FOREIGN DIRECT INVESTMENT AND TRADE:
A CAUSALITY ANALYSIS

Oscar Bajo-Rubio
(Universidad Pública de Navarra)

María Montero-Muñoz
(Universidade de Vigo)

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ABSTRACT

We analyse in this paper the relationship between outward FDI and exports, with Spanish quarterly data for the period 1977-1992. The empirical methodology makes use of Granger-causality tests in a cointegration framework, where the order of lags for each variable has been selected by means of Hsiao's sequential approach. The tests have been performed both in a bivariate and multivariate setting, in the latter case including as additional variables a proxy for world income and the relative price of exports. Since cointegration between outward FDI and exports was found, error-correction mechanisms under three alternative specifications were included, which allowed us to discriminate between short- and long-run Granger-causality. Our results point to the existence of long-run Granger-causality from outward FDI to exports, according to a complementary relationship, for the Spanish case during the period of analysis.

Key words: Foreign direct investment, exports, Granger-causality

JEL codes: F21, F40

1. INTRODUCTION

As it is well known, last years have witnessed a growing internationalization of economic activities, in the context of a progressive liberalization of international economic relations, which has led to a spectacular increase in both goods and services exchange as well as capital movements. In particular, after 1985 a surge in foreign direct investment (FDI) would have occurred, mainly among industrialized nations, and with Southern Europe as one of its most important destinations in relative terms (Graham and Krugman, 1993). In this way, according to International Monetary Fund data, during the second half of the eighties world FDI and imports increased at yearly average rates of 41 and 12.6 per cent, respectively, whereas world gross domestic product (GDP) grew at 3.8 per cent.

However, and despite the important growth experienced by both foreign trade and FDI, the relationship between them has not been extensively explored. There is some empirical work, within the industrial organization tradition, which analyses the export behaviour of multinational firms' affiliates towards host country markets. The most common finding of this literature is that of a complementarity relationship between FDI and exports, both with industry and individual firms' data; see, among others, Lipsey and Weiss (1981,1984), Blomström, Lipsey and Kulchycky (1988), Yamawaki (1991), Graham (1996) or Pfaffermayr (1996).

On the other hand, the macroeconomic relationship between FDI and trade flows has been hardly tested. An exception is a paper by Pfaffermayr (1994), who applies Granger-causality analysis to Austrian outward FDI and exports, obtaining a complementary relationship between these variables, with causality running in both directions. Also, Lin (1995) estimates exports and imports equations augmented with several FDI variables for Taiwan *vis-à-vis* four ASEAN countries (Indonesia, Malaysia, the Philippines and Thailand), and finds a positive and significant effect for outward FDI on exports.

Our aim in this paper will be to analyse the empirical relationship between outward FDI and exports for the Spanish case at a macroeconomic level, by means of Granger-causality tests in a cointegration framework. The Spanish economy has experienced a growing integration in world markets since the early sixties, strengthened after the restoration of democracy in 1977, and definitively established after joining the then European Community (EC) in 1986. As can be seen in Table 1, the increasing outward-looking orientation of the Spanish economy has resulted in growing GDP shares of exports and (especially in last years) outward FDI. Also, the links with the other EC countries have become substantial, in particular after 1986, although they are stronger for exports than outward FDI. In this way, the Spanish experience could be of interest for countries engaged in liberalization and opening processes, such as those of Eastern Europe.

Regarding the available empirical evidence on the subject for the Spanish case, there are some studies reporting a greater trade orientation for those industries and firms receiving higher FDI inflows (Bajo and López, 1998; Moreno and Rodríguez, 1998). Turning to the studies using aggregate data, in Bajo and Montero (1995), a positive effect for accumulated FDI outflows and inflows was found, respectively, when estimating exports and imports equations; however, Alguacil and Orts (1998) obtain a negative Granger-causality running from FDI outflows to exports.

Notice that, unlike other similar studies on the subject [e. g., Pfaffermayr (1994) or Alguacil and Orts (1998)], we use data on real figures for exports and outward FDI, instead of GDP shares. Also, regarding outward FDI, we take its accumulated sum over a long period (which would be a proxy of the foreign capital stock) rather than data on FDI flows. This could be justified on the grounds that FDI strategies should be a long-run phenomenon, which might be blurred when looking at the year-to-year evolution of FDI flows.

