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The Cost of Export Subsidies: Evidence from Costa Rica

Prepared by Alexander Hoffmaister 1/

Authorized for distribution by Mohsin S. Khan

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Abstract

This paper develops a model to estimate the effects of export subsidies on the supply of exports. Using data for Costa Rica over the 1980's, it is shown that while the export subsidy scheme in operation led to an increase in exports, the direct fiscal costs of the scheme were quite large. Furthermore, the subsidy scheme led to a significant increase of imports. These results suggest that elimination of export subsidies would not have a particularly harmful effect on the trade balance, and would increase the fiscal position and generate economic efficiency besides.

JEL Classification Numbers:

C22, F13, F17

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Contents

I. Introduction

The development strategy of import substitution (IS) of the 1950's and 1960's was undertaken to foster rapid industrialization. Much of Latin America, following the advice of the Economic Commission for Latin America, levied import tariffs to protect infant industries from foreign competition. However, the debt crisis and the experiences of high growth Asian countries, fueled by remarkable export growth, gave rise to interest in export promotion (EP).

Policymakers have been creative in designing export incentives. Most EP programs involve a combination of fiscal and direct incentives. A variation of a drawback scheme, allowing for exporters to "drawback" 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: income, sales and value added taxes. Other programs allow for preferential rates on public utilities, subsidized interest rates, generous wastage allowances for imported inputs, accelerated depreciation of capital goods, etc.

As widespread as EP programs are, empirical evidence regarding their effectiveness in increasing exports and their costs is scarce. These costs include fiscal expenditures on export subsidies, foregone 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 EP and the costs of administering the programs.

This paper intends to measure the impact of export subsidies on export supply and evaluate their cost. A simple model, that is useful to analyze the impact of export subsidies, is presented in Section II.

The model is estimated with the data of Costa Rica. The subsidy scheme was introduced in 1972 and enhanced with the export contract in 1984. The direct subsidy functions as a tax credit (CATS) worth 15 percent of fob 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 Stock and Watson (1991) estimator that allows for valid hypothesis testing. Section III presents the estimates.

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

1/ The subsidy rate varies with the destination of exports. Exports received 20 percent for nontraditional products shipped to Europe. However, the vast majority of nontraditional exports receives 15%.

In general the export subsidy has been very costly way to foster exports: it has averaged 1.2 percent of GDP between 1988 and 1989. This has prompted 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.

II. 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 exporting country. Goldstein and Khan (1985) argue that support for this assumption is based upon 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.

We consider 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 explain intra-industry trade, has modeled this simultaneity. There are two major explanations for intra-industry trade. First, the reciprocal dumping of homogeneous products presented by Brander and Krugman (1983), and second a combination of product differentiation and increasing returns to scale following Helpman and Krugman (1985). These models have been developed using general functional forms and do not seem to lend themselves to an estimable form.

We have chosen to use a simplified version of a model presented by Behrman and Levy (1988). 1/ The representative domestic firm maximizes the profit function:

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

where Π is profits, P is an exact price index of the composite output Q, 2/ Px is the export price 3/ the export price, inclusive of export subsidies (S) multiplied by the exchange rate E, and Pd is the price for the firm's product in the domestic market. 4/ L and K are the labor and capital quantities used

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1/ These authors model the firm's labor type decision as well as the firm's intermediate input type decision. Our analysis is not, however, concerned with either of these issues.

2/ This index is such that: $P(Px,Pd) \cdot Q = Px\ Qx + Pd\ Qd$.

3/ $Px = (1+S)EPx^*$

4/ It should be noted that Pd is potentially endogenous to the model, this issue will be discussed in Section 3.2.

in the production process. The convention of denoting levels with upper case letters, while lower case letters will denote logs will be used throughout the text.

Equation (1) is maximized subject to:

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

where (2) describes a CES relationship between domestic and export output.

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

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

Figure 1 depicts the firms' maximization problem. Point A shows a firm 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. If initially the firm is at an equilibrium, the new composite output will require an increase in the capital stock. The new subsidy increases the attractiveness of exports relative to domestic output, so that the ratio of exports to domestic output increases. 1/

To obtain the export supply curve requires combining (3) with the remaining three first order conditions (requiring the value of the marginal product of labor and capital to equal their corresponding prices, and the constraint (2)). In log form the export supply curve will be: 2/

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

1/ The diagram 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 be 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 would be a straight line, so that the firms allocate all their output to one market. In that case the supply of exports would be discontinuous.

2/ Where: $b_0 = \ln\left[(1-\beta)\left(\frac{\beta}{1-\beta}\right)^{1+\Omega}\right]\frac{\Omega}{1+\Omega}$, $b_1 = \Omega$, $b_2 = 1$.

Notice that this is very similar to the original Goldstein-Khan supply equation. The difference is the scale variable that is the composite output of the firm while Goldstein-Khan use capacity, or trend, income. We will use real GDP as a proxy for Q. 1/

III. Empirical Results

This section presents estimates of the model using data from Costa Rica. We first test the series for the existence of units roots and then proceed to estimate the model. 2/ The two step procedure to estimate error correction models, suggested by Engle and Granger (1987) is followed. When regressors are endogenous or residuals are serially correlated, hypotheses tests using ordinary least squares (OLS) estimates of the cointegration vector are not valid. We follow recent literature, e.g. Phillips and Hansen (1989), Stock and Watson (1991) and Phillips and Loretan (1989), in handling these issues.

1. Time-series properties

To test for unit roots we have applied three standard tests: (i) augmented Dickey-Fuller (ADF), (ii) augmented Phillips-Perron (APP) and (iii) Sargan-Bhargava (SB). The number of lags included in each of these tests was determined following Campbell and Perron (1991). It has been shown by Hall (1990) that this procedure will to come up with the correct number of lags with probability one asymptotically. 3/

Most test results suggest the existence of one, but not two, unit roots. 4/ 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 tests is discussed by Campbell and Perron (1991). In the following sections the model will be estimated presuming that the series are stationary in first differences. Unit root test results are included in the Appendix.

