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The Cost of Export Subsidies

Evidence from Costa Rica

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A model is developed to estimate the effects of export subsidies on the supply of exports. With data for Costa Rica over the 1980s, it is shown that although the export subsidy scheme in operation led to an increase in exports, the direct fiscal costs of the scheme were substantial. Furthermore, the subsidy scheme led to a significant increase in imports. These results suggest that elimination of export subsidies would not have a particularly harmful effect on the trade balance, and would, in addition, increase the fiscal position and generate economic efficiency. [JEL C22, F13, F17]

IN THE 1970s, the appeal of export promotion as a development strategy began to overshadow import substitution, particularly in Latin American countries hit by the debt crisis. The export-fueled growth of several Asian countries further sparked interest in export promotion schemes.

Policymakers have been creative in designing export incentives. Most export promotion programs involve a combination of fiscal and direct incentives. A drawback scheme, or some variation of it, allowing exporters to "draw back" taxes paid on imported inputs used in the production of exported goods, is a standard incentive. Many programs offer additional tax incentives such as exemptions from domestic taxes. Other programs allow for preferential rates on public utilities, subsidized interest rates, generous wastage allowances for imported inputs, and accelerated depreciation of capital goods.

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As widespread as export promotion programs are, empirical evidence on their effectiveness in increasing exports and on their costs is scarce. These costs include fiscal expenditures on export subsidies, forgone tax revenues, indirect subsidies related to public utility rates, and the costs associated with subsidized interest rates. Full measurement of the costs would ideally also account for distortions introduced by export promotion and the costs of administering the programs.

This paper measures the impact of export subsidies on export supply and evaluates their cost. A simple model is presented in Section I. The model is estimated with data from Costa Rica, where an export subsidy scheme was introduced in 1972 and enhanced by an export contract in 1984. The direct subsidy functions as a tax credit (CATS) worth 15 percent (f.o.b.) of nontraditional exports.[1] Other export incentives are available under the export contract, such as a drawback scheme; however, data on these incentives are not available. The time-series properties of the data are evaluated, and the estimation is accomplished using a Stock and Watson (1991) estimator that allows for valid hypothesis testing on the cointegrating vector. Section II presents the estimates.

The model is used to gauge the impact of the export subsidy. It is shown in Section III that, first, exports increased by roughly 10 percent; second, each dollar spent on the program increased exports by $1.35; and finally, imports of intermediates used in the production of exports increased significantly.

In general, the export subsidy has been a very costly way to promote exports, averaging 1.2 percent of gross domestic product (GDP) between 1988 and 1989, and prompting policymakers to consider alternatives to the subsidization scheme. The model indicates that the 15 percent subsidy could be offset by an average quarterly depreciation of 7 percent. It should be noted that this rate of depreciation would replicate the behavior of total exports, and thus implicitly assumes that the rate of growth of exports obtained under the subsidy is desirable. The socially optimum level of exports, however, is not addressed in the paper. The main findings are summarized in the concluding section.

I. The Basic Model

The key assumption underlying the standard empirical trade model is that exports are not perfect substitutes for the domestic good of the ex-

[1] The subsidy rate varies with the destination of exports. Nontraditional products shipped to Europe receive 20 percent, but the majority of nontraditional exports receive 15 percent.

porting country. Goldstein and Khan (1985) argued, that support for this assumption is based on the existence of two-way trade (precluded in a perfect substitute model) and evidence of significant and persistent deviation from the law of one price.

The basic model here begins with a firm that is able to allocate production between the domestic and the export market. Thus, the firm will simultaneously determine its supply of exports together with the domestic supply. Recent theoretical work seeking to account for intraindustry trade has modeled this simultaneity.[2]

Here, a simplified version of a model formulated by Behrman and Levy (1988) is used.[3] The representative domestic firm maximizes the profit function:

$$\Pi = P(Px, Pd)Q(L, K) - (WL + RK), \tag{1}$$

where Π denotes profits; P is an exact price index of the composite output, Q;[4] Px is the export price inclusive of export subsidies, S, multiplied by the exchange rate, E;[5] Pd is the price for the firm's product in the domestic market;[6] and L and K are the labor and capital quantities used in the production process. Throughout the text, uppercase symbols will denote levels, while lowercase symbols are reserved for logs. Equation (1) is maximized subject to

$$Q = [\beta Qx^{1+\Omega/\Omega} + (1-\beta)Qd^{1+\Omega/\Omega}]^{\Omega/1+\Omega}. \tag{2}$$

Equation (2) describes a constant elasticity of substitution (CES) relationship between domestic and export output.

Profit maximization will require the firm to choose Qx, Qd, L, and K, subject to equation (2). The first two first-order conditions from the lagrangian ($\partial 1/\partial Qx$ and $\partial 1/\partial Qd$) imply

$$\frac{Qx}{Qd} = \left(\frac{1-\beta}{\beta}\right)^{\Omega} \cdot \left(\frac{Px}{Pd}\right)^{\Omega}. \tag{3}$$

Figure 1 depicts the firm's maximization problem. At point A, the

[2] There are two major explanations for intraindustry trade: first, the reciprocal dumping of homogeneous products (Brander and Krugman (1983)); and second, a combination of product differentiation and increasing returns to scale (Helpman and Krugman (1985)). These models were developed using general functional forms and do not lend themselves to an estimable form.

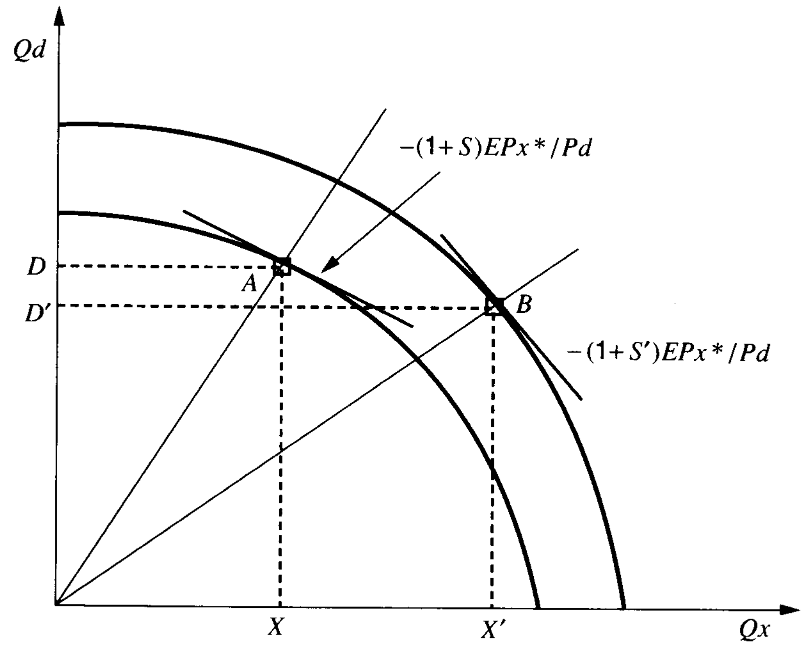
[3] These authors modeled the firm's labor and intermediate input decision. However, the present analysis is not concerned with either of these issues.

[4] This index is such that $P(Px, Pd) \cdot Q = PxQx + PdQd$.

[5] $Px = (1 + S)EPx^*$.

