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RELATIVE RESPONSIVENESS OF TRADE FLOWS TO A CHANGE IN PRICES AND EXCHANGE RATE IN DEVELOPING COUNTRIES

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In 1950 Orcutt conjectured that a country's trade flows could respond to a change in exchange rate quicker than they do to a change in relative prices. Previous research that supported Orcutt's hypothesis employed non-stationary data rendering the results to suffer from spurious regression problem. When we account for stationarity of the data by using cointegration and error-correction modeling, no strong evidence is found in support of the Orcutt's hypothesis. The findings in this paper for developing countries are similar to those found for developed countries.

Keywords: Trade Flows, Relative Prices, Exchange Rate, Bounds Testing
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1. INTRODUCTION

Global economic integration has introduced new growth opportunities for developing countries as well as severe macro problems, such as increasing trade deficits. To cope with trade deficits, policy makers face various policy decisions that affect trade flows. In addition to the level of economic activity, the exchange rate and relative prices are the two important determinants of the trade flows. Policy makers could improve the trade balance either by devaluing the currency or by altering the relative prices through imposing tariff or providing subsidy. When deciding which policy tool to utilize, policy makers could be concerned with the time path and magnitude of the response of trade flows to changes in exchange rate and changes in relative prices so that they make sound and efficient decisions. Indeed, Orcutt (1950) was the first to argue that trade flows might respond differently to changes in exchange rate and changes in relative prices. A few empirical studies that have tested the hypothesis, have provided mixed results.

Junz and Rhomberg (1973) who used data from thirteen industrialized countries, measured the partial correlations between trade flows, prices and exchange rates. They concluded that the response of trade flows to changes in exchange rate and prices are

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similar. However, Wilson and Takacs (1979), after criticizing Junz and Rhomberg (1973) used data from Canada, France, Germany, Japan, United Kingdom, and the U.S.A. for the period 1957-71 and estimated export and import demand functions to show that the response time of trade flows to exchange rate is shorter that their response time to a change in relative prices. Bahmani-Oskooee (1986) and Tagene (1989, 1991) found similar results when they considered experiences of a few developing countries.

The above studies were recently criticized by Bahmani-Oskooee and Kara (2003) on the ground that they all employed non-stationary data. When regression results are based on non-stationary data, they could be considered spurious. After incorporating the time series properties of the variables into testing procedure by using cointegration and error-correction modeling, Bahmani-Oskooee and Kara (2003) investigated the trade flows of nine industrialized countries (i.e., Australia, Austria, Canada, Denmark, France, Germany, Italy, Japan and the USA) and concluded that the trade flows of different countries react differently to changes in exchange rate and to changes in relative prices, rejecting the Orcutt's original conjecture.

The main purpose of this paper is to test the relative responsiveness of the trade flows to changes in exchange rate and changes in relative prices by drawing data from developing countries. We estimate directly price and exchange rate response patterns for Columbia, Greece, Hong Kong, Hungary, Indonesia, Israel, Korea, Pakistan, Poland, Singapore, South Africa, and Turkey using quarterly data over the 1973-2002 period. To that end, in section 2 we introduce the trade models and the methods to estimate them. Section 3 presents the empirical results. Finally, section 4 summarizes and concludes the study. Data definition and sources are cited in the Appendix.

2. THE MODELS

The models adopted in this study are conventional in the sense that they are also employed by previous studies. Therefore, only a short account will be given. In formulating the import demand function we follow Bahmani-Oskooee and Kara (2003) and adopt the following log linear formulation:

$$\ln M_t^d = a + b\ln Y_t + c\ln\left(\frac{PM}{PD}\right)_t + d\ln E_t + \varepsilon_t\,, \tag{1}$$

where M is the volume of imports; Y = domestic income; PM= the price of imports;

PD = price of domestic goods; E= nominal effective exchange rate; and ε is an error term. It is expected that an estimate of b to be positive and c to be negative. Furthermore, if a decrease in E or a depreciation is to discourage imports, an estimate of d should be positive.