2. Estimation

Despite the "super consistency" of OLS estimates of cointegrating relationships, proven by Stock (1990b), OLS estimates have two important drawbacks: (1) they are not asymptotically optimal, and (2) it is not possible to test hypotheses about the parameters of the cointegrating vector. Park and Phillips (1988, 1989) have discussed the issue of inference in models with unit roots and cointegration. Unfortunately, OLS estimates of the

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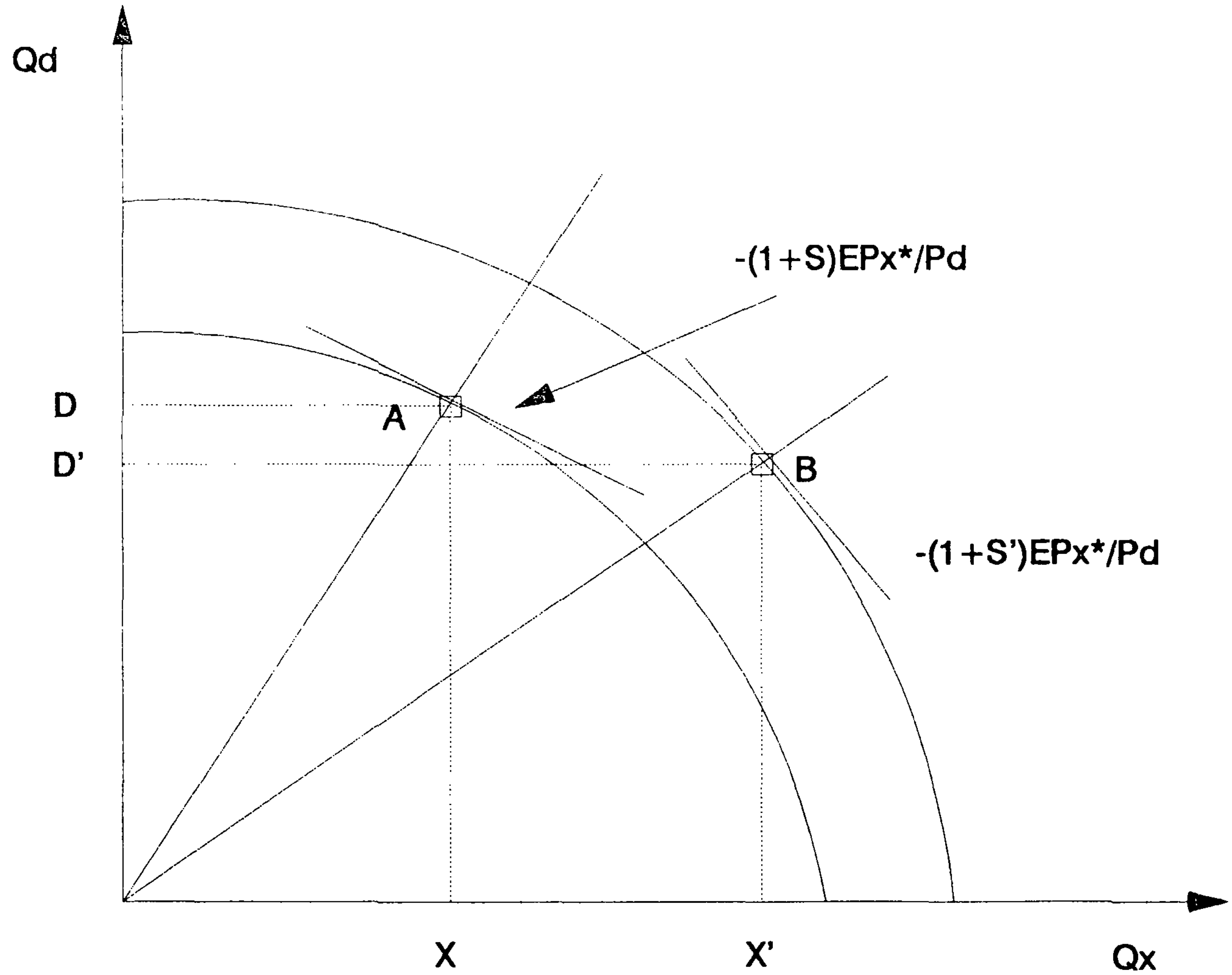
1/ We will not explicitly model the demand for exports. However, the endogeneity of regressors will be tested in Section III.

2/ The data are described in the Appendix.

3/ Provided that the procedure start with a sufficiently high number of lags.

4/ This result also holds true when a trend and/or drift is added to the null hypothesis.

Figure 1: Supply Decision



cointegrating vector depends upon nuisance parameters for two reasons: (i) serial correlation in the errors, and (ii) contemporaneous correlation between the innovations in the right hand side variable (RHS) and the left hand side (LHS) variables, that is endogeneity of regressors. 1/

Three single equation methods that account for serial correlation and endogeneity of regressors are available. 2/ All three methods are asymptotically optimal. Phillips and Hansen (1990) (P-H) propose a fully non-parametric estimator to correct for both serial correlation and endogeneity. Stock and Watson (1991) (S-W) and Phillips and Loretan (1989) (P-L) share the same parametric correction for endogeneity. However, the former uses a non-parametric correction to deal with serial correlation, while the latter suggests a parametric procedure to deal with this problem. 3/

Although all three methods are asymptotically equivalent, they do not have the same small sample properties. Both S-W and P-L present Monte Carlo simulation results showing that the P-H estimator has greater bias and mean squared error than simple OLS. Monte Carlo simulations comparing S-W with P-L are not presently available. Thus, there is no a priori reason to favor either S-W or P-L method. Nonetheless, preliminary estimation of the cointegration equation has favored S-W. 4/

a. 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) (1+S), the export subsidy, (2) E, the nominal exchange rate defined as the price of foreign currency, and (3) Px^*/Pd, the relative world price of exports in terms of the domestic price.