[6] It should be noted that Pd is potentially endogenous to the model; this issue will be discussed in Section II.

Figure 1. *Supply Decision*



firm is maximizing its profits. The firm first determines the level of composite output, Q, and allocates it according to the relative price, $(1 + S)EPx^*/Pd$. An increase in the export subsidy will have two effects. The price of the composite output increases, triggering an increase in the composite level of output, denoted by the outward shift of the output allocation curve.[7] The new subsidy increases the attractiveness of exports relative to domestic output, so that the ratio of exports to domestic output increases.[8]

To obtain the export supply curve requires combining equation (3) with the remaining three first-order conditions (requiring the value of the

[7] If initially the firm is at an equilibrium, the new composite output will require an increase in the capital stock.

[8] Figure 1 presents the case where there are increasing costs of shifting output from one market to the other. By reducing the elasticity of substitution, the transformation curve would become a right angle and the production technology would be that of joint production. The effect of a subsidy would then be exclusively an increase in the composite output. If the elasticity is very large the transformation curve turns into a straight line, so that the firm allocates all its output to one market. In that case, the supply of exports would be discontinuous.

marginal product of labor and capital to equal their corresponding prices, and the constraint (2)). In log form the export supply curve will be[9]

$$qx = b_0 + b_1(\tilde{p}x - pd) + b_2 q(W,R). \tag{4}$$

Notice that this is very similar to the original Goldstein and Khan (1978) supply equation. The difference is the scale variable, which is the composite output of the firm, whereas Goldstein and Khan used capacity, or trend, income. Here, real GDP will be used as a proxy for Q.[10]

II. Empirical Results

This section presents estimates of the model using data from Costa Rica. The series are all integrated of order one. The model is estimated accordingly, following the two-step procedure suggested by Engle and Granger (1987).[11] If regressors are endogenous or residuals are serially correlated, standard hypotheses tests on ordinary-least-squares (OLS) estimates of the cointegration vector are not valid.[12]

Four single-equation estimation methods for cointegrating vectors, which account for serial correlation and endogeneity of regressors, are available.[13] All four methods are asymptotically optimal. Phillips and Hansen (1990) proposed a fully nonparametric estimator to correct for both serial correlation and endogeneity. Saikkonen (1991), Stock and Watson (1991), and Phillips and Loretan (1989) shared the same parametric correction for endogeneity. However, Stock and Watson used a nonparametric correction to deal with serial correlation, and Phillips and Loretan suggested a parametric procedure to deal with this problem.[14]

Although all four methods are asymptotically equivalent, they do not have the same small sample properties. Both Stock and Watson and Phillips and Loretan presented Monte Carlo simulation results showing that the Phillips-Hansen estimator has greater bias and mean squared

[9] Where $b_0 \equiv -\Omega \cdot \ln(\beta)$, $b_1 \equiv \Omega$, $b_2 \equiv 1$.

[10] The demand for exports will not be modeled explicitly, but the endogeneity of regressors will be tested in Section II.

[11] The series and their unit root tests are described in the Appendix.

[12] OLS estimates of the cointegrating vector depend upon nuisance parameters for two reasons: serial correlation in the errors, and endogeneity of regressors. See Park and Phillips (1988, 1989).

[13] Phillips and Hansen (1990) showed that instrumental variable methods, although they reduce the simultaneity bias for cointegration vectors, do not eliminate the bias asymptotically. Saikkonen (1991) developed an asymptotically efficient instrumental variable estimator. He argued that the use of instruments was advisable only when the instruments and the regressors were cointegrated.

[14] The Appendix contains a brief description of these single-equation methods.

error than simple OLS. There are no Monte Carlo simulations that compare Stock and Watson with Phillips and Loretan or relate to Saikkonen's estimator. Thus, there is no a priori reason to favor either method. Nonetheless, preliminary estimation of the cointegration equation has favored the Stock-Watson approach.[15]

The Cointegration Equation

The relative price that exporters face can be expressed as $(1 + S)EPx^*/Pd$. This relative price is the combination of three elements: (1) the export subsidy, $(1 + S)$; (2) the nominal exchange rate defined as the price of foreign currency, E; and (3) the relative world price of exports in terms of the domestic price, Px^*/Pd.

If exporters are indifferent about the origin of their export revenues, one would expect that each component of this relative price would have the same effect on export supply. However, if subsidies are perceived as temporary, one would expect a relatively large short-run response to changes in the subsidy, relative to their long-run effect. This reasoning is analogous to Calvo's (1987) temporary trade liberalization argument. A temporary subsidy could induce exporters to increase supply today to take advantage of the subsidy that will not be there tomorrow. A long-run effect could occur to the extent that investment plans were shifted forward in an effort by exporters to further increase exports during the life of the subsidy. It seems plausible that a short-lived subsidy would not change investment plans and thus would not have long-run effects.

The perception that the subsidy is temporary might come from a law that states the subsidy's life span, such as Costa Rica's 1984 export contract, but this is not necessary. This perception can also be prompted by the expectation of medium-term changes in trade policies, such as joining the General Agreement on Tariffs and Trade (GATT). If the fiscal deficit is an issue, then the subsidy might come under attack because of its fiscal impact. Note that as the f.o.b. value of exports increases so will the expense of the program, increasing the likelihood that the program could be cut as exports grow. Regardless of the origin of the perception of temporariness, it will impinge on the effectiveness of the effort to promote exports.[16]

[15] Specifically, estimates of the price coefficients using Phillips and Loretan were less precise than either the Stock-Watson estimates or OLS estimates; see Hoffmaister (1991).

[16] It is also possible that the exporters might discount the nominal exchange rate, if they perceive that the authorities are not committed to keeping the exchange rate at market clearing levels.

The cointegration equation may be expressed as follows:

$$qx = \tilde{\beta}_0 + \tilde{\beta}_1 \log(1 + S) + \tilde{\beta}_2 e + \tilde{\beta}_3(px^* - pd) + \tilde{\beta}_4 q + w_t. \quad (5)$$

The analysis of Costa Rica's export supply response to export subsidies will require testing several hypotheses regarding the price coefficients: $\tilde{\beta}_1$, $\tilde{\beta}_2$, and $\tilde{\beta}_3$. The model, presented in Section I, suggests that all $\tilde{\beta}_i$ will be equal. It is also conceivable that $\tilde{\beta}_1$ will differ when exporters discount the export subsidy relative to EPx^*/Pd.[17]

If all three $\tilde{\beta}$'s were found to be equal, this would imply that exports respond equally to all three price components. This response would suggest that the export subsidy was not viewed as temporary, which would have implied a weak long-run response by exports. Since the Costa Rican subsidies are indeed temporary (their life span is ten years), a possible interpretation of that result would be that exporters expect the subsidies to be extended indefinitely, thereby suggesting that the temporary subsidy scheme is time inconsistent.

The cointegration equation (5) was estimated using Stock and Watson and 80 quarterly observations covering 1970 through 1989. The coefficients of $(1 + S)$, E, and Px^*/Pd have been allowed to differ.[18] The results of the estimation of the static model are presented in Table 1; column (1) contains the unconstrained estimation, and columns (2) and (3) present two different constrained estimations described below.[19]

Two cointegration tests were performed—the augmented Dickey-Fuller (ADF) and the augmented Phillips and Perron (APP).[20,21] Notice

[17] The hypothesis $\beta_2 = \beta_3$ is also tested. This hypothesis suggests that exporters base their output decision on the relative price, EPx^*/Pd. Rejection of this hypothesis could be accounted for by data measurement problems. It is not unlikely that exporters know Px^*, since most exports are contractual. However, it is likely that exporters face larger uncertainty surrounding Pd and E when the output decision is made. Since the data consist of actual Pd and E, it is conceivable that these series imperfectly reflect expectations regarding these variables.

