\Delta \ln REX_{t-10}$						
EC(–1)	–0.41 (3.75)	–0.76 (3.47)	–1.18 (5.31)	–0.34 (3.46)	–0.38 (4.29)	–0.19 (1.64)

Note: Number inside the parentheses below each coefficient is the absolute value of t-statistic.

followed by positive ones will support the J-Curve phenomenon. It is clear from Table 2 that in none of the cases, the coefficient estimates follow any specific pattern. For example, while in the case of Canada, both coefficients are positive, in the results for Germany, Japan, and Italy they are all negative. In the cases of France and the U.K., there are positive as well as negative coefficients with no specific pattern. The negative coefficients obtained for the exchange rate variable in most cases, however, should not be interpreted as an adverse effect of depreciation on the trade balance. Hsiao (1981: 95) has argued that "the negative autoregressive coefficients are not counter-intuitive because they are coefficients of the filtered data. If the model is represented in terms of the original variables, then most autoregressive coefficients will become positive". In our case, after representing the error-correction models in terms of the original variables (level rather than first difference), we gather that in each case there are positive and negative coefficients with no specific short-run pattern. This general finding supports Magee (1973) who was the first to analyze the short-run effects of exchange rate changes on the trade balance at the theoretical level. He concluded that, theoretically, the trade balance can go either way in each period.

We can now turn to the long-run effects of depreciation on the trade balance. Again, it is clear from Table 2 that the lagged error-correction term (EC_{t-1}) carries its correct negative sign and is highly significant in all cases except the last, supporting cointegration. Unfortunately, the long-run sign and coefficient estimates of variables cannot be inferred from the error-correction terms. Thus, we need to report the estimates of δ_1, δ_2, δ_3, and δ_4 from (2) that were used to form the error-correction terms in Table 2. These estimates are reported in Table 3.

It is clear from Table 3 that in all cases the real exchange rate carries a positive coefficient and it is highly significant in all cases except in the case of the U.K. The implication is that even though the short-run effects were mixed, the long-run effects of a real depreciation of the dollar against each trading partner's currency seems to have a favorable effect on their bilateral trade balances.

Table 3 – *Long-Run Coefficient Estimates of the Bilateral Trade Balance Model*

Country j	*Constant*	$\ln Y_{U.S.}$	$\ln Y_j$	$\ln REX$
Canada	4.21	–4.42	3.51	0.84
	(5.67)	(4.78)	(4.12)	(4.08)
France	–3.91	2.49	–1.92	0.78
	(2.33)	(2.05)	(1.26)	(5.27)
Germany	–3.85	1.37	–0.52	0.87
	(15.3)	(3.35)	(1.40)	(7.92)
Italy	–9.76	–2.10	2.98	0.89
	(2.38)	(0.93)	(1.43)	(2.62)
Japan	–18.3	2.67	–0.37	1.73
	(5.21)	(1.97)	(0.44)	(6.00)
U.K.	9.41	12.70	–14.70	0.81
	(1.46)	(1.70)	(1.68)	(1.03)

Note: Numbers inside the parentheses are absolute values of the t-ratios.

As indicated before, only Rose and Yellen (1989) have used stationary bilateral data and cointegration analysis to investigate the short-run and the long-run effects of a real depreciation on the U.S. bilateral trade balance. While our short-run findings, of no J-Curve pattern, are consistent with Rose and Yellen (1989), the long-run effects are inconsistent. This could be attributed to the importance of error-correction modeling in this paper versus a simple autoregressive formulation by Rose and Yellen.[4]

IV. Conclusion and Summary

Almost all previous research which investigated the relationship between the trade balance and its determinants employed aggregate data. In this paper we employed disaggregated bilateral data from the U.S. and six of her largest trading partners to investigate the short-run and the long-run response of the trade balance to a currency depreciation. The methodology was based on a new cointegration technique advanced by Pesaran and Shin (1995) and Pesaran et al. (1996), known as the ARDL approach. The main conclusion of the paper could be summarized by saying that while there was no specific short-run pattern supporting the J-Curve phenomenon, the long-run results supported the economic theory, indicating that a real depreciation of the dollar has a favorable long-run effect on the U.S. trade balance with her six trading partners. This later finding contradicts at least one previous research paper.

Appendix

Data Definition and Sources

Quarterly data over 1973I–1996II are employed to carry out the empirical work. The data come form the following sources: (*i*) *Direction of Trade Statistics* of IMF, various issues; (*ii*) *International Financial Statistics* of IMF, various issues.

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[4] We also applied Johansen's cointegration technique to the bilateral trade models. After varying the order of VAR, cointegration was confirmed in all the cases. Furthermore, estimates of cointegrating vectors revealed that real depreciation has a favorable effect on the trade balance in each case, a result consistent with that found using Pesaran and Shin's (1995) method.

Variables:

TB_j = U.S. trade balance with trading partner j defined as the ratio of U.S. imports from country j over her exports to j. All data come from source (i).

$Y_{U.S.}$ = U.S. real GDP from source (ii). It is set in index form to make it unit-free.

Y_j = Real GDP of trading partner j from source (ii) except German GDP which comes from *OECD Main Economic Indicators*. Again, it is set in index form to make it unit-free.

EX_j = Real bilateral exchange rate between U.S. dollar and each trading partner's currency. It is defined as $(P_{U.S.} \cdot NEX_j)/P_j$, where $P_{U.S.}$ is the U.S. GDP deflator from source (ii), P_j is the GDP deflator in each trading partner from source (ii), and NEX_j from source (ii) is the nominal bilateral exchange rate defined as number of j's currency per unit of the U.S. dollar. Thus, a decline in EX is a reflection of real depreciation of the dollar.

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