The import demand function expressed by Equation (1) outlines the long-run relationship between imports and their determinants. In order to test the responsiveness of imports to a change in relative prices and to a change in the exchange rate, we need to incorporate the dynamic adjustment into Equation (1). Following recent advances in time series econometrics, this amounts to specifying (1) in an error-correction format. Following Pesaran *et al.* (2001), we employ their bound testing approach to cointegration and specify (1) in an error-correction modeling format as in Equation (2):

$$\Delta \ln M_t^d = \alpha + \sum_{i=0}^{n} \beta_i \Delta \ln Y_{t-i} + \sum_{i=0}^{n} \gamma_i \Delta \ln(\frac{PM}{PD})_{t-i} + \sum_{i=0}^{n} \lambda_i \Delta \ln E_{t-i}$$
$$+ \sum_{i=1}^{n} \theta_i \Delta \ln M_{t-i}^d + \delta_1 \ln Y_{t-1} + \delta_2 \ln(\frac{PM}{PD})_{t-1} \tag{2}$$
$$+ \delta_3 \ln E_{t-1} + \delta_4 \ln M_{t-1} + u_t .$$

Note that (2) is a standard VAR model with the addition of a linear combination of the lagged level of variables as a proxy for lagged error term in Engle-Granger (1987) sense. Thus, the first step in estimating (2) is to justify the retention of the lagged level of all variables. Pesaran *et al.* (2001) show that this could be done by using the standard *F*-test with new critical values that they tabulate. If calculated *F* is significant, then the lagged level of variables are to be retained and thus, they are said to be cointegrated. A main feature of this relatively new approach is that it does not require pre-unit root testing of the variables and the new critical values take account of unit-root properties. Of course, it is clear that the results of the *F*-test will be sensitive to the number of lags imposed on each first differenced variable (Bahmani-Oskooee and Brooks (1999) and Bahmani-Oskooee and Kara (2003)). Once cointegration is established, Pesaran *et al.* (2001) propose estimating (2) by using a set criterion in selecting the optimum number of lags to infer the dynamics of the model. This second step is important for our analysis because we are interested in finding out whether lags are shorter on the relative price term or on nominal exchange rate.

In formulating the export demand function, again, we adopt the formulation by Bahmani-Oskooee (1986) and Bahmani-Oskooee and Kara (2003) and assume that the export demand function takes the following form:

$$\ln X_t^d = a' + b' \ln YW_t + c' \ln (\frac{PX}{PXW})_t + d' \ln E_t + \omega_t , \tag{3}$$

where X^d is the world demand for a country's exports; *YW* is world income, (*PX/PXW*) is the relative price of a country's exports (*PX*) compared to the world (*PXW*), *E* is the nominal effective exchange rate and ω is an error term. It is hypothesized that an increase in world income increases the world demand for a country's exports. Thus, an estimate of b' is expected to be positive. Since an increase in a country's export price relative to the world's export price is expected to discourage exports, an estimate of c' is expected to be negative. Finally, if depreciation, i.e., a decrease in E is to stimulate exports, an estimate of d' is expected to be negative. Once again, in order to be able to assess the relative responsiveness of exports to a change in relative prices and to a change in exchange rate, we express (3) in an error-correction format as in (4):

$$\begin{aligned}\Delta \ln X_t^d = {} & \alpha' + \sum_{i=0}^{m} \beta_i' \Delta \ln YW_{t-i} + \sum_{i=0}^{m} \gamma_i' \Delta \ln\left(\frac{PX}{PXW}\right)_{t-i} + \sum_{i=0}^{m} \lambda_i' \Delta \ln E_{t-i} \\ & + \sum_{i=1}^{m} \phi_i \Delta \ln X_{t-i}^d + \theta_1 \ln YW_{t-1} + \theta_2 \ln\left(\frac{PX}{PXW}\right)_{t-1} \\ & + \theta_3 \ln E_{t-1} + \theta_4 \ln X_{t-1} + v_t. \end{aligned} \tag{4}$$

In estimating (4) we follow the same two steps procedure that was explained above for import demand function.