[18] The coefficients of px^* and pd were found to be equal and of opposite signs. The data did not reject this hypothesis.

[19] These estimates are subject to two qualifications. First, the estimates suffer from aggregation bias, because the measure of nontraditional exports includes maquila exports that do not qualify for the subsidy. However, this bias is probably small since these products have been growing at a steady rate, reaching about 9.5 percent of nontraditional exports in 1989. For a discussion of the aggregation bias, see Goldstein and Khan (1985). Second, since the subsidy is redeemed after a period of time, the relevant measure of the subsidy is its discounted value. Unfortunately, it has not been possible to discount the subsidy, because the maturity has changed on several occasions, so that each individual export contract has a different maturity period.

[20] Campbell and Perron's (1991) suggested method determined that four lags were needed in these tests.

[21] The critical values are taken from Engle and Yoo (1987).

Table 1. *Export Supply Static Estimation*

Dependent Variable	*qx* (1)	*qx* (2)	*qx* (3)
Observations	80	80	80
R^2	0.940	0.932	0.940
$\bar{R}^2$	0.937	0.930	0.937
Sum of squared residuals	1.112	1.262	1.117
Standard error of estimate	0.122	0.128	0.121
Durbin-Watson statistic	1.164	0.970	1.144
Q-statistic	72.393	83.370	73.602
Augmented Dickey-Fuller	−3.14	−2.72	−3.09
Augmented Phillips-Perron	−5.39***	−4.82**	−5.33***
Constant	−21.92	−20.39	−22.14
	(−14.81)	(20.19)	(15.93)
$\log(1 + S)$	0.18	0.08	0.15
	(1.38)	(1.60)	(2.50)
e	0.13	0.08	0.15
	(1.85)	(1.60)	(2.50)
$px^* - pd$	0.23	0.08	0.24
	(2.09)	(1.60)	(2.40)
q	2.31	2.17	2.33
	(12.53)	(16.69)	(15.53)
H_0, χ^2_2-statistic	3.504	—	—
H_1, χ^2_1-statistic	0.116	—	—

Note: Two asterisks denote significance at the 5 percent level; three asterisks denote significance at the 1 percent level; t-statistics in parentheses.

that the ADF tests failed to reject the presence of a unit root at the 10 percent significance level; that is, according to this test, the equations do not cointegrate. However, the APP rejected a unit root at the 1 percent significance level, implying that the equations do cointegrate. The failure of ADF to reject noncointegration is likely due to the fact that this test was developed for the case where disturbances are independent and identically distributed (i.i.d.).[22]

The unconstrained estimation, shown in column (1), suggests an upward-sloping supply curve of nontraditional exports, although it is relatively price inelastic. Casual observation of the results suggests that exports respond less to nominal exchange rates or subsidies than they do to changes in relative prices. This observation provides the motivation for the null hypotheses: H_0—all price coefficients are equal; and H_1—the

[22] Campbell and Perron (1991) discuss this issue.

subsidy and the exchange rate are equal. The results of these two tests are reported at the bottom of Table 1. Both hypotheses are supported by the data.

Columns (2) and (3) present the constrained regression results under H_0 and H_1, respectively.[23] It is clear from column (2) that price elasticity falls dramatically and cointegration is obtained at 5 percent, not 1 percent significance. Furthermore, the results suggest that supply is perfectly price inelastic. This result, although statistically valid, is not persuasive. Strictly speaking, it means that regardless of the subsidy or exchange rate policy, the quantity supplied of exports remains unchanged. Furthermore, the relative profitability of exports over the domestic market, measured by the relative price, does not play a role in the long-run export supply. Thus, an increase in domestic prices vis-à-vis export prices, such as when tariff barriers are increased, does not change the firm's allocation of its output between markets, implying that tariffs do not create an anti-export bias.

Column (3) shows the estimation results when the subsidy and the exchange rate are constrained to be equal. Notice that the price elasticities are comparable to those obtained in the unconstrained regressions. It is also interesting to note that, once again, cointegration is attained at 1 percent significance.

This leaves the analyst with a dilemma. While the hypothesis tests suggest that price coefficients are equal, imposing this equality on the data renders exports perfectly price inelastic. However, when equality is imposed between the subsidy and the exchange rate, the estimates make more sense—that is, one obtains a small significant price elasticity and stronger evidence of cointegration. Also notice that the standard errors of the estimates (SEE) of the regressions in column (3) are smaller than those from both column (2) and those obtained from the unconstrained regression reported in column (1). This suggests the efficiency of imposing the second hypothesis over the unconstrained regressions.

One possible explanation for these contradictory results is that these Wald tests are asymptotically χ^2, and therefore might not perform adequately in finite samples. Monte Carlo experiments reported by Phillips and Hansen (1990) suggest that the probability distribution function is adequately approximated for sample sizes as small as 50 observations. Nonetheless, as Campbell and Perron (1991) note, these results have

[23] The Breusch (1978) and Godfrey (1978) test for serial correlation rejected the null of no serial correlation of up to fourth order; thus, OLS estimates are not efficient and standard hypothesis tests are not valid. Nonetheless, the standard F-tests were performed on OLS estimates of the cointegrating equation. These tests rejected H_0, but maintained H_1.

been obtained for small-scale models with only two or three variables in the cointegrating vector. It is not known whether these simulation results hold when the model is larger. In the rest of the paper, the estimates obtained in column (3) of Table 1 will be referred to, since they seem reasonable and are not rejected by the data.

The above results suggest that, in the Costa Rican experience, exporters have discounted the joint variations of subsidies and the nominal exchange rate relative to Px^*/pd. This evidence is consistent with the temporariness of the export subsidy as established by the export contract during 1984. Although it makes sense for a temporary change in policy to have a smaller long-run impact than permanent changes, it is hard to generalize this result, because estimates of the impact of export subsidies are scarce.

Balassa and others (1986) studied the export incentives implemented by Greece and the Republic of Korea. Their estimates for Korea—which is an important case, since it is part of the so-called Asian miracle—are comparable to those reported here for Costa Rica. Estimating the standard Goldstein-Khan export supply curve, using annual data from 1965 to 1979, they found that the elasticity of exports to $(1 + S)E$ differed from and exceeded that of Px^*/Pd—which is precisely the opposite of the result obtained here.[24,25] According to Balassa and others, exporters perceived the upward movement of $(1 + S)E$ as permanent (nonreversible), while the fluctuations of Px^*/Pd were less so. Indeed, $(1 + S)E$ increased continuously throughout their sample, whereas export incentives increased only up until 1971, reaching close to 32 percent (from about 13 percent in 1965), falling thereafter to about half this amount in 1979. Thus, it would seem that exporters perceived the depreciation of the exchange rate as permanent, which could explain the large elasticity with respect to $(1 + S)E$.