The paper is organized as follows. Some theoretical arguments concerning the relationship between FDI and trade are reviewed in Section 2. The econometric methodology is presented in Section 3, and the empirical results are shown in Section 4. Section 5 concludes.

2. THEORETICAL ISSUES

On principle, the relationship between FDI and trade is far from being unambiguous, according to a theoretical point of view. The traditional view (Mundell, 1957) stated, in the context of the (two-good, two-factor, two-country) Heckscher-Ohlin trade model, that goods movements and factor movements were substitutes. Factor mobility induced by differences in factor prices between countries would eliminate price differentials in both goods and factor markets, so removing the basis for trade. Then, trade impediments would enhance factor movements and conversely, so that exports and FDI would be alternative ways of involvement in foreign markets. However, this result would be highly dependent on the specific assumptions made (Schmitz and Helmberger, 1970).

Later contributions showed that trade and foreign investment might be complements rather than substitutes. For instance, once a certain thresold is reached, exports could result in FDI in the destination market, aimed to exploit certain advantages intrinsic to the host country as well as trying to satisfy in a better way the specific requirements of that market. Hence, FDI would be a mean of consolidating and enlarging exportation markets (Purvis, 1972).

More generally, Markusen (1983) discusses several models in which factor movements generated by international factor-price differences lead to an increase in the volume of trade. Retaining the assumption of identical relative factor endowments between two countries, several models embodying alternative bases for trade are presented (including differences in production technology, production taxes, imperfect competition, returns to scale, and factor market distortions). In all of these cases, factor mobility leads to differences in factor proportions, which means an additional motive for trade in goods. Therefore, Markusen concludes, Mundell's result of trade in goods and factors being substitutes would be a special case which is only true if trade is based on differences in relative factor proportions (i. e., for the Heckscher-Ohlin trade model).

On the other hand, and starting from Hymer’s (1976) pioneering contribution, the theories of the multinational enterprise (MNE) state that MNEs must own some particular advantage over domestic firms in the host country. Given such an ownership advantage, it must be beneficial for the MNE to internalise it within the firm by means of FDI, provided that the foreign country possess a location advantage over the home country making FDI more profitable than exporting. This is the essence of the well-known Dunning's OLI (ownership-location-internalisation) paradigm (Dunning, 1977).

These considerations have been incorporated in formal general equilibrium models in which MNEs arise endogenously. Helpman (1984) and Helpman and Krugman (1985) combine ownership and location advantages in a monopolistic competition model with horizontally differentiated goods, where MNEs develop some specific and highly specialized inputs (such as management, marketing, and product-specific R&D), that are not tradable. So, if differences in factor endowments exist, the firms from the country relatively abundant in headquarter services become MNEs, and both intra-industry trade in differentiated products and intra-firm trade in such specialized inputs will appear. Ethier (1986) endogeneizes the internalisation decision of the MNE. He finds that both a greater uncertainty faced by the firm and (unlike the models by Helpman, and Helpman and Krugman) a greater similarity in factor endowments between countries, make FDI more likely, leading to two-way FDI and a relatively higher intra-industry and intra-firm trade. In a similar line, Barrios (1997) shows that, for a peripheral country engaged in a process of economic integration, both intermediate imports and exports of the final good would be higher as integration deepens.

The previous models refer to "vertical" FDI, i. e., when MNEs locate each stage of the production process in different countries according to relative cost advantages, which results in FDI and trade being complements. However, there are also models for "horizontal" MNEs, aimed to gain an easier access to a foreign market (for reasons of transport costs, or being closer to the final customer), which might lead to FDI and trade being substitutes rather than complements.

Brainard (1993) develops a two-sector, two-country model where firms in a differentiated-products sector choose between exporting and FDI as alternative methods of foreign market penetration. This sector is characterized by increasing returns to scale at the firm level due to some specialized input (such as R&D), scale economies at the plant level, and transport costs increasing with distance. From here, an equilibrium with MNEs is more likely the higher are scale economies at the firm level relative to those at the plant level, and the higher are transport costs relative to plant-level scale economies. Also, for intermediate ranges of transport costs and firm-level scale economies relative to those at the plant level, there can be an equilibrium with MNEs and domestic firms in the differentiated sector, with two-way trade in both differentiated products and intangible inputs. Similar results are found by Markusen and Venables (1995), who add what they refer as the “convergence hypothesis”: MNEs become more important relative to trade as countries become more similar in size, relative endowments, and technologies.