3. THE RESULTS

As indicated before, in this paper we estimates the models for developing countries. Quarterly data over 1973I-2002Q4 period are used from twelve developing countries, i.e., Columbia, Greece, Hong Kong, Hungry, Israel, Korea, Pakistan, the Philippines, Poland, Singapore, South Africa and Turkey. As indicated above, the first step in estimating (2) and (4) involves testing for cointegration or for joint significance of δ_1 - δ_4 in Equation (2) and θ_1 - θ_4 in Equation (4). As indicated, the results of the F-test will be sensitive to the number of lags imposed on each first differenced variable (Bahmani-Oskooee and Brooks (1999)). We tried four, six, eight, ten and twelve lags. Although the calculated F statistic was different, it was significant in most cases. For brevity of presentation we report the significant results in Table 1 but consider them preliminary. Note also that a trend variable was also included in the testing procedure. It was only excluded when it carried an insignificant coefficient.

Table 1 reveals that in most cases our calculated F-statistic is greater than its critical value which supports joint significance of all lagged level variables or cointegration among them in both models. We now shift to the second stage of estimation and impose maximum of 12 lags on each first differenced variable. In this stage, we employ AIC criterion to select the optimum number of lags. Due to volume of the results, we report them in several tables. First, in order to judge the central theme of the paper, we report in Table 2 the optimum number of lags on relative price terms and on nominal effective exchange rate.

Table 1. The Results of the F-Test for Cointegration

	Calculated F-Statistics	
	EXPORTS	IMPORTS
Columbia	3.6898 (6)	4.1879 (12)
Greece	5.053 (12)	4.756 (12)
Hong Kong	4.7071 (12)	3.6873 (6)
Hungary	5.9767 (12)	6.0846 (12)
Israel	5.1410 (10)	3.6657 (8)
Korea	4.581 (12)	4.514 (12)
Pakistan	7.3683 (12)	4.1063 (10)
Philippines	16.789 (12)	8.057 (12)
Poland	3.7971 (8)	5.1152 (12)
Singapore	5.5833 (8)	4.4363 (4)
South Africa	5.3494 (6)	3.8623 (6)
Turkey	4.233 (12)	3.978 (12)

Notes: Numbers inside the parentheses are the number of lags. The critical value of the F test at 5% level of significance is 4.08. This comes from Pesaran *et al.* (2001, p. 300).

Table 2. Lag Length on Relative Price and Nominal Exchange Rate Selected by AIC

	Import Demand		Export Demand	
	ln(PM/PD)	lnE	ln(PX/PXW)	lnE
Columbia	10	5	1	6
Greece	1	1	4	2
Hong Kong	1	1	7	11
Hungary	9	6	4	5
Israel	4	7	1	9
Korea	1	8	1	1
Pakistan	10	2	1	1
Philippines	1	8	1	1
Poland	12	11	6	12
Singapore	9	10	6	1
South Africa	5	5	5	6
Turkey	1	4	2	1

It is clear from Table 2 that there is no specific pattern. In the import demand function, the lags on exchange rate are shorter than the lags on relative prices in the results for Columbia, Hungary, Pakistan, and Poland. Exactly opposite is true for Israel, Korea, the Philippines, Singapore and Turkey. For the remaining three countries, i.e., Greece, Hong Kong, and S. Africa the lags are the same. The same is true of the export demand function. Considering the results all together, only seven out of twenty-four cases support Orcutt's (1950) conjecture and Wilson and Takacs' (1979) and Bahmani-Oskooee's (1986) empirical results that the trade flows adjust faster to a change in exchange rate than to a change in relative prices.[1] These contradictory findings could be due to stationary data employed in this study versus non-stationary data used by previous research. It is clear that each country demonstrates different response path to changes in the relative prices and the exchange rate. Thus, we may not be able to reach a general conclusion. In order to learn the size and significance of estimated short-run coefficients, we report them in Tables 3 and 4.

Table 3 reveals that in most cases there is no specific short-run response pattern. The long-run coefficient estimates are reported in Table 5. Note that in each case and for each country we report not only the coefficient estimates of the lagged level variables (normalized on the import and export variables) but also the size and significance of the linear combination of the lagged level variables represented by a lagged error-correction term (EC) in the estimation procedure. Bahmani-Oskooee and Brooks (1999) have demonstrated that a negative and significant lagged error-correction term is another way of establishing cointegration.