Exporters appear to perceive the origin of their export revenue differently. Tyler (1976) and Faini (1988) found, respectively, that Brazilian and Turkish exporters responded more to subsidies than to the relative price, but Moroccan exporters did not (Faini (1988)). It would therefore seem important that policymakers keep the perceptions of exporters in mind when evaluating the effects of reducing export subsidies. Specifically, the elasticity with respect to $(1 + S)E$ was found here to be less than

[24] It should be noted that the subsidies in Balassa and others (1986) are not direct export subsidies, as in these estimates. Rather, Balassa and others constructed an implicit indirect subsidization consisting of tax exemptions and other indirect subsidies.

[25] This result was confirmed independently by Jung and Lee (1986), although no hypothesis test was performed.

that of Px^*/Pd. This is important for policy decisions: using the elasticity of Px^*/Pd to evaluate the effect on exports of a reduction of export subsidies would tend to overstate, in the case of Costa Rica (or understate for the Republic of Korea), the negative impact on export revenues.

Another important empirical issue for the model is whether prices are endogenous. This will be important in Section III, since the model will be used to simulate the effect of the subsidy on export revenues. If prices were endogenous, a demand curve would be needed to measure correctly the impact of the subsidy on export revenues.

The data, however suggest that the regressors in the cointegrating equation (5) are exogenous.[26] This result is not trivial, since it implies that both Px^* and Pd are exogenous. It is also partly expected, at least for Px^*, given the size of Costa Rica's exports relative to the size of the major export market, the United States. Although Pd was more likely to be endogenous, its exogeneity is explained by the fact that the market for domestic goods is formed by a large number of suppliers, including some exporters. The data support the idea that Pd is determined by the actions of the exclusively domestic producers, while exporters take Pd as given. These results are important, since they allow one to concentrate exclusively on export supply, disregarding demand.[27]

Before examining the short-run dynamics, let us refer to the export elasticity of nontraditional exports with respect to the composite output, Q. The estimation results suggest that it is greater than 2. This value means that in the long run, for every percentage increase of the composite output, exports increase more than proportionally. This, of course, is not possible forever. Eventually, all or most output will be exported, and an increase in composite output should translate approximately into a one-to-one increase in exports. However, the typical Costa Rican exporting firm exports less than 30 percent of its output, so that for the long-run horizon it is possible that exports will increase more than proportionally. However, this long-run elasticity is not expected to fall closer to unity, as predicted by equation (4), when firms export a larger portion of their output.

[26] The evidence stems from the fact that the leds of the regressors are insignificant, either for the Stock-Watson procedure presented in the text or for the Phillips-Loretan. Additional evidence supporting the exogeneity of regressors comes from the standard Hausman specification test performed on the Phillips-Hansen estimates, which is also unable to reject the exogeneity of the regressors.

[27] This does not mean that Pd will always remain exogenous; Pd will eventually become endogenous as more and more firms allocate part of their output to exports, thereby reducing Qd. The data suggest that this has not yet occurred.

Estimation of an Error-Correction Model

The Granger representation theorem tells us that the short-run dynamics of the cointegrated process can be expressed by an error-correction mechanism of the following form:

$$\Delta y_t = \rho[y_{t-1} - \beta' x_{t-1}] + h(L)\epsilon_t, \qquad (6)$$

where y_t is the endogenous variable; x_t corresponds to a vector containing exogenous regressors; $[1, -\beta]$ is the cointegrating vector; and $h(L)$ is a lag polynomial.

Table 2 presents the estimation results for both the unconstrained and constrained error-correction models—using the cointegrating vector from column (3) in Table 1—in columns (2) and (3), respectively.[28] Note that the constrained model results suggest a relatively fast pace for the adjustment of nontraditional exports to disturbances. The estimate for ρ is approximately one half, implying that 95 percent of the adjustment is made within the first year (four quarters).[29] Notice that imposing the error-correction restriction reduces, slightly, the standard error of the estimate. This suggests the efficiency gain obtained by imposing the restriction on the data.[30]

This final model was subjected to a series of diagnostic tests. Godfrey's (1978) and Breusch's (1978) generalization of Durbin's h-test was used to test for serial correlation of up to order one and up to order four. Neither serial correlation nor autoregressive conditional heteroscedasticity effects were found, and the residuals from the regression did not exhibit significant skewness or kurtosis.

III. Effect of the Export Subsidy

The export contract is the cornerstone of Costa Rica's export promotion policy. The contract, which was established during 1984:2, governs all export incentives, including the direct export subsidy, CATS. The

[28] The specification for the error-correction model presented was arrived at after testing four lags of the difference of each variable in the cointegrating vector. Using standard F-tests, none of the lags were significant and have thus been excluded from the estimates.

[29] It should be noted that the constrained error-correction model imposes the same speed of adjustment for all variables, whereas the unconstrained version allows for speed of adjustment to vary.

[30] Also note that the estimates for the unconstrained error-correction model have appropriate signs and sizes, but are not significant. They imply that the long-run elasticity of $\log[(1 + S)E]$ is about half that of Px^*/Pd, about 0.07; the elasticity with respect to composite output is about 2.27.

Table 2. *Error-Correction Model*

Dependent Variable	Δqx (1)	Δqx (2)
Observations	79	79
R^2	0.251	0.244
$\bar{R}^2$	0.211	0.234
Sum of squared residuals	0.792	0.800
Standard error of estimate	0.104	0.102
Durbin-Watson statistic	1.972	1.920
Q-statistic	39.569	38.779
Constant	−10.38	−21.92
	(−4.38)	(−14.81)
ρ	—	−0.49
	—	(−4.90)
qx_{t-1}	−0.49	—
	(−4.90)	—
$\log[(1 + S)E]_{t-1}$	0.03	—
	(0.43)	—
$(px^* - pd)_{t-1}$	0.06	—
	(0.60)	—
q_{t-1}	1.11	—
	(4.40)	—

Note: t-statistics in parentheses.

contract has a life span of ten years during which firms are granted incentives to export.

This section measures the impact of the export subsidy on export revenues. The results suggest that exports have increased about 10 percent. The impact on export revenues is compared with the budgetary cost, which constitutes a lower bound for the cost of the subsidy. One important policy implication of the program emerges from the analysis: roughly half of the total expenditure on the subsidy has been used to increase imports of intermediate inputs.

A frequently mentioned alternative to export subsidies is exchange rate policy. The model is used to determine the impact and trade-off of reducing the export subsidy and compensating with a higher rate of depreciation.

Forecasting Performance

Before the model is used to simulate the effect of the export subsidy, its forecasting performance is gauged. To establish the model's ability to track the data during this period, it has been used to generate static forecasts of dollar exports. Roughly two thirds of the one-period forecast errors, from 1984:2 through 1989:4 (23 quarters), were less than 10 percent of the dollar value of exports. Of the remaining eight errors, five were less than 15 percent. The models' ability to forecast exports can be measured through dynamic forecasts. Accordingly, it is simulated dynamically starting from 1984:2 through 1989:4. This simulation uses the export revenue forecast for one period in the forecast for the next. Thus, the simulation forecasts just under six years into the future. Under these circumstances, roughly half of the forecast errors are under 10 percent; the other half are distributed equally between 10–15 percent, 15–20 percent, and 20 percent and above. Figure 2 shows the static forecasts in panel A and the dynamic forecasts in panel B.