The above arguments show that there are not *a priori* theoretical reasons to ascertain a clear-cut relationship between FDI and trade, so we now turn to an empirical evaluation for the Spanish case. Our econometric methodology is presented in the next section.

3. ECONOMETRIC METHODOLOGY

As stated before, we are going to analyse the empirical relationship between aggregate outward FDI and exports for the Spanish case, by means of Granger-causality tests. As it is well known, the results from these tests are highly sensitive to the order of lags in the autoregressive process. An inadequate choice of the lag length would lead to inconsistent model estimates, so that the inferences drawn from them would be likely to be misleading. In this paper, we will identify the order of lags for each variable by means of Hsiao's (1981) sequential approach, which is based on Granger's concept of causality and Akaike's final prediction error criterion, and avoids imposing often false or spurious restrictions on the model.

Suppose two stationary variables, X_t and Y_t, on which we would like to test for Granger-causality. Consider the models:

$$X_t = \alpha + \sum_{i=1}^{p} \beta_i X_{t-i} + u_t \tag{1}$$

$$X_t = \alpha + \sum_{i=1}^{p} \beta_i X_{t-i} + \sum_{j=1}^{q} \gamma_j Y_{t-j} + v_t \tag{2}$$

and then the following steps are used to apply Hsiao's procedure:

(i) Take X_t to be a univariate autoregressive process as in (1), and compute its final prediction error criterion (FPE hereafter) with the order of lags i varying from 1 to P. Choose the lag that yields the smallest FPE, say p, and denote the corresponding FPE as $FPE_X(p,0)$.

(ii) Treat X_t as a controlled variable with p lags, add lags of Y_t to (1) as in (2), and compute the FPEs with the order of lags j varying from 1 to Q. Choose the lag that yields the smallest FPE, say q, and denote the corresponding FPE as $FPE_X(p,q)$.

(iii) Compare $FPE_X(p,0)$ with $FPE_X(p,q)$. If $FPE_X(p,0) > FPE_X(p,q)$, then Y_t is said to Granger-cause X_t, whereas if $FPE_X(p,0) < FPE_X(p,q)$, then X_t would not be Granger-caused by Y_t.

Finally, by repeating steps (i) to (iii) with Y_t as the dependent variable, whether or not X_t Granger-causes Y_t can be established.

Recall that before it was assumed that X_t and Y_t were stationary variables. However, if they are integrated of order one (i. e., first-difference stationary) and are cointegrated, equations (1) and (2) need to be amended to:

$$\Delta X_t = \alpha + \sum_{i=1}^{p} \beta_i \Delta X_{t-i} + \delta z_{t-1} + u_t \qquad (3)$$

$$\Delta X_t = \alpha + \sum_{i=1}^{p} \beta_i \Delta X_{t-i} + \sum_{j=1}^{q} \gamma_j \Delta Y_{t-j} + \delta z_{t-1} + v_t \qquad (4)$$

where z_t is the error-correction mechanism (ECM) (Engle and Granger, 1987). Notice that if X_t and Y_t are I(1) but are not cointegrated, the coefficient δ in equations (3) and (4) is assumed to be equal to zero.

Now, the previous definitions of Granger-causality for stationary variables can be applied to the case of I(1) variables from equations (3) and (4). In particular, if $FPE_{\Delta X}(p,0)>FPE_{\Delta X}(p,q)$, Y_t is said to Granger-cause X_t in the short run; and if δ is significantly different from zero, Y_t is said to Granger-cause X_t in the long run. Conversely, if $FPE_{\Delta X}(p,0)<FPE_{\Delta X}(p,q)$, X_t would not be Granger-caused by Y_t in the short run; and if δ is not significantly different from zero, X_t would not be Granger-caused by Y_t in the long run. As before, by repeating the procedure with ΔY_t as the dependent variable, the hypothesis of short-run and long-run Granger-causality from X_t to Y_t could be tested.

4. EMPIRICAL RESULTS

In this section we will show the results of applying the methodology presented in the last section to Spanish data on exports and outward FDI. We have explored Granger-causality relationships between exports and outward FDI both in a bivariate and a multivariate setting, in order to avoid possible spurious results due to the omission of some relevant variables. In the latter case, we have included as additional variables a proxy for world income and the relative price of exports.