As can be seen from Table 5, EC_{t-1} carries its expected negative and highly significant coefficient in most of the cases, supporting the cointegration results reported in Table 1. According to the theory, the estimated coefficient of lagged error correction term should be small and carry a negative sign (between zero and one in absolute value). In addition, this coefficient shows how fast is the adjustment toward the long run values: the bigger the value (in absolute value), the faster the adjustment. Concentrating on the import demand function, we note that the domestic income (lnY) carries its expected positive and significant coefficient in most cases. The estimated elasticity is greater than one for many countries, except for Columbia, Hong Kong, Hungary, Pakistan, and South Africa. The relative import price carries its expected negative sign in all cases but it is only significant in half of the cases. Similarly, the nominal effective exchange rate (lnE) is highly significant in most instances with its expected positive sign with the exception of South Africa.

[1] However, these findings are in line with Bahmani-Oskooee and Orhan (2003) who did similar analysis for industrial countries. Note that Orcutt’s (1950) conjecture could also be interpreted as indicating more persistent influence of relative prices and short-lived influence of exchange rate on trade flows.

Table 5. Long-Run Coefficient Estimates

	IMPORT DEMAND					
	Constant	lnY	ln(PM/PD)	lnE	Trend	EC_{t-1}
Columbia	4.73 (1.8)	0.30 (2.67)	-0.61 (-2.34)	0.25 (1.36)		-0.47 (-2.64)
Greece	13.88 (1.90)	1.05 (1.22)	-1.74 (-1.18)	2.284 (0.98)		-0.29 (-1.91)
Hong Kong	1.87 (1.24)	0.53 (5.29)	-0.23 (-0.66)	-0.07 (-0.21)	0.02 (4.19)	-0.31 (-3.23)
Hungary	8.28 (4.68)	0.51 (1.85)	-0.07 (-1.77)	0.73 (7.85)		-0.29 (-2.82)
Israel	-2.47 (-5.18)	1.48 (13.80)	-0.32 (-1.93)	0.38 (1.49)		-0.54 (-2.52)
Korea	3.07 (0.53)	1.62 (7.62)	-2.94 (-3.4)	1.58 (4.29)		-0.16 (-3.27)
Pakistan	1.89 (0.75)	0.34 (1.79)	-0.13 (-1.39)	0.034 (1.97)	0.05 (1.83)	-0.76 (-7.11)
Philippines	-22.93 (-6.42)	3.27 (9.21)	2.36 (4.40)	0.535 (1.21)		-0.55 (-3.96)
Poland	-4.05 (-4.23)	1.21 (6.45)	1.01 (4.63)	0.21 (1.89)	-0.01 (-1.63)	-0.17 (-2.40)
Singapore	19.51 (2.79)	1.31 (6.81)	-3.09 (-2.45)	0.60 (0.76)	-0.01 (-1.25)	-0.35 (-4.32)
S. Africa	4.96 (1.28)	0.22 (1.66)	-2.72 (-1.55)	-2.39 (-1.79)		-0.16 (-3.06)
Turkey	-2.32 (-1.53)	1.67 (1.82)	-0.97 (-2.77)	0.88 (2.32)		-0.61 (-4.05)
	EXPORT DEMAND					
	Constant	lnY	ln(PX/PXW)	lnE	Trend	EC_{t-1}
Columbia	-1.15 (-0.43)	1.28 (2.16)	-0.48 (-5.68)	0.457 (4.12)		-0.48 (-5.06)
Greece	39.57 (2.52)	1.04 (3.40)	-0.10 (-0.20)	-1.07 (-1.37)	0.07 (3.85)	-0.06 (-3.78)
Hong Kong	15.88 (1.17)	4.41 (5.02)	-5.47 (-1.02)	-1.07 (-1.18)		-0.07 (-2.85)
Hungary	-4.25 (-0.95)	2.77 (0.69)	-1.89 (-1.97)	-0.08 (-1.16)	0.16 (1.72)	-0.15 (-3.05)
Israel	3.32 (1.12)	0.48 (1.74)	-0.50 (-1.53)	-0.024 (-1.74)	0.02 (6.13)	-0.40 (-3.49)
Korea	20.36 (4.39)	3.00 (2.79)	-1.12 (-8.00)	-0.03 (-0.53)		-0.58 (-2.43)
Pakistan	1.81 (1.45)	0.58 (1.62)	-0.55 (-2.37)	-1.26 (-2.65)		-0.63 (-3.74)
Philippines	7.22 (0.94)	0.85 (1.75)	-0.28 (-1.64)	-1.17 (-6.26)		-0.34 (-1.97)
Poland	10.55 (3.51)	0.51 (1.12)	-0.16 (-1.49)	-1.05 (-3.41)	0.01 (9.38)	-0.75 (-4.42)
Singapore	26.80 (4.47)	3.45 (3.51)	-1.48 (-9.71)	-0.51 (-1.66)	0.04 (8.14)	-0.54 (-4.17)
S. Africa	1.89 (0.35)	0.47 (1.41)	-1.55 (-3.94)	-0.49 (-1.73)		-0.34 (-3.84)
Turkey	17.26 (3.58)	1.29 (0.87)	-0.68 (-1.29)	-0.12 (-7.61)		-0.89 (-3.44)