To further evaluate the model's ability to forecast exports, a series of statistics that summarize in-sample forecasts during the export contract have been compiled. The model is re-estimated each quarter and used to forecast up to 12 quarters. These in-sample forecasts were used to calculate the mean error (ME), the mean absolute value error (MAE), the root mean square error (RMSE), and Theil's U-statistic. Table 3 presents the results.

The results do not indicate a problem of consistently over- or underpredicting the data, since the ME and the MAE have very different magnitudes. Notice that the model's one-step forecast erred by an average of \$4.0 million, while the absolute forecast erred by \$9.5 million. Considering that quarterly export revenue averaged \$117 million during this period, these errors are quite small. Notice, however, that the model tends to underpredict actual exports; the Theil U-statistics for forecasts for three quarters and less are poor. However, as the forecasting horizon increases, the model consistently outperforms the naive forecast. These simulations suggest that the model can track and forecast Costa Rica's dollar exports with reasonable accuracy during the period of interest.

Simulations

Once the model's ability to forecast has been evaluated, the role of the subsidy in stimulating exports can be explored. First, the model will be used to simulate baseline exports, which are compared to a simulated counterfactual where the export subsidy is set to zero during the export contract.[31] The additional export revenues will be compared with the bud-

[31] At this point it is worthwhile to refer to the Lucas critique of policy evaluation. There is growing recognition that policy evaluation is not useless. Both Cooley, Leroy, and Raymond (1984) and Sims (1987) have argued that the usual interpretation of the critique is logically flawed. Sims (1987) argues that the Lucas

Figure 2. *Model Predictive Capacity*
(1984:2–1989:4)

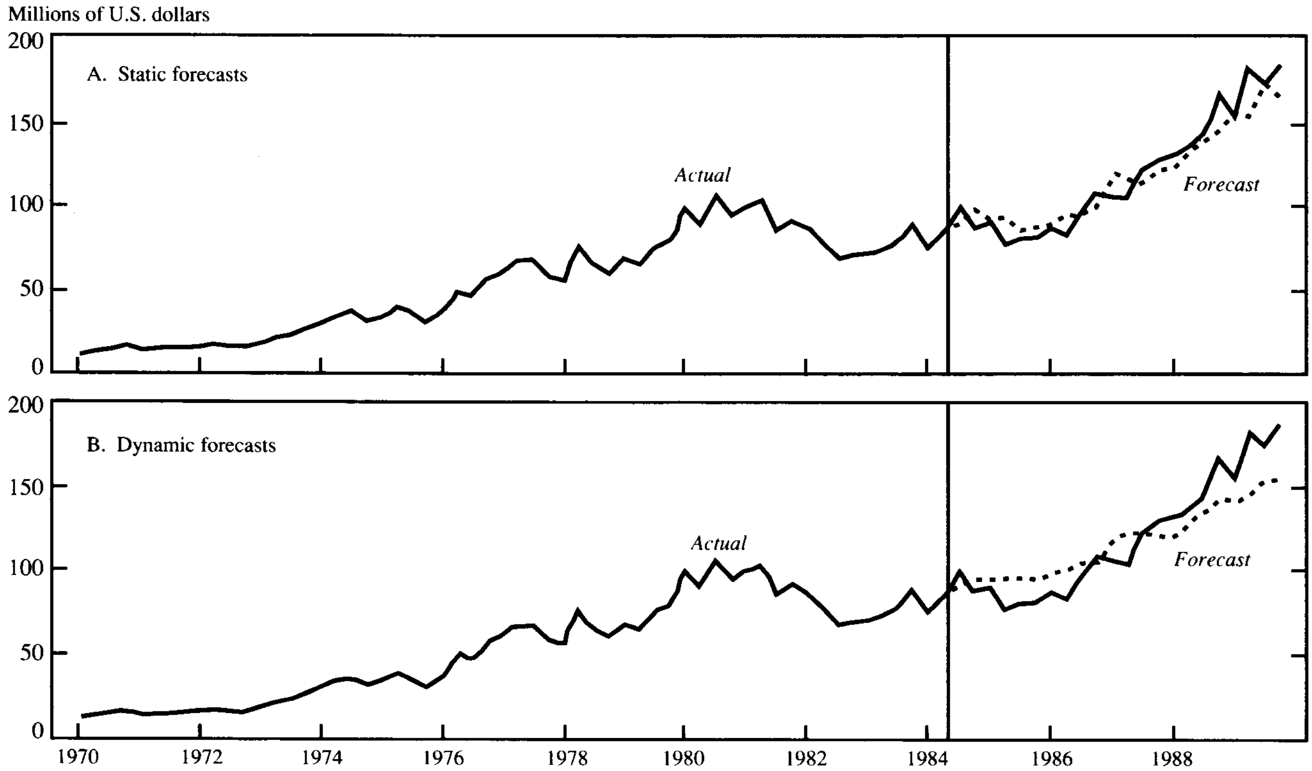


Table 3. *Forecasting Statistics*
(1984:2–1989:4)

Steps	Mean Error (In millions of U.S. dollars)	Mean Absolute Value Error (In millions of U.S. dollars)	Root Mean Square Error (In millions of U.S. dollars)	Theil *U*-Statistic	Observations
1	4.0	9.5	12.0	1.0	23
2	7.2	16.8	19.2	1.3	22
3	7.7	16.8	21.7	1.1	21
4	8.2	18.1	22.1	0.9	20
6	10.1	20.0	25.1	0.7	18
8	13.3	21.7	26.8	0.6	16
10	15.3	21.5	26.4	0.5	14
12	16.6	21.6	27.1	0.4	12

getary cost of the subsidy. Second, the model is used to simulate a common policy prescription to foster exports: exchange rate depreciation. The trade-off between export subsidy and exchange rate depreciation is assessed.

The Cost of the Subsidy

The model is used to evaluate the impact of export subsidies during the export contract period, 1984:2–1989:4. The baseline is obtained by dynamically simulating the model starting from 1984:1; in the following quarters, the subsidy was set to zero. The model was subsequently simulated to generate the counterfactual. Figure 3 shows the evolution of exports in both cases.

The model estimates that the impact on dollar exports was approximately $275 million over these 23 quarters. Given that total nontraditional exports totaled about $2.7 billion during this period, this represents roughly a 10 percent increase. This dollar amount should be compared with the cost of the subsidies. Table 4 presents the relevant data. The direct cost of the subsidization program is estimated at about $205 million,[32] corresponding to an average of 0.8 percent of GDP over the six years. Nonetheless, the cost has been increasing, averaging 1.2 percent of GDP during 1988 and 1989.

Comparing this cost with the additional exports generated suggests that each dollar spent on export subsidies has yielded a gain of about 34 percent in export revenues over the 23 quarters. However, this yield is subject to two qualifications. First, note that the cost of the export subsidies consists exclusively of the direct cost and, as such, represents a lower bound for costs. Important administrative costs are associated with the program. Each firm's application is reviewed by a joint commission—composed of representatives of the Ministry of Finance, the Ministry of Economics, and the Commission to Promote Exports (CEMPRO). The most important requirement is that the product contain at least 35 percent domestic value added. When the application is approved, the conditions—markets and subsidy rate—are published in the *Official Gazette*. For every shipment, the central bank provides the exporter with the

critique does not raise a problem when the model is "... one in which policy is already optimal and persists in being so. Thus the process of policy choice does not change the expectations formation behavior implicit in the model's structure" (p. 305). It is in this context that policy simulations are conducted later in this section.