The particular variables used in our empirical exercise are as follows (all of them transformed into logarithms):

- Exports (*LX*): Spanish total exports, valued at 1985 prices.
- Outward FDI (*LINV*): accumulated sum of gross payments for Spanish investment abroad from 1966 on, net of disinvestments receipts, valued at 1986 prices using the GDP deflator[1].
- World income (LY^*): industrialized countries' imports, valued at 1985 prices.
- Relative price of exports (*LPRX*): ratio of Spanish export prices to the industrialized countries' industrial prices, the latter adjusted by the Pta-US $ exchange rate.

The data are quarterly and cover the period 1977-1992[2]. All the data have been taken from the Spanish Ministry of Economy and Finance's database (*Síntesis de Indicadores Económicos*), except for the industrialized countries' imports and the industrialized countries' industrial prices, taken, respectively, from the IMF (*International Financial Statistics*) and the OECD (*Main Economic Indicators*) data sets.

As a first step of our empirical analysis, we tested for the order of integration of the series by means of the Dickey-Fuller test [see Dickey and Fuller (1979)]. According to the test results, shown in Table 2, the null hypothesis of a unit root was not rejected in all cases; indeed, the null of a second unit root was rejected for all the series.

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1 This procedure was the same used in Bajo and Sosvilla (1994) to compute the value of the stock of foreign capital in Spain.

2 Notice that a methodological change in the elaboration of the Spanish Balance of Payments took place in 1993, when the V edition of the Balance of Payments Manual of the IMF was adopted. This change, which mainly affected to the accounting of capital flows, makes strictly impossible the comparison of FDI figures before and after that year, so we have preferred to

Next, we tested for the presence of cointegration between exports and outward FDI. To this end, we estimated long-run equations for both bivariate and multivariate models by means of the method proposed by Phillips and Hansen (1990), robust to the presence of serial correlation and endogeneity bias. The results are shown in Table 3, where the figures in brackets below each estimated coefficient are not the standard *t*-statistics but the Phillips and Hansen's fully-modified Wald test statistics, distributed asymptotically as a χ^2 with one degree of freedom.

As can be seen in the table, in both cases a positive and statistically significant relationship between exports and outward FDI is obtained. Together with the estimated equations, we also present the results from some cointegration tests: the cointegrating regression Durbin-Watson, augmented Dickey-Fuller, and Phillips-Ouliaris (denoted by CRDW, CRADF, and CR$\hat{Z}_t$, respectively). The results from these tests allow us to reject the null hypothesis of no cointegration at the 1 per cent significance level in all cases.

Therefore, the results from tables 2 and 3 suggest that Granger-causality tests should be performed with all the variables transformed into first differences, and including an ECM in the equations to estimate. As in Ngama and Sosvilla (1991), three alternative specifications for the ECM have been tested:

a) The restricted ECM: i. e., including, in a short-run equation estimated by ordinary least squares, the lagged equilibrium error from the corresponding long-run equation estimated by the Phillips-Hansen method.
b) The unrestricted ECM (Inder, 1993): i. e., jointly estimating in one step a short-run and a (lagged) long-run equation by the method of non-linear least squares.
c) The general ECM (Banerjee, Dolado, Hendry and Smith, 1986): i. e., including, in a short-run equation estimated by ordinary least squares, all the variables appearing in the corresponding long-run equation, lagged and without imposing any restriction on them.

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end our sample period in 1992.

The results of the Granger-causality tests are shown in Table 4. We started in all cases by taking until four lags of all the variables involved, comparing the FPEs, and choosing as the best specification that with the minimum FPE; the number of lags for each variable in all the selected model specifications is indicated in the table. As stated in Section 3, short-run Granger causality is assessed by comparing the FPEs from the models excluding and including the additional variable assumed to presumably Granger-cause the dependent variable. In its turn, long-run Granger causality is tested from the significance of the coefficient on the ECM, which is evaluated from, respectively: the standard *t*-statistic (restricted ECM); the *t*-statistic according to the critical values from Banerjee, Dolado and Mestre (1998) (unrestricted ECM); and a Wald test on the joint significance of the coefficients on all the variables included in the ECM, distributed as a χ^2 (general ECM)[3].

As can be seen in the table, in all the cases considered, FPEs increase when either outward FDI or exports are added to the equations estimated with exports or outward FDI as the dependent variable and not including them, so that short-run Granger-causality is not found in any direction.