Note: t-ratios inside the parentheses.

Turning to the long-run estimates of export demand function, we gather that the results are similar to those of import demand function. Again, the EC_{t-1} carries a negative and highly significant coefficient in all cases supporting cointegration among the variables of export demand function. The world income carries its expected positive and significant coefficient in most cases. The relative export price term carries its expected negative coefficient in all cases, and it is significant in the results for Columbia, Hungary, Korea, Pakistan, Singapore, and South Africa. In addition, the nominal effective exchange rate carries its expected negative and significant coefficient in the results for Columbia, Pakistan, the Philippines, Poland, and Turkey cases. Thus, for these countries currency depreciation is expected to stimulate their exports.

Table 6. Diagnostic Test Results

IMPORTS	Serial Correlation (A)	Functional Form (B)	Normality (C)	Heteroscedasticity (D)
Columbia	7.42	4.31	8.22	1.27
Greece	10.56	11.61	30.24	1.53
H. Kong	5.86	2.51	2.07	1.31
Hungary	17.65	2.03	4.25	0.49
Israel	7.86	3.86	1.12	1.76
Korea	10.52	1.41	30.85	5.09
Pakistan	0.89	2.10	7.11	2.77
Philippines	29.16	29.08	25.15	16.767
Poland	33.01	4.74	21.96	5.49
Singapore	4.23	4.36	9.07	0.13
S. Africa	5.94	1.61	35.77	1.58
Turkey	9.89	1.19	1.39	3.30
EXPORTS				
Columbia	15.08	2.26	3.59	7.02
Greece	40.15	1.85	68.78	13.26
H. Kong	12.87	3.20	2.36	0.24
Hungary	6.50	4.16	1.63	1.21
Israel	5.05	2.39	1.55	1.72
Korea	22.12	3.94	9.28	1.45
Pakistan	11.85	3.55	2.30	1.58
Philippines	16.49	14.07	41.18	6.82
Poland	19.34	2.06	9.60	2.01
Singapore	12.78	1.19	7.91	1.03
S. Africa	9.45	2.05	1.65	1.17
Turkey	14.92	10.94	5.79	1.62

Notes: (A) Lagrange multiplier test of residual correlation (df=4). (B) Ramsey's RESET test using the square of the fitted values (df=1). (C) Based on a test of skewness and Kurtosis of residuals (df=2). (D) Based on the regression of squared residuals on squared fitted values (df=1). The critical values of X^2 statistic with four degrees of freedom at the 5% level of significance is 9.48. The comparable values with two and one degrees of freedom are 5.99 and 3.84 respectively.

Finally, to justify the models we report in Table 6 some diagnostic statistics. Based on these statistics, our models pass the diagnostic tests in most cases except serial correlation. All these statistics are distributed as X^2 with degrees of freedom reported at the bottom of the table. Furthermore, following Bahmani-Oskooee and Kara (2003) we test for stability of all coefficient estimates of the error-correction models using CUSUM and CUSUMSQ tests. The results available from the authors upon request revealed that in most cases the plot of CUSUM and CUSUMSQ stayed within the critical values reflected by two straight lines, indicating that the estimated coefficients are stable in most cases.