[32] The export subsidy, CATS, has been converted into dollars using the average exchange rate.

Figure 3. *Export Subsidies*
(Simulation: 1984:2–1989:4)

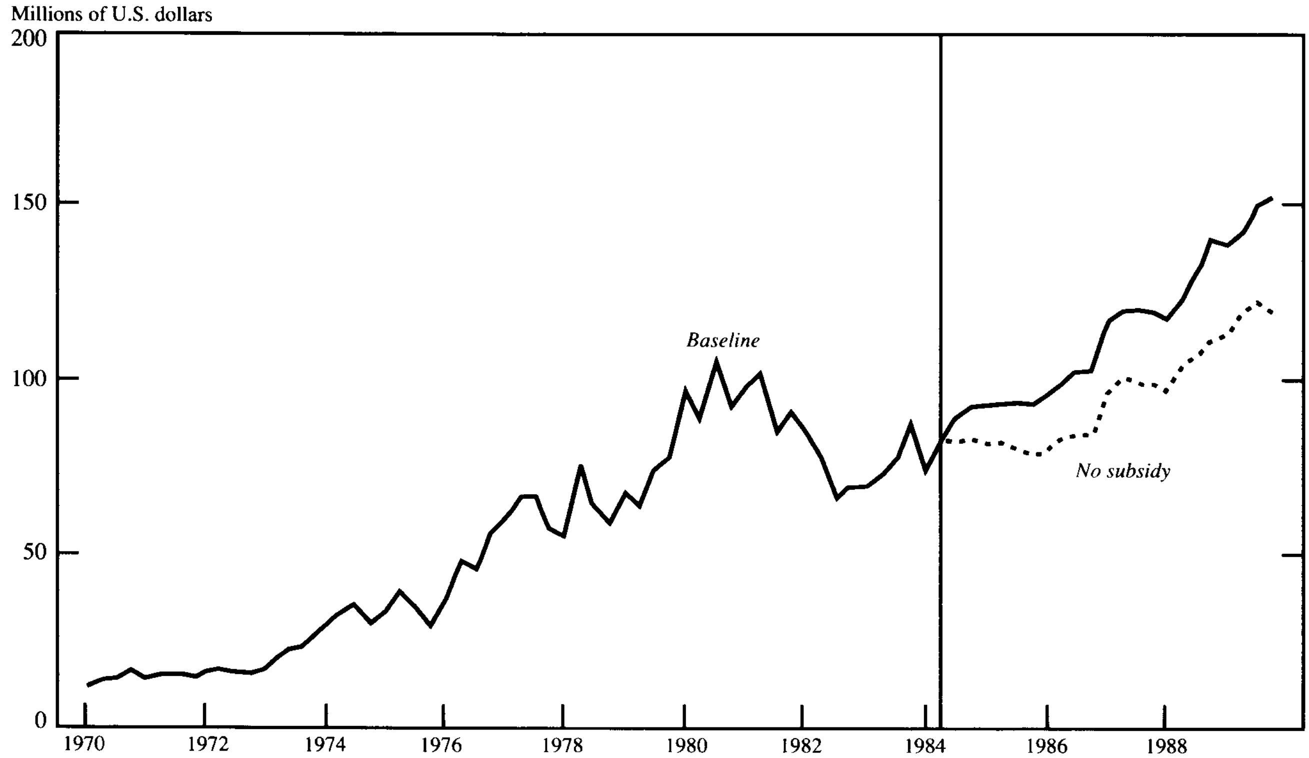


Table 4. *Simulation of the Export Subsidy*

Year	Exchange Rate (colón/U.S. dollar)	CATS		Export Response (In millions of U.S. dollars)
		colones	U.S. dollars	
1984	44.98	480.30	10.68	9.25
1985	51.31	973.50	18.97	30.98
1986	56.71	1,553.80	27.40	44.60
1987	64.15	2,030.50	31.65	54.34
1988	76.84	3,880.20	50.50	62.64
1989	82.09	5,394.90	65.72	74.73
Total		14,473.30	204.92	276.55

appropriate tax credit papers. These costs are difficult to measure, and have not been accounted for in the 34 percent yield.

The second qualification concerns the measurement of additional export revenues. Strictly speaking, the $275 million increase corresponds to gross exports, but these exports have a significant import component. On average, nontraditional exports contain about 40 percent of domestic value added.[33] This means that only $110 million has been generated, net of imports, over the 23 quarters. If the lower-bound estimate for cost is used to determine the yield of the subsidy program, the result is a net of 54 cents generated for each dollar spent. This implies that out of each dollar transferred from taxpayers to exporters via the export subsidy, 46 cents ended up subsidizing the import of intermediate inputs.[34]

Exchange Rate Depreciation

Exchange rate depreciation is frequently suggested as a way to compensate for a reduction in export subsidies. As already discussed, exports have the same elasticity with respect to the nominal exchange rate as they do with respect to the subsidy. This suggests that a reduction of the subsidy $(1 + S)$ could be offset by an equal percent change of the exchange rate.

The exact trade-off between the exchange rate and the subsidy is simple to calculate. The estimates were obtained using an index, S, for the export subsidy: S_t'/S_0'. Notice that the percentage change of $(1 + S)$

[33] Domestic value added is obtained by summing up the domestic value added of each input used to produce the final export good. Thus, the domestic value added in the final stage of production is typically less than 40 percent. Data for 1988 and 1989 provided by the Ministry of Finance were used to calculate an average for value added.

[34] It should be noted that imported intermediates used to produce exports are duty free; thus, the subsidy is not offset by tariff revenues.

can be expressed in terms of the export subsidy, S', as $S'_t/(S'_t + S'_0) \cdot \hat{S}'$. This implies that $\hat{E} < -\hat{S}'$, as long as the base used to calculate the subsidy index is positive. Thus in the long run, the reduction of the export subsidy can be compensated by a smaller percentage depreciation.

To determine the average depreciation required to compensate for the elimination of the export subsidy, a counterfactual was generated by setting the rate of depreciation constant throughout the simulated period. The rate of depreciation was set so that total dollar exports during these six years was the same as the baseline, about $2.7 billion. Full compensation requires an increase in the quarterly depreciation by 7 percent.[35] Figure 4 depicts the trajectory of exports compensated with an increase of 7 percent over the baseline.[36]

The results imply that a 25 percent reduction of the export subsidy, via the proposed tax on CATS, will reduce nontraditional exports by approximately 2.5 percent in the long run, which could be compensated by an increase of about 1.75 percent in the quarterly rate of depreciation.