If exporters are indifferent about the origin of their export revenues, we expect that each component of this relative price to have the same effect upon export supply. However, if subsidies are perceived as temporary we would expect relatively large short run responses to changes in the subsidy, relative to their long run effect. This is analogous to Calvo's (1987) temporary trade liberalization argument. A temporary subsidy could induce

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1/ If the error term from the cointegrating equation is serially uncorrelated and the innovations of the LHS variables are weakly exogenous and do not Granger cause the innovations in the RHS variables, then the asymptotic distribution of the cointegrating vector is free of nuisance parameters. In such a case, OLS is asymptotically optimal and standard hypothesis tests are valid.

2/ Phillips and Hansen (1990) show that instrumental variable methods, although they reduce the simultaneity bias for cointegration vectors, do not eliminate the bias asymptotically.

3/ The Appendix contains a brief description of these single equation methods.

4/ Specifically, estimates of the price coefficients using P-L were less precise than either S-W estimates or OLS estimates.

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 very short-lived subsidy would not change investment plans, and thus not have long run effects.

The perception of temporariness of the subsidy might be due to a to law that states the life span of the subsidy, as in the case of Costa Rica's 1984 export contract, but this is not necessary. This perception can also be due to the expectation of medium-term changes in trade policies, such as joining the GATT. If the fiscal deficit is an issue, then the subsidy might come under attack because of its fiscal impact. Note that as the fob 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 upon the effectiveness to promote exports. 1/

The cointegration equation can be expressed as follows:

$$qx = \beta_0 + \beta_1 \log(1+S) + \beta_2 e + \beta_3 (px^* - pd) + \beta_4 q + w_t \tag{5}$$

The analysis of Costa Rica's export supply response to export subsidies will require testing several hypothesis regarding the price coefficients: β_1, β_2, and β_3. The model suggests that all β_i will be equal. It is also conceivable that β_1 differ when exporters discount the export subsidy relative to EPx^*/Pd. 2/

If all three β's were found to be equal this would imply that exports respond equally to all three price components. This would suggest the export subsidy was not viewed as temporary, as that would have implied a weak long run response by exports. Since these 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.

1/ It is also possible that the exporters might discount the nominal exchange rate, if they perceive the authorities are not committed to keep the exchange at market clearing levels.

2/ In addition, the hypothesis $\beta_2 = \beta_3$ is 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 what Px^*, as 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 consists of actual Pd and E, it is conceivable that these series reflect imperfectly expectations regarding these variables.

Table 1. Export Supply Static Estimation
(t-statistics in parenthesis)

Dependent Variable	qx (1)	qx (2)	qx (3)
Observations	80	80	80
R**2	0.940	0.932	0.940
R-BAR**2	0.937	0.930	0.937
SSR	1.112	1.262	1.117
SEE	0.122	0.128	0.121
DW	1.164	0.970	1.144
Q	72.393	83.370	73.602
ADF	-3.14	-2.72	-3.09
APP	-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-stat	3.504	---	---
H_1, χ^2_1-stat	0.116	---	---

The cointegration equation (5) was estimated using S-W, using 80 quarterly observations covering 1970 through 1989. We have allowed the coefficients of (1+S), E, and Px*/Pd to differ. 1/ The results of the estimation of the static model are presented in Table 1 where column (1) contains the unconstrained estimation, while columns (2) and (3) present two different constrained estimations described below. 2/

1/ We have tested that the coefficients of px* and pd were equal and of opposite signs. The data did not reject this hypothesis. Restrictions will be placed upon price coefficients when supported by the data.

2/ These estimates are subject to two qualifications. First, the estimates suffer from aggregation bias, as the measure of non-traditional exports includes maquila exports that do not qualify for the subsidy. However, this bias is probably small as these products have been growing at a steady rate, reaching about 9.5% of non-traditional exports in 1989. For a discussion of the aggregation bias see Goldstein and Khan (1985). Second, since the subsidy is redeem 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.

Before we discuss the estimates, let us first consider the issue of cointegration. Since all our variables are integrated of order one (denoted as I(1)) it is crucial to determine if a linear combination of them is stationary. The basic intuition is that when w_t is stationary it will revert to its mean value, zero, thus the long-run relationship between exports, relative price and Q has empirical meaning. On the contrary if w_t is I(1) it will diverge indefinitely, hardly ever crossing zero, thus the equation (5) lacks empirical validity. In a nutshell, the model used to derive (5) is incorrect if the variables do not cointegrate.

Cointegration implies that in the static regression w_t does not have a unit root. This has led to tests for cointegration based upon the residuals of the cointegration regression. 1/ The same tests for unit roots described before can be applied to the residuals. It should be noted that while the tests are the same, the significance levels are not. In general, these will depend upon the number of integrated regressors, as well as on trends or other seasonal variables included in the cointegration regression. Engle-Yoo (1987) provide critical values for ADF tests for up to five regressors.

We have performed two cointegration tests, namely ADF and APP. 2/ Notice that ADF tests fail to reject the presence of a unit root at 10 percent significance level. That is, according to this test the equations do not cointegrate. However APP rejects a unit root at 1 percent significance level implying that the equation does cointegrate. The failure of ADF test to reject non-cointegration is likely due to the fact that this test was developed for the case where disturbances are independent, and identically distributed (iid). 3/ Our interpretation of these test results is that the equations do cointegrate, and the presence of non-iid disturbances have adversely affected the power of the ADF test.

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

1/ Alternative tests for cointegration based upon the significance of spurious regressors, as well as tests based upon the long run covariance matrix have been developed in addition to Johansen's likelihood ratio procedure. For details see Campbell-Perron (1991) and the references therein.