On the other hand, regarding long-run Granger-causality, the coefficients on the ECMs are always significant in the equations estimated with exports as the dependent variable (except for the unrestricted ECM in the multivariate model), but not in the case of outward FDI[4]. Therefore, it can be concluded that outward FDI would Granger-cause exports in the long run (with a positive sign), but not otherwise.

In general, our results would agree with those previously found in microeconomic studies using industry or individual firms' data (see references in the Introduction). They would also be in line with other studies with a macroeconomic focus, either by estimating equations for exports (Bajo and Montero, 1995; Lin, 1995), or by means of Granger-causality tests as in this paper (Pfaffermayr, 1994). An exception is the recent paper by Alguacil and Orts (1998), who

[3] The equations estimated with LX as the dependent variable include seasonal dummies; this is not the case, however, for the equations estimated with LINV as the dependent variable, since then seasonal dummies never proved to be significant, according to a Wald test of joint significance.

[4] Notice, however, that, in the case of the unrestricted ECM-multivariate model with LX as the dependent variable, the coefficients on the ECMs would be significant at the 1% level according to the standard *t* distribution, unlike to Banerjee, Dolado and Mestre's (1998) more stringent

reply Pfaffermayr's analysis to find a negative Granger-causality running from outward FDI to exports, using Spanish quarterly data for the period 1970-1992. However, our analysis differs from these papers in that we use a stock measure of outward FDI and not data on flows (which seems to be more appropriated), as well as real figures for both exports and outward FDI instead of GDP shares. In addition, as compared to Alguacil and Orts (1998), we take a shorter but more homogeneous sample period.

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critical values.

5. CONCLUSIONS

We have analysed in this paper the empirical relationship between outward FDI and exports, using Spanish quarterly data for the period 1977-1992, by means of Granger-causality tests in a cointegration framework. Unlike other similar studies on the subject, outward FDI was taken as its accumulated sum over a long period, instead of taking flow figures. The tests have been performed both in a bivariate and a multivariate setting, in the latter case including as additional variables a proxy for world income and the relative price of exports. Since cointegration between outward FDI and exports was found, error-correction mechanisms under three alternative specifications (i. e., restricted, unrestricted, and general) were included, which allowed us to discriminate between short- and long-run Granger causality.

The main result of the paper was the long-run Granger-causality found, in all cases but one, from outward FDI to exports, with a positive sign, which would point to a complementary relationship between outward FDI and exports running from the former to the latter. However, no short-run Granger-causality between both variables was found in any of the alternative specifications tested.

The results of this paper, then, would suggest a potentially positive effect from outward FDI into increased exports. Also, the complementarity relationship found between outward FDI and exports would suggest that an increased outward FDI is not necessarily associated with deindustrialization and employment losses in the home country, as is often claimed. In addition, our results would agree with the predictions from theoretical models of "vertical" FDI, even though in some cases a complementary relationship between FDI and trade can also be inferred from models of "horizontal" FDI. In this regard, it should be noticed that a significant share of the Spanish outward FDI has as its destination the Latin American countries, which could be presumed to be relatively less abundant in the kind of specialized inputs stressed by those theories.

Therefore, and according to the results of this paper, increased capital outflows, in the context of a process of liberalization and external opening such as the one experienced by the Spanish economy during our period of analysis, might lead to higher exports. This, in turn, would illustrate the potentially important role to be played by an increased FDI abroad as a useful tool in order to promote exports.

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TABLE 1

EXPORTS AND FOREIGN DIRECT INVESTMENT OUTFLOWS: SPAIN, 1977-1992

	Exports (billion Pta.)	Exports /GDP (%)	Exports to the EC (%)	FDI outflows (billion Pta.)	FDI outflows /GDP (%)	FDI outflows to the EC (%)
1977	775.3	8.4	49.4	11.1	0.1	-
1978	1001.4	8.9	49.4	10.5	0.1	-
1979	1221.2	9.3	51.1	15.0	0.1	24.3
1980	1462.2	9.6	53.4	23.1	0.2	13.3
1981	1888.4	11.1	46.0	29.6	0.2	15.3
1982	2258.0	11.5	48.6	63.8	0.3	14.6
1983	2838.6	12.8	50.0	37.6	0.2	19.0
1984	3778.1	15.0	51.4	45.6	0.2	31.7
1985	4099.2	14.7	52.4	52.9	0.2	30.0
1986	3804.7	11.9	60.3	52.0	0.2	40.7
1987	4195.6	11.7	63.8	93.0	0.3	63.6
1988	4686.4	11.8	65.6	156.5	0.4	49.0
1989	5134.5	11.5	66.9	228.6	0.5	57.6
1990	5642.8	11.4	69.3	326.0	0.7	53.7
1991	6225.7	11.5	70.9	391.8	0.7	62.2
1992	6605.7	11.2	71.2	329.1	0.6	55.9