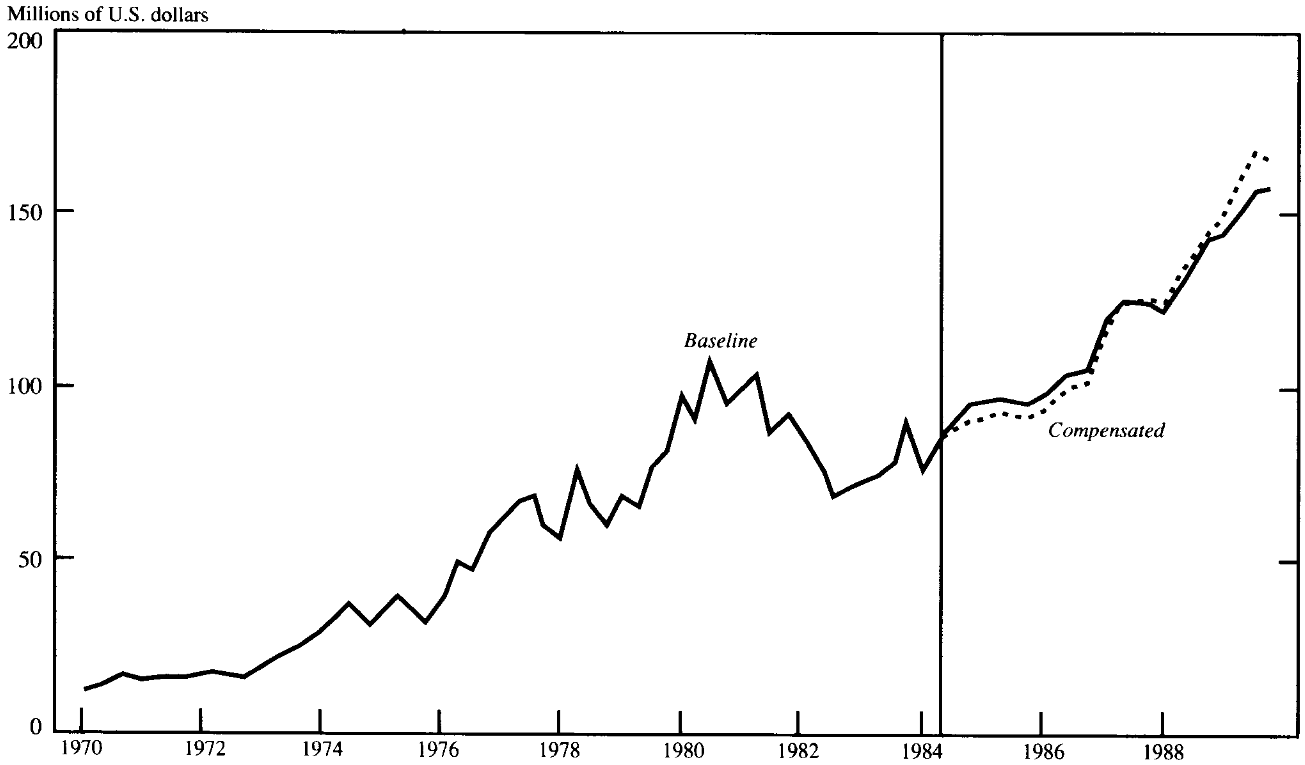
A final comment should be made about the compensating depreciation. It is possible that the depreciation will affect the domestic price of exportables, reducing its impact on exports. A higher rate of depreciation will tend to increase the domestic cost of imported goods, and can thus contribute to higher prices, which will tend to reduce the effectiveness of nominal depreciation. However, eliminating the export subsidy reduces public expenditure and thus contracts aggregate demand. This will tend to reduce inflationary pressures. In addition, the depreciation will improve the position of the Central Administration by increasing tax revenues, primarily import taxes, while expenditures in the rest of the public sector will tend to increase. The net impact on domestic prices is an empirical issue that can only be measured by a complete macro model of the Costa Rican economy. The present calculations of the compensating depreciation assume that the effects on inflation offset each other, thus replicating a concept analogous to real depreciation.[37]

[35] The rate of depreciation required was 10 percent. However, since the baseline included a 3 percent depreciation, compensation is attained with the reported rate.

[36] Notice that during the first two years the level of exports of the counterfactual is less than the baseline, while during the last two years it is larger. This implies that the 7 percent compensation does not necessarily result in the same discounted flow of export revenues as the baseline. However, the differences are relatively small.

[37] The caveat on real depreciation is due to the asymmetry between the effect of *pd* and the subsidy. The estimates here suggest that export supply is more sensitive to domestic prices than to the nominal exchange rate. The calculated effect on exports would require measuring the impact of the depreciation on the general price level and, in turn, the response by the price that exporters face in the domestic market. Given that the elasticities are different, this is not exactly the rate of depreciation accounting for inflation.

Figure 4. *Compensating Depreciation*
(Simulation: 1984:2–1989:4)



IV. Conclusions

In recent years, many countries have switched their development strategies from import substitution to export promotion. Empirical evidence regarding the effectiveness and costs of these export promotion policies, and particularly direct and indirect incentives to exports, is limited. In this paper, a direct export subsidy was introduced into a model that featured a firm facing two markets (domestic and world). The subsidy was found to increase output and switch sales to the world market.

The model analyzed the long-run supply of exports, and the short-run dynamics were generated by the data. However, explicitly modeling the short-run dynamics could prove worthwhile. A generalization of this model, where firms maximize a discounted stream of future profits, would shed light on the dynamics of export subsidies. It is likely that subsidies could trigger both intertemporal and intratemporal responses through their effect upon investment decisions. This model would be analogous to models that have analyzed the effect of terms of trade shocks on the trade balance (see, for example, Ostry (1988)). Indeed, it is likely that export subsidies would have very different effects when they are viewed as being temporary as opposed to being permanent, and modeling the short-run dynamics could be a fruitful avenue for future research.

The estimates of the long-run relationship between export supply and relative prices for Costa Rica showed strong evidence of cointegration, making it possible to estimate a constrained error-correction model, to capture the short-run dynamics of export supply. The estimates suggested that exports are price inelastic, and firms adjust within the year to shocks to the system. The forecasting performance of the estimated model was adequate.

The estimated model was used to measure the impact of the export subsidy. Exports increased by about $275 million during the six-year period, roughly a 10 percent increase in response to the 15 percent export subsidy. However, the impact on net exports was much smaller, estimated to be only about $110 million. The direct cost of the subsidy, not accounting for administrative costs, totaled about $205 million. This suggests that on average, a dollar spent on the program increased net exports by only 54 cents.

The cost of the subsidy averaged 1.2 percent of GDP during 1988 and 1989, prompting policymakers to consider modifying the scheme. The model indicates that about half of the amount spent on the program subsidized imports. Thus, it would seem that a more efficient way to spend tax dollars would be to subsidize the domestic value added of exports, which would reduce the cost of the incentive by avoiding the subsidization of imports. Alternatively, a lump-sum transfer could also avoid the subsidization of imports. Such a transfer could be set up to cover initial investment costs or the initial costs of penetrating foreign markets.

Compensating depreciation is often prescribed as a substitute for export subsidies. The simulations suggest that compensating for the 15 percent export subsidy would require an increase of 7 percent of the quarterly rate of depreciation, or about 31 percent on an annual basis. This calculation implicitly assumes that the growth of exports attained by the export subsidy is socially desired—an issue not addressed in this paper.

A subsidy is not a first-best policy; it introduces distortions that offset its benefits. Many countries have introduced export incentives to reduce the anti-export bias caused by import barriers. Given the cost of introducing export subsidies—direct on the fiscal budget and indirect through their effect on production and consumption decisions—the economically preferable policy is to eliminate the source of the anti-export bias. Thus, the first-best policy is trade liberalization.

APPENDIX

Description of Data, Test Results, and Methodology

This Appendix provides a description of the data, as well as the results of the unit-root tests and the methods for solving single equations.