2/ Four lags were included in the tests, as indicated by Campbell-Perron's suggestion.

3/ Campbell and Perron (1991) discusses this issue.

Columns (2) and (3) present the constrained regression results under H_0 and H_1 respectively. 1/ 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 supply of exports will remain 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-a-vis export prices, i.e. 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 price elasticities are comparable to those obtained in our unconstrained regressions. It also interesting to note that once again cointegration is attained at 1 percent significance.

This leaves us with a dilemma. While our hypothesis tests suggest that price coefficients are equal, imposing this on our data renders exports perfectly price inelastic. However, when we impose equality between the subsidy and the exchange rate, estimates are more sensible, i.e. one obtains a small significant price elasticity, and stronger evidence of cointegration. Also notice that the standard error of the estimates, SEE, of the regressions in column (3) are smaller than both those from 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 the fact that these Wald tests are asymptotically χ^2 and therefore might not perform adequately in finite samples. Monte Carlo experiments reported by P-H suggest that the probability distribution function is adequately approximated for samples size as small as 50 observations. Nonetheless, as Campbell and Perron (1991) note, these results have been obtained for small scale models with only two or three variables in the cointegrating vector. It is unknown whether these simulation results hold when the model is larger. In the rest of the paper, we will refer to the estimates obtained in column (3) of Table 1, as they seem sensible and are not rejected by the data.

1/ 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, we performed the standard F-tests on OLS estimates of the cointegrating equation. These tests reject H_0, but maintain H_1.

The above results suggest that exporters seemed to 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. It is sensible for a temporary change in policy to have a smaller long-run impact than permanent changes. However, it is hard to generalize this result about export subsidies as estimates of the impact of export subsidies are scarce.

We are aware of only one such study. Balassa et al (1986) study the Greek and South Korean export incentives. However, they report their estimates, breaking up the relative price faced by exporters only for South Korea. Nonetheless, South Korea is a very important case to examine, as it is part of the so called Asian-miracle. Balassa et al. estimate the standard Goldstein-Khan export supply curve, using annual data from 1965-1979. They find that the elasticity of exports to (1+S)E differed and exceeded that of Px^*/Pd, which is precisely the opposite of our result. 1/ Exporters, they explain, 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, however export incentives did not. In fact, export incentives increased up until 1971, when they reached 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 permanent depreciation of the exchange rate, and that could explain the size large elasticity with respect to (1+S)E.

In any event, exporters seem to perceive differently the origin of their export revenue. It would seem important that policymakers keep in mind, when evaluating the effects of reducing export subsidies the perception that exporters entertain. Specifically, we have found that the elasticity with respect to (1+S)E to be less than that of Px^*/Pd. This is important for policy decisions: using the elasticity of Px^*/Pd to evaluate the effect upon exports of a reduction of export subsidies would tend to overstate, in the case of Costa Rica (or understate for South Korea), the negative impact upon export revenues.

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

1/ It should be noted that Balassa et al.'s subsidy is not direct export subsidies, as in these estimates. Rather Balassa et al. construct an implicit indirect subsidization consisting of tax exemptions and other indirect subsidies.

The data, however suggests that the regressors in the cointegrating equation (5) are exogenous. 1/ This result is not trivial, as it implies that both Px^* and Pd are exogenous. This result is partly expected given the size of Costa Rica's exports relative to the size of the major export market U.S.A. However, Pd was more likely to be endogenous. Nonetheless, the exogeneity of Pd is explained by the fact that the market for domestic goods is formed by a large number of suppliers, including some exporters. The data supports the idea that Pd is determined by the actions of the exclusively domestic producers, while exporters take Pd as given. These results are important as they allow us to concentrate exclusively on export supply, disregarding demand. 2/

Before examining the short run dynamics let us refer to the export elasticity of non-traditional 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 increases more than proportionally. This of course is not possible forever. Eventually, all or most output will be exported, and an increase of composite output should translate approximately to a one to one increase of 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 increase more than proportionally. However, we do expect this long run elasticity to fall closer to one, as predicted by equation (4), when firms export a larger portion of their output.

b. Estimation of an error-correction model

Given our estimates of the long relationship obtained in the previous Section we wish to estimate the short run dynamics associated with them. Since our long run relation cointegrates, the Granger representation theorem tells us that the short run dynamics can be expressed by an error correction mechanism (ECM) of the following form:

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

where y_t is the endogenous variable, x_t corresponds to a vector containing exogenous regressors, β is the vector of unknown coefficients and h(L) is a lag polynomial.

1/ The evidence stems from the fact that the leds of the regressors are insignificant, either for the S-W procedure presented in the text or for the P-L. This is analogous to the work of Sims (1972) testing for causality. Additional evidence supporting the exogeneity of regressors comes from the standard Hausman specification test preformed on the P-H estimates, which is also unable to reject the exogeneity of the regressors. See the appendix for a brief description of these estimators.

2/ 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 suggests that this has not yet occurred.

Table 2 contains the estimation results. The first column of Table 2 shows the estimation results for the unconstrained ECM. The specification for the ECM's presented have been arrived at after testing four lags of the difference of each variable in our cointegrating vector. Using standard F-tests, none of the lags are significant, thus they have been excluded from the estimates. The second column presents the constrained ECM, using the cointegrating vector from column (3) in Table 1.

Table 2. Error Correction Model
(t-statistics in parenthesis)

Dependent Variable	Δqx (1)	Δqx (2)
Observations	79	79
R**2	0.251	0.244
R-BAR**2	0.211	0.234
SSR	0.792	0.800
SEE	0.104	0.102
DW	1.972	1.920
Q	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)	---

The appropriateness of the error correction representation for the model can be checked by the unconstrained ECM. Notice that the estimates for the unconstrained ECM have appropriate signs and sizes, but not significant. They imply that the long run elasticity of log[(1+S)E] is about half that of Px*/Pd, about 0.07. While the elasticity with respect to composite output is about 2.27.