Before closing we though to engage in some sensitivity analysis by relying upon a different lag selection criterion such as SBC. Table 7 reports the optimum lags selected for the exchange rate and for the relative prices.

Table 7. Lag Length on Relative Price and Nominal Exchange Rate Selected by AIC & SBC

	Import ln(PM/PD)		Demand lnE		Export ln(PX/PXW)		Demand lnE	
	AIC	SBC	AIC	SBC	AIC	SBC	AIC	SBC
Columbia	10	1	5	1	1	2	6	1
Greece	1	1	1	1	4	3	2	1
Hong Kong	1	1	1	1	7	1	11	3
Hungary	9	1	6	2	4	2	5	1
Israel	4	2	7	1	1	1	9	1
Korea	1	3	8	1	1	1	1	1
Pakistan	10	1	2	2	1	1	1	1
Philippines	1	2	8	1	1	1	1	1
Poland	12	5	11	10	6	6	12	5
Singapore	9	1	10	2	6	1	1	1
South Africa	5	2	5	5	5	2	6	2
Turkey	1	1	4	3	2	2	1	1

Clearly, the SBC selects a different lag order on both variables, in both models compared to the AIC. Once again response of trade flows to a change in relative prices and the exchange rate is country specific. Furthermore, results seem to be sensitive to lag selection criterion. For example, in the import demand model of Hungary while AIC selected shorter lags for the exchange rate as compared to the lags on relative prices supporting Orcutt, the SBC did exactly the opposite. Not only the lag orders are sensitive to lag selection criterion, so are the long-run coefficient estimates reported in Table 8.

Table 8. Long-Run Coefficient Estimates Based on the SBC Criterion

	IMPORT DEMAND					
	Constant	lnY	ln(PM/PD)	lnE	Trend	EC_{t-1}
Columbia	14.47 (2.65)	0.46 (2.49)	-1.66 (-3.63)	0.43 (1.03)		-0.19 (-2.89)
Greece	13.88 (1.90)	1.05 (1.22)	-1.74 (-1.18)	2.284 (0.98)		-0.29 (-1.91)
Hong Kong	1.87 (1.24)	0.53 (5.29)	-0.23 (-0.66)	-0.07 (-0.21)	0.02 (4.19)	-0.31 (-3.23)
Hungary	6.258 (3.84)	0.54 (2.39)	-0.09 (-0.97)	-0.74 (7.21)		-0.20 (-2.39)
Israel	-2.38 (-5.16)	1.50 (14.95)	-0.14 (-0.82)	0.16 (0.89)		-0.43 (-3.43)
Korea	3.44 (0.31)	1.45 (3.61)	-2.65 (-1.73)	1.41 (2.18)		-0.16 (-3.48)
Pakistan	-0.13 (-0.78)	0.32 (1.38)	0.08 (0.55)	0.44 (2.01)	0.01 (2.79)	-0.77 (-6.08)
Philippines	57.16 (2.15)	-9.05 (-2.16)	3.91 (2.82)	-7.21 (-2.41)		-0.46 (-2.67)
Poland	8.99 (2.78)	-0.05 (-0.15)	-0.08 (-0.24)	-1.06 (-3.03)	0.03 (2.85)	-0.27 (-2.59)
Singapore	30.74 (2.43)	1.63 (4.60)	-4.87 (-2.18)	-2.29 (-2.58)	-0.01 (-1.49)	-0.21 (-3.27)
S. Africa	-9.2 (-2.13)	2.91 (3.51)	-0.48 (-1.74)	0.54 (2.31)		-0.29 (-4.43)
Turkey	8.17 (13.91)	-0.40 (-3.13)	-2.57 (-41.87)	2.24 (40.97)		-0.17 (-8.97)
	EXPORT DEMAND					
	Constant	lnY	ln(PX/PXW)	lnE	Trend	EC_{t-1}
Columbia	-7.59 (-3.14)	2.12 (1.62)	0.31 (1.58)	0.19 (0.86)		-0.27 (-2.72)
Greece	11.08 (0.68)	-2.26 (-0.88)	-0.37 (-0.54)	-0.72 (-0.53)	0.03 (1.69)	-0.33 (-1.53)
Hong Kong	-0.67 (-0.91)	2.88 (3.50)	6.45 (1.80)	-7.83 (-2.24)		-0.07 (-3.22)
Hungary	-15.07 (-0.96)	5.39 (1.71)	1.84 (2.51)	-1.448 (-0.91)	-0.09 (-1.72)	-0.10 (-2.21)
Israel	2.18 (0.79)	0.75 (1.21)	-0.52 (-1.41)	0.04 (2.03)	0.02 (6.09)	-0.41 (-3.51)
Korea	41.36 (1.19)	7.33 (1.04)	-1.54 (-1.74)	-0.29 (-0.28)		-0.11 (-1.39)
Pakistan	1.81 (1.45)	0.58 (1.62)	-0.55 (-2.37)	-1.26 (-2.65)		-0.63 (-3.74)
Philippines	7.22 (0.94)	0.85 (1.75)	-0.28 (-1.64)	-1.17 (-6.26)		-0.34 (-1.97)
Poland	3.72 (1.56)	1.19 (2.52)	-0.24 (-1.41)	-1.11 (-7.11)	0.04 (10.53)	-0.83 (-5.76)
Singapore	25.26 (1.17)	-2.74 (-0.90)	-1.22 (-2.16)	-1.05 (-0.69)	0.04 (2.11)	-0.15 (-1.55)