Data

The following quarterly series were taken from *International Financial Statistics,* (International Monetary Fund, various years): the exchange rate; domestic price; and Px^*. The latter series was used to distribute the export price of nontraditional exports using Chow and Lin (1971).

The following annual series came from Banco Central de Costa Rica (BCCR): (1) U.S. dollar exports of nontraditional exports, which was distributed using Litterman (1984); and (2) prices of nontraditional exports, which was distributed using Chow and Lin (1971). Dollar exports were deflated using the price of nontraditional exports to obtain the quantity of exports.

The Ministry of Finance of Costa Rica provided the CATS subsidy series. An annual series for "CATS Entregados" was distributed using Chow and Lin (1971) with the quarterly series "CATS Efectivos." The "Entregados" version is analogous to a commitment series of subsidy, while "Efectivos" corresponds to cash payments. The Ministry also provided information on the domestic value added of nontraditional exports.

Quarterly GDP figures are from Hoffmaister (1992). All relevant series have been indexed to 1985.

Unit-Root Test Results

Two standard unit-root tests were applied: (1) augmented Dickey-Fuller (ADF); and (2) augmented Phillips-Perron (APP). The number of lags included in each of these tests was determined following Campbell and Perron (1991). Hall (1990) shows that this procedure will come up with the correct number of lags with probability one asymptotically, provided that the procedure starts with a sufficiently high number of lags. The test results are included in Table 5.

The test results suggest the existence of one, but not two, unit roots. This is also true when a trend and/or drift is added to the null hypothesis. Notice that the augmented Dickey-Fuller test could not reject the existence of two unit roots in most cases. The lack of power of the augmented Dickey-Fuller test is discussed by Campbell and Perron (1991).

Single-Equation Methods

To discuss the four methods mentioned above, let us introduce the following equations:

$$y_t = \beta' x_{1t} + \mu_t^{(1)} \tag{7}$$

$$\Delta x_{1t} = \mu_t^{(2)} \tag{8}$$

$$\Delta x_{2t} = \mu_t^{(3)}. \tag{9}$$

Equation (7) is the cointegrating equation, (8) is a vector of k_1 regressors included in (7), and (9) is a vector of k_2 instruments that do not appear in equation (7) and are cointegrated with the regressors in equation (8). Let $\mu = [\mu^{(1)}, \mu^{(2)}]'$ be the $(k_1 + 1)$ vector of residuals in the system (7)–(8) and let its covariance matrix be

$$\Sigma = E[\mu \cdot \mu'] = \begin{bmatrix} \sigma_{11} & \sigma'_{21} & \sigma'_{31} \\ \sigma_{21} & \Sigma_{22} & \sigma'_{32} \\ \sigma_{31} & \sigma_{32} & \Sigma_{33} \end{bmatrix}, \tag{10}$$

partitioned to conform with equations (7)–(9).

Phillips and Hansen (1990) note that for time series

$$\sigma_{21} = \sum_{s=0}^{\infty} E[\mu_0^{(2)} \mu_s^{(1)}]. \tag{11}$$

Table 5. *Unit-Root Tests*

Test	Series			
	qx	*px* − *pd*	log(1 + *S*)*e*	*q*
Augmented Dickey-Fuller				
Level	−1.440	−1.827	−0.259	−1.487
(1 − *L*)	−34.298**	−2.108	−2.050	−1.833
Augmented Phillips-Perron				
Level	−2.522	−1.609	−2.095	−2.200
(1 − *L*)	−74.770**	−31.450**	−23.788**	−600.153**

Note: Two asterisks indicate significance at the 1 percent level.

Their nonparametric correction for serial correlation adjusts OLS estimates obtained from equation (7) by adding to it: $-[x_1' x_1]^{-1} T\hat{\sigma}_{21}$, where $\hat{\sigma}_{21}$ is a consistent estimator of σ_{21}. This adjustment purges the OLS estimates of the nuisance parameters due to serial correlation.

Their "fully modified" estimator requires two corrections that are accomplished as follows. First, the left-hand-side variable in equation (7) is purged of endogeneity by the following transformation: $y_t^+ = y_t - \hat{\sigma}_{21}' \hat{\Sigma}_{22}^{-1} \Delta x_{1t}$. OLS is performed with this transformed variable, and in turn corrected for serial correlation by adding to it $-[x_1' x_1]^{-1} T\hat{\delta}^+$, where $\hat{\delta} = \hat{\Phi} \cdot [1, -\hat{\Sigma}_{22}^{-1} \hat{\sigma}_{21}]'$ and $\hat{\Phi}$ is a consistent estimate of $\Phi = \Sigma_{s=0}^{\infty} E[\mu_0^{(2)} \mu_s']$.

Stock and Watson (1991) suggest the following parametric method to deal with endogeneity of regressors. The basic idea is to make $\mu_t^{(1)}$ independent of $\mu_t^{(2)}$; to this effect they note that since $\mu_t^{(1)}$ is assumed Gaussian and stationary, then

$$E[\mu_t^{(1)} / \{\Delta x_{1t}\}] = E[\mu_t^{(1)} / \{\mu_t^{(2)}\}] = d_1(L)\Delta x_{1t},$$

where $d_1(L)$ is a two-sided lag polynomial. It should be noted that $d_1(L) = \Sigma_{i=-\infty}^{\infty} d_{1,i} \cdot L^i$ in practice is truncated. By adding and subtracting this term to (7)

$$y_t = \beta' x_{1t} + d_1(L)\Delta x_{1t} + c_{22}(L)\tilde{\mu}_t^{(2)}, \tag{12}$$

where $\tilde{\mu}^{(2)} = \mu^{(2)} - E[\mu^{(2)} / \{\mu_t^{(2)}\}]$ is independent of innovations from the left-hand-side variable by construction. Stock and Watson suggest using OLS on the dynamic equation (12). They call this estimator dynamic OLS.

This parametric correction for endogeneity is shared by Saikkonen (1991) and by Phillips and Loretan (1989), and is based on the work of Sims (1972) on causality tests. Recall that when a variable y_t causes (in Granger terms) x_{1t}, then y_t can be expressed as a linear combination of past, future, and present values of x_{1t}. Thus, future values of x_{1t} will provide information that helps in the prediction of y_t. These future values of x_{1t} are in essence Sims's causality test. Significant values for future x_{1t} provide evidence that x_{1t} is not weakly exogenous.

Equation (11) still contains serial correlation. Stock and Watson dealt with the serial correlation by correcting the covariance matrix used in the estimation of (11). The covariance matrix should be estimated using nonparametric methods, such as using a Bartlett window. They have also suggested estimating the covariance matrix using an autoregressive spectral estimator. Alternatively, they also model the errors as autoregressive processes, suggesting dynamic generalized least squares. Saikkonen (1991) suggested a different nonparametric correction that basically adds to equation (12), $d_2(L)x_{3t}$, where $d_2(L)$ is a two-sided lag polynomial. Phillips and Hansen (1990) suggested a parametric correction to deal with serial correlation. They proposed adding to equation (12) the term $d_3(L)(y_t - \beta' x_t)$, where $d_3(L)$ is a one-sided lag polynomial defined as $\Sigma_{i=1}^{\infty} d_{3,i} \cdot L^i$. Their estimator implies that the cointegrating vector enters nonlinearly; thus, it is estimated using nonlinear least squares.

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