The constrained ECM 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). 1/ Notice that imposing the ECM restriction reduces, slightly, the standard error of the estimate. This suggests the efficiency gain obtained by imposing the restriction to our data.

1/ It should be noted that the constrained ECM imposes the same speed of adjustment for all variables, while the unconstrained ECM allows for speed of adjustment to vary.

This final models is subjected to a series of diagnostic tests. Godfrey (1978) and Breusch (1978) generalization of Durbin's h test has been used to test for serial correlation of up to order one and up to order four. We do not find, serial correlation. We do not find autoregressive conditional heteroskedasticity (ARCH) effects either. The residuals from our regression do not show significant skewness or kurtosis.

IV. Effect of the Export Subsidy

The export contract is the cornerstone of export promotion policy in Costa Rica. It was established during 1984:II and governs all export incentives, including the direct export subsidy, CATS. The 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 suggests 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. An important policy implication of the program emerges: 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 tradeoff of reducing the export subsidy and compensating with a higher rate of depreciation.

1. Forecasting performance

Before the model is used to simulate the effect of the export subsidy, its forecasting performance is gauged. To establish the models ability to track the data during this period we have used the models to generate static forecasts of dollar exports. Roughly two-thirds of the one period forecasts errors were less than 10 percent the dollar value of exports. The remaining 8 errors, 5 were less than 15 percent. The models' ability to forecast exports can be measured through dynamic forecasts. To this effect the model is simulated dynamically starting from 1984:II through 1989:IV. 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 forecasts errors are under 10 percent; the other half is 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 models' ability to forecast exports, we have compiled a series of statistics that summarize in-sample forecasts during the export contract. 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 (M.E.), the mean absolute value error (M.A.E.), the root mean square error (R.M.S.E.) and Theil's U statistic. Table 3 presents the results.

Table 3. Forecasting Statistics
(QII 1984- QIV 1989)

Steps	M.E.	M.A.E.	R.M.S.E	Theil U	Obs.
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

Comparing the magnitudes of the M.E. and the M.A.E. gives us an idea of the randomness of the forecasting error. A model that consistently errs in one direction can be usually improved by changing its specification. When a model consistently overpredicts the data, we expect the M.E. to be negative and roughly the same magnitude as the absolute value of the M.A.E. When it under-predicts the data the M.E. is positive and roughly the same magnitude of the M.A.E.

The results do not indicate a problem of consistently over or under predicting the data, as the M.E. and the M.A.E. have very different magnitudes. Notice that our models' one step forecast erred by an average of $4.0 million, while the absolute forecast erred by $9.5 million. Considering that export revenue averaged $117 million during this period, these are quite small. Notice, however, that the model tends to under-predict actual exports.

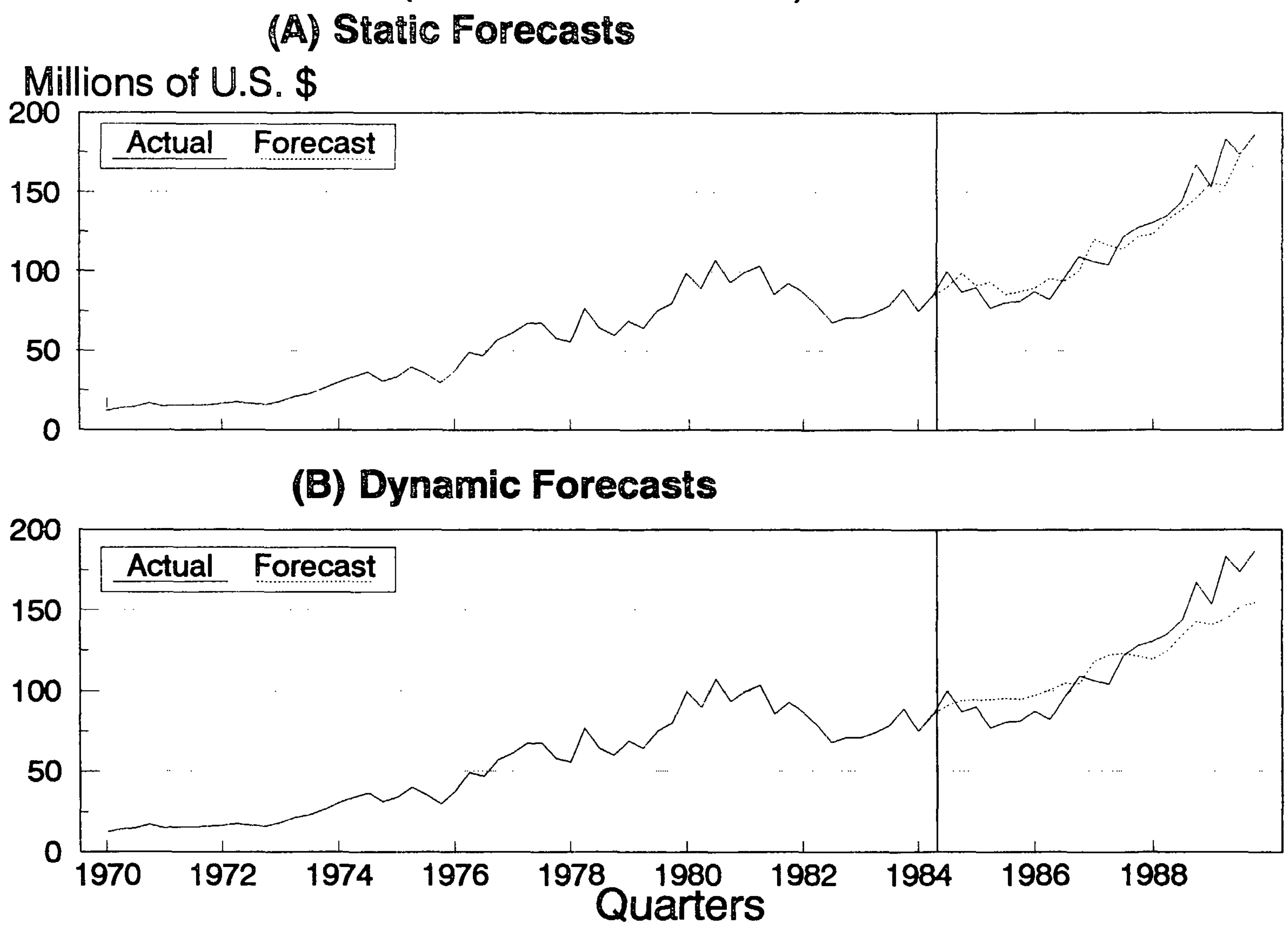
The Theil U statistics for forecasts for three quarters and less are not good. They indicate that a naive forecast of no change outperforms the model. However, as the forecasting horizon increases, the model consistently does better than the naive forecast.