Source: Spanish Ministry of Economy and Finance

TABLE 2

UNIT ROOT TESTS

A) First differences

	ΔLX	ΔLINV	ΔLY*	ΔLPRX
τ_τ	-13.59[a]	-4.41[a]	-7.35[a]	-5.51[a]
τ_μ	-13.65[a]	-4.12[a]	-7.42[a]	-5.64[a]
τ	-10.22[a]	-2.36[b]	-6.56[a]	-5.69[a]

B) Levels

	LX	LINV	LY*	LPRX
τ_τ	-2.52	-1.51	-2.55	-1.38
τ_μ	-1.02	1.48	-0.52	-1.43
τ	5.35	8.60	1.70	0.25

Notes:

(i) τ_τ, τ_μ and τ are the Dickey-Fuller statistics with drift and trend, with drift, and without drift, respectively.

(ii) (a) and (b) denote significance at the 1% and 5% levels, respectively. The critical values are taken from MacKinnon (1991).

TABLE 3

COINTEGRATION TESTS

A) Bivariate model

$$\underset{(882.4)}{\text{LX} = 4.24} + \underset{(330.4)}{0.46}\ \text{LINV}$$

$R^2 = 0.88$, CRDW = 1.73, CRADF = -5.25, CR $\hat{Z}_t$ = -7.17

B) Multivariate model

$$\text{LX} = \underset{(1.77)}{-1.09} + \underset{(7.03)}{0.16}\ \text{LINV} + \underset{(25.81)}{0.88}\ \text{LY*} - \underset{(42.04)}{0.39}\ \text{LPRX}$$

$R^2 = 0.94$, CRDW = 1.87, CRADF = -8.27, CR $\hat{Z}_t$ = -8.23

Notes:

(i) CRDW, CRADF and CR $\hat{Z}_t$ are the cointegrating regression Durbin-Watson, augmented Dickey-Fuller and Phillips-Ouliaris statistics, respectively.

(ii) The null hypothesis of no cointegration is rejected at the 1% significance level. The critical values are taken from MacKinnon (1991) and Phillips and Ouliaris (1990).

TABLE 4

GRANGER-CAUSALITY TESTS

A) Bivariate model

Dependent variable: LX

	Restricted ECM	Unrestricted ECM	General ECM
Only lags of LX			
Number of lags	p=1	p=1	p=1
FPE	0.00427	0.00395	0.00440
Test on ECM	-3.01669[a]	-3.00880[c]	9.11780[b]
Adding lags of LINV			
Number of lags	p=1; q=1	p=1; q=1	p=1; q=1
FPE	0.00438	0.00453	0.00453
Test on ECM	-3.05565[a]	-3.02457[c]	9.22289[a]
Decision			
Short-run	LINV does not cause LX	LINV does not cause LX	LINV does not cause LX
Long-run	LINV causes LX	LINV causes LX	LINV causes LX

TABLE 4 (continued)

Dependent variable: LINV

	Restricted ECM	Unrestricted ECM	General ECM
Only lags of LINV			
Number of lags	p=1	p=1	p=1
FPE	0.00123	0.00123	0.00123
Test on ECM	0.02956	0.09740	1.68128
Adding lags of LX			
Number of lags	p=1; q=1	p=1; q=1	p=1; q=1
FPE	0.00126	0.00127	0.00127
Test on ECM	-0.32248	0.06530	1.85125
Decision			
Short-run	LX does not cause LINV	LX does not cause LINV	LX does not cause LINV
Long-run	LX does not cause LINV	LX does not cause LINV	LX does not cause LINV

Notes:

(i) The equations estimated in Table 4A are:

$$\Delta LX_t = \alpha + \sum_{i=1}^{p} \beta_i \Delta LX_{t-i} + \delta ECM_{t-1} + u_t$$

$$\Delta LX_t = \alpha + \sum_{i=1}^{p} \beta_i \Delta LX_{t-i} + \sum_{j=1}^{q} \gamma_j \Delta LINV_{t-j} + \delta ECM_{t-1} + v_t$$

$$\Delta LINV_t = \alpha + \sum_{i=1}^{p} \beta_i \Delta LINV_{t-i} + \delta ECM_{t-1} + u_t$$

$$\Delta LINV_t = \alpha + \sum_{i=1}^{p} \beta_i \Delta LINV_{t-i} + \sum_{j=1}^{q} \gamma_j \Delta LX_{t-j} + \delta ECM_{t-1} + v_t$$

(ii) (a) and (c) denote significance at the 1% and 10% levels, respectively. The critical values are taken from the standard *t* distribution (restricted ECM); Banerjee, Dolado and Mestre (1998) (unrestricted ECM); and the $\chi^2(2)$ distribution (general ECM).

TABLE 4 (continued)

B) Multivariate model

Dependent variable: LX

	Restricted ECM	Unrestricted ECM	General ECM
Only lags of LX, LY*, LPRX			
Number of lags	p=1; r=2; s=2	p=1; r=2; s=2	p=1; r=2; s=2
FPE	0.00382	0.00377	0.00377
Test on ECM	-2.53270[b]	-3.25364	12.68200[b]
Adding lags of LINV			
Number of lags	p=1; q=1; r=2; s=2	p=1; q=1; r=2; s=2	p=1; q=1; r=2; s=2
FPE	0.00395	0.00389	0.00389
Test on ECM	-2.39738[b]	-3.08437	11.9070[b]
Decision			
Short-run	LINV does not cause LX	LINV does not cause LX	LINV does not cause LX
Long-run	LINV causes LX	LINV does not cause LX	LINV causes LX

TABLE 4 (continued)

Dependent variable: LINV

	Restricted ECM	Unrestricted ECM	General ECM
Only lags of LINV, LY*, LPRX			
Number of lags	p=1; r=1; s=1	p=1; r=1; s=1	p=1; r=1; s=1
FPE	0.00125	0.00136	0.00116
Test on ECM	-1.60740	0.03976	5.52857
Adding lags of LX			
Number of lags	p=1; q=1; r=1; s=1	p=1; q=1; r=1; s=1	p=1; q=1; r=1; s=1
FPE	0.00129	0.00140	0.00135
Test on ECM	-1.60439	0.04881	5.37296
Decision			
Short-run	LX does not cause LINV	LX does not cause LINV	LX does not cause LINV
Long-run	LX does not cause LINV	LX does not cause LINV	LX does not cause LINV

Notes:

(i) The equations estimated in Table 4B are:

$$\Delta LX_t = \alpha + \sum_{i=1}^{p} \beta_i \Delta LX_{t-i} + \sum_{k=1}^{r} \lambda_k \Delta LY^*_{t-k} + \sum_{l=1}^{s} \mu_l \Delta LPRX_{t-l} + \delta ECM_{t-1} + u_t$$

$$\Delta LX_t = \alpha + \sum_{i=1}^{p} \beta_i \Delta LX_{t-i} + \sum_{j=1}^{q} \gamma_j \Delta LINV_{t-j} + \sum_{k=1}^{r} \lambda_k \Delta LY^*_{t-k} + \sum_{l=1}^{s} \mu_l \Delta LPRX_{t-l} + \delta ECM_{t-1} + v_t$$

$$\Delta LINV_t = \alpha + \sum_{i=1}^{p} \beta_i \Delta LINV_{t-i} + \sum_{k=1}^{r} \lambda_k \Delta LY^*_{t-k} + \sum_{l=1}^{s} \mu_l \Delta LPRX_{t-l} + \delta ECM_{t-1} + u_t$$

$$\Delta LINV_t = \alpha + \sum_{i=1}^{p} \beta_i \Delta LINV_{t-i} + \sum_{j=1}^{q} \gamma_j \Delta LX_{t-j} + \sum_{k=1}^{r} \lambda_k \Delta LY^*_{t-k} + \sum_{l=1}^{s} \mu_l \Delta LPRX_{t-l} + \delta ECM_{t-1} + v_t$$

(ii) (b) denotes significance at the 5% level. The critical values are taken from the standard *t* distribution (restricted ECM); Banerjee, Dolado and Mestre (1998) (unrestricted ECM); and the $\chi^2(4)$ distribution (general ECM).