S. Africa	2.54 (0.54)	-0.59 (-0.59)	1.55 (4.52)	-0.51 (-2.03)		-0.39 (-4.96)
Turkey	17.26 (3.58)	1.29 (0.87)	-0.68 (-1.29)	-0.12 (-7.61)		-0.89 (-3.44)

Note: t-ratios inside the parentheses.

4. SUMMARY AND CONCLUSIONS

In 1950 Orcutt conjectured that trade flows could react to changes in the exchange rate quicker than they do to changes in relative prices. Earlier studies provided some support for Orcutt's argument, mostly by using non-stationary data. When non-stationary data are used, the results could be considered spurious. Bahmani-Oskooee and Kara (2003) who employed cointegration and error-correction modeling and data from several industrial countries to account for deficiencies of the previous research, did not find any support for Orcutt's hypothesis. They showed that the hypothesis is country specific.

In this paper we follow the methodology of Bahmani-Oskooee and Kara (2003) and consider the experiences of several developing countries. We estimate import and export demand functions of Colombia, Greece, Hong Kong, Hungary, Israel, Korea, Pakistan, the Philippines, Poland, Singapore, South Africa, and Turkey using quarterly data over the 1973-2002 period. Like Bahmani-Oskooee and Kara (2003), we find that response time of the trade flows to a change in relative prices and to a change in nominal exchange rate is country specific and there is no general pattern. Furthermore, the results were found to be sensitive to lag selection criterion.

Appendix. Data Definitions and Sources

The quarterly data were extracted from the "IMF Financial Statistics CD, January 2004 and 2000". The study period is somewhat different for some countries. For Colombia (1979Q1-2003Q1), Greece (1973Q1-1993Q4), Hong Kong (1982Q1-2003Q2), Hungary (1979Q1-2002Q4), Israel (1979Q1-2003Q1), Korea (1973Q1-1997Q3), Pakistan (1979Q1-2003Q1), the Philippines (1979Q2-1991Q4), Poland (1985Q1-2003Q2), Singapore (1979Q1-2002Q3), South Africa (1979Q1-1999Q2), and Turkey (1989Q1-1997Q3). In addition, partial data on effective exchange rate were drawn from Bahmani-Oskooee and Mirzaie (2000).

Variables:	
M	Index of volume of imports (1995=100),
PM	Index of unit value of imports (1995=100),
PD	Index of wholesale prices (1995=100),
X	Index of volume of exports (1995=100),
PX	Index of unit value of exports (1995=100),
PXW	World export price index (1995=100),
Y	Real GDP expressed as an index 1995=100. In some cases, quarterly data were not available for some countries. Following Bahmani-Oskooee (1986), quarterly data were generated from annual GDP,
YW	Index of industrial production in industrial countries (1995=100). This is used as a measure of world income,
E	Index of nominal effective exchange rate (1995=100). Note that an increase in E represents appreciation of the domestic currency.

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