These simulations suggest that the model can track and forecast dollar exports of Costa Rica with reasonable accuracy during the period of interest. One step errors are relatively small, while dynamic forecast errors are larger, they remain reasonable.

2. Simulations

Having evaluated the model's ability to forecast, this sub-section evaluates the role of the subsidy in stimulating exports. Two issues are addressed. First, the model will be used to simulate baseline exports, that are compared with a simulated counterfactual where export subsidy is set to

Figure 2: Model Predictive Capacity
(QII 1984 - QIV 1989)



zero during the export contract. 1/ The additional export revenues will be compared with the budgetary 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.

a. The cost of the subsidy

The model is used to evaluate the impact of the export subsidies during the export contract period, 1984:II-1989:IV. The baseline is obtained by dynamically simulating the model starting from 1984:I. Then, 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 the impact on dollar exports to be approximately $275 *million over these 23 quarters. Given that total non-traditional exports* totalled 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 contains the relevant data.

Table 4. Simulation of the Export Subsidy
(Millions)

Year	Exchange Rate	CATS		Export Response
		Colones	Dollars	
1984	44.98	480.30	10.68	9.25
1985	51.31	973.50	18.97	30.98
1986	56.71	1553.80	27.40	44.60
1987	64.15	2030.50	31.65	54.34
1988	76.84	3880.20	50.50	62.64
1989	82.09	5394.90	65.72	74.73
Total		14473.30	204.92	276.55

1/ At this junction it is important to reference the Lucas critique of policy evaluation. There has been a growing recognition that policy evaluation is not useless. Both Cooley, Leroy and Rahman (1984) and Sims (1982) have argued that the usual interpretation of the critique is logically flawed. Sims (1987) argues that the Lucas 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 expectation formation behavior implicit in the model's structure." It is in this context that policy simulations are conducted later in this Section.

The direct cost of the subsidization program is estimated at roughly $205 million. 1/ This corresponds to an average of 0.8 percent of GDP over these 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, it seems that a dollar spent on export subsidies has yielded about 34 percent gain 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. There are important administrative costs associated with the program. Each contract is reviewed by the Ministry of Finance to determine eligibility; the most important requirement is that of 35 percent domestic value added. The Central Bank receives the paper work, and keeps track of the contracts. For each and every shipment, the Central Bank determines the appropriate subsidy and issues the tax credit papers. These tax credits are then submitted to the Ministry of Finance when firms file their taxes. 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. However, these exports have a significant import component. On average, non-traditional exports contain roughly 40 percent of domestic value added. 2/ This means that only $110 million has been generated net of imports, over the 23 quarters. Using the lower bound estimate for cost to determine the yield of the subsidy program, renders a net generation of 54 cents for each dollar spent. What this implies is that out of each dollar transferred from tax payers to exporters, via the export subsidy, 46 cents ended up subsidizing the imports of intermediate inputs. 3/

A more efficient way to transfer resources to exporters would be via a direct transfer; the same incentive could have been accomplished with roughly half the tax resources. If, for example, firms face some sort of barrier to start up their exporting business, the transfer could be administered as a lump-sum to cover initial investment or initial cost of penetrating foreign

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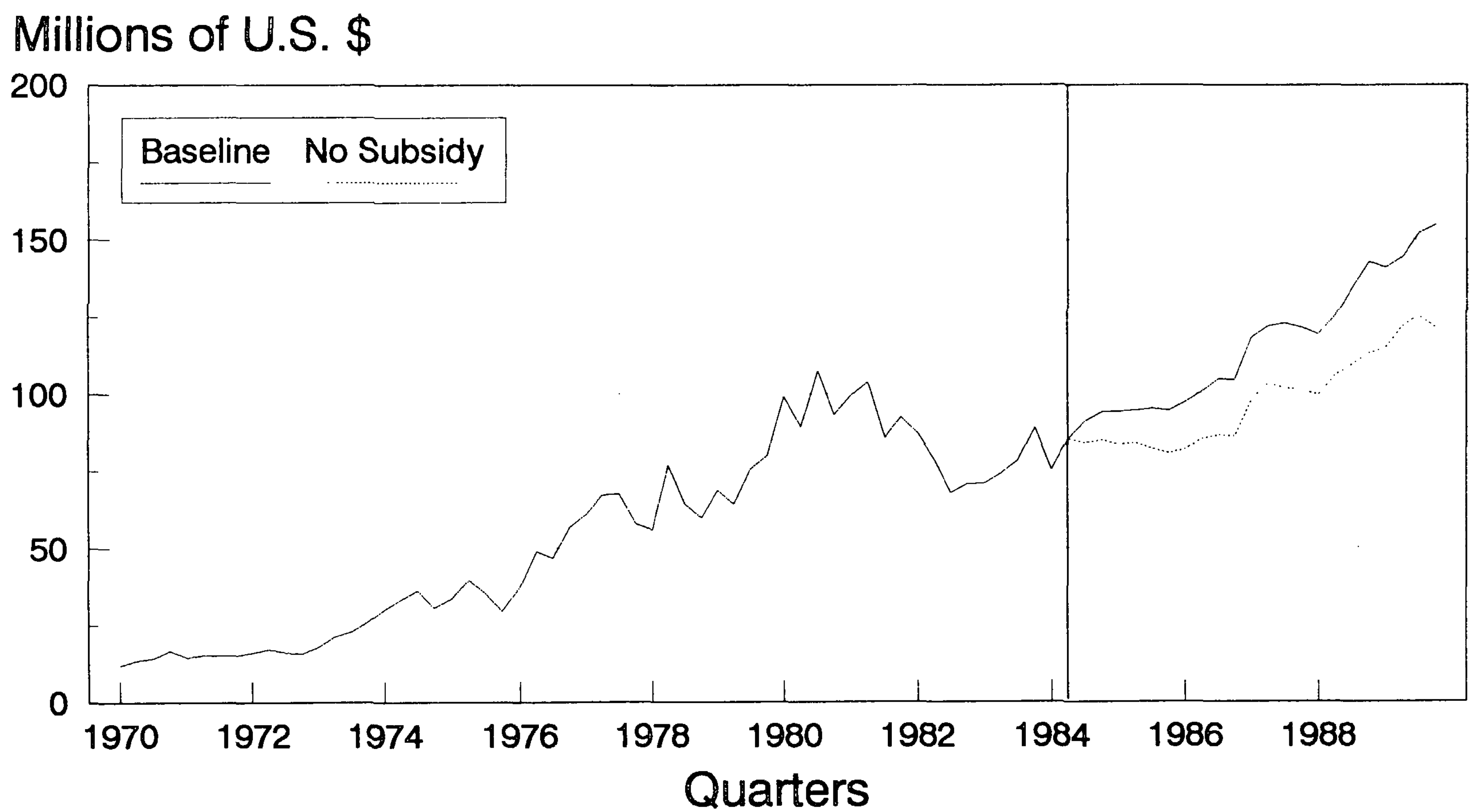
1/ The export subsidy, CATS, has been converted into dollars using the average exchange rate.

2/ The export contract calculation of domestic value added, is obtained by adding 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 was used to calculate an average value for value added.

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

Figure 3: Export Subsidies

Simulation QII 1984 - QIV 1989



markets. Alternatively, the subsidy could be re-designed so as to apply to the domestic value added of the exports. This could provide the same incentive as a subsidy on the value of exports, at a fraction of the cost. 1/

b. Exchange rate depreciation

A policy frequently mentioned to compensate the reduction of export subsidies is exchange rate depreciation. As discussed above, 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, thus $\hat{E}$ equals minus the percentage change of (1+S). The exact trade off between the exchange rate and the subsidy is simple to calculate. The estimates have been obtained using an index, "S", for the export subsidy: S'_t/S'_0. Notice that the percentage change of (1+S) 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 the total dollar exports during these six years was the as the baseline, roughly $2.7 billion. Compensating requires increasing the quarterly depreciation by 7 percent. 2/ Figure 4 depicts the trajectory of exports compensated with an increase of 7 percent over the baseline.

Notice that during the first two years the level of exports under the constant 10 percent depreciation is less than the baseline, while the last two years they are 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. These 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, that could be compensated by an increase of the quarterly rate of depreciation of about 1.75 percent.

A final comment should be made about the compensating depreciation. It is possible that the compensating depreciation will affect the domestic price of exportables, reducing the impact of the depreciation on exports. The substitution of the subsidy for a higher rate of depreciation has offsetting effects upon inflation. A higher rate of depreciation will tend to increase the domestic cost of imported goods, and thus can contribute to higher prices.

1/ Unfortunately, the model cannot measure the reduction of the cost for this subsidy.

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

This 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 tend to 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 requires a complete macro model of the Costa Rican economy to measure. Our calculation of the compensating depreciation assumes that the effects upon inflation offset each other. Thus, our calculations refer to a concept analogous to real depreciation. 1/

V. Conclusions

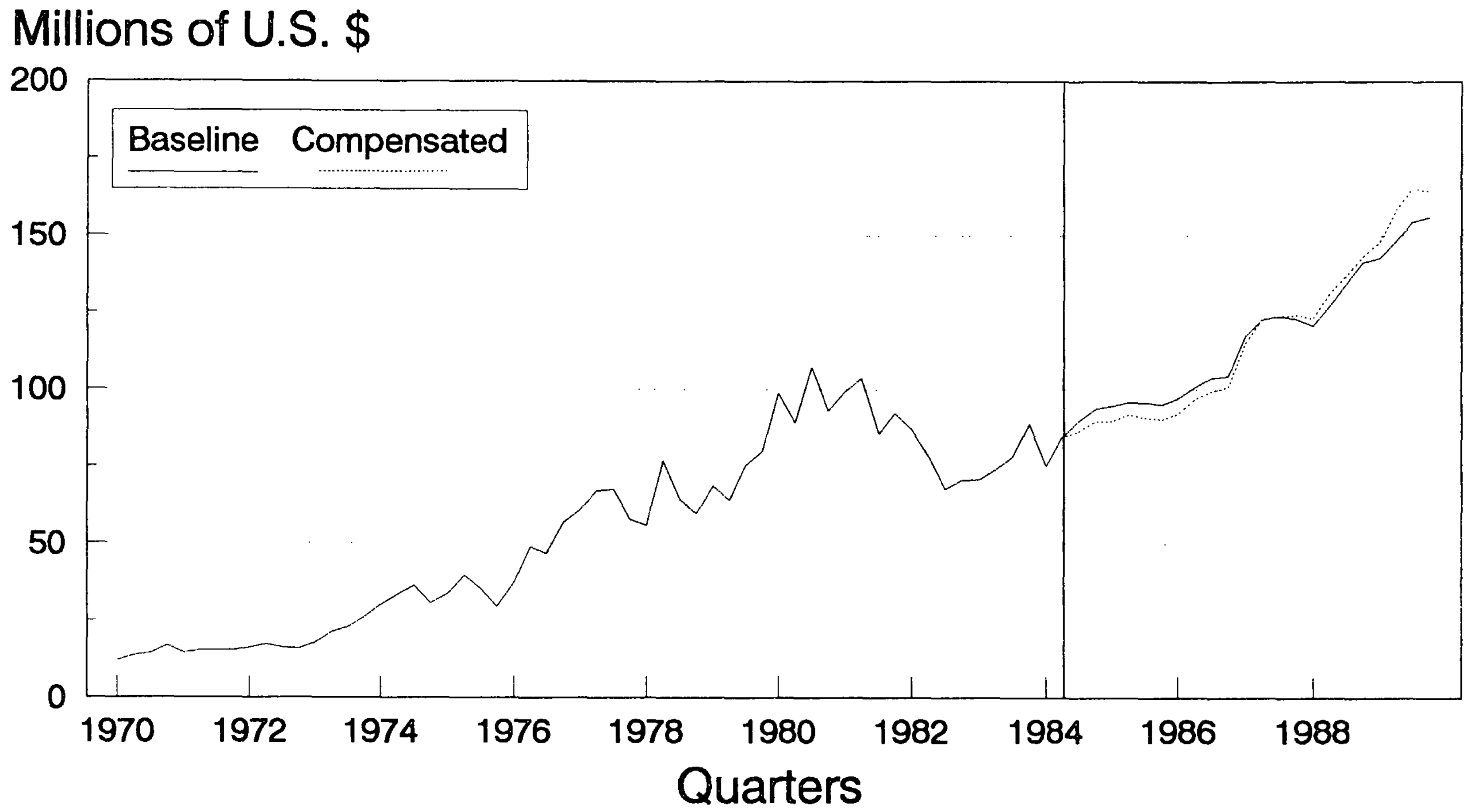
In recent years, many countries have switched their development strategies from import substitution to export promotion. Empirical evidence regarding the effectiveness and costs of these export promotion policies and specifically direct and indirect incentives to exports, is limited. To analyze these issues we have modelled a firm that faces two markets (domestic and world) and introduced a direct export subsidy. The subsidy tends to increase output and switch sales to the world market.

With regards to the model, it is worth noting the following characteristics. The model analyzes the long-run supply of exports and have relied on the data to generate the short run dynamics. 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 allow us to shed light upon the dynamics of export subsidies. It is likely that subsidies could trigger both intertemporal and intra-temporal responses through the effect upon investment decisions. This model would be analogous to models that have analyzed the effect of terms of trade shock upon the trade balance, for example Ostry (1988). Indeed, it is quite likely that export subsidies would have very different effects when they are viewed as temporary as opposed to being permanent. However, this paper does not elaborated upon this. Our strategy has been to use the theory to determine the long run determinants and let the data determine the short run dynamics. Nonetheless, modeling the dynamics could be fruitful avenue for future research.

1/ The caveat on real depreciation is due to the asymmetry between the effect of pd and the subsidy. Our estimates suggests that export supply is more sensitive to domestic prices they face, than to the nominal exchange rate. The calculated effect upon 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 to the rate of depreciation accounting for inflation.

Figure 4: Compensating Depreciation

Simulation: QII 1984 - QIV 1989



The estimates of the long-run relationship between export supply and relative prices for Costa Rica show strong evidence of cointegration. This allowed the estimation a constrained error correction model, to capture the short run dynamics of export supply. The estimates suggest that exports are price inelastic, and firms adjust within the year to shocks in the system. The forecasting performance of the estimated model is 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 is much smaller, estimated to be only by $110 million. The direct cost of the subsidy, not accounting for administrative costs involved, totalled about $250 million.

The cost of the subsidy has averaged 1.2 percent of GDP during 1988 and 1989. This has prompted 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 could be to subsidize the domestic value added of exports. This would reduce the cost of the incentive by avoiding the subsidization of imports. Alternatively, a lump-sum transfer could also avoid subsidizing imports. This could be set up to cover initial investment cost or the initial costs to penetrate foreign markets.

Compensating depreciation is common prescription to 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 that we do not address in this paper. The purpose of the estimates and simulations is to provide some evidence regarding export subsidies to help quantify its impact and suggest possible trade offs. No claim is made with regards to the optimality of the export subsidy or of export promotion in general.

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 upon production and consumption decisions, the economically preferable policy is to eliminate source of the anti-export bias. Thus, the first best policy is trade liberalization.

Data

The following quarterly series have been taken from IFS: (1) the exchange rate, (2) domestic price and (3) Px* These have been defined as codes: ahx, 63 country 238 (Costa Rica) and 74 i.r country 111 (U.S.A.) respectively. 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 export of nontraditional exports, and (2) price of nontraditional exports. The former series has been distributed, using Litterman [1984]; the related series was the category of "Other Exports" provided by BCCR. The latter series was distributed using Chow and Lin as mentioned above. The dollar exports has been deflated using the price of nontraditional exports to obtain the quantity of exports.

The Ministry of Finance provided the CATS subsidy series. An annual series for CATS entregados was distributed using Chow and Lin 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 regarding the domestic value added of nontraditional exports.

Quarterly GDP figures have been taken from Hoffmaister (1991). All relevant series have been indexed to 1985.

Unit Root Test Results

Table A1. Unit Root Tests

	Series			
Test	qx	px-pd	log(1+S)e	q
ADF:				
Level	-1.440	-1.827	-0.259	-1.487
(1-L)	-34.298**	-2.108	-2.050	-1.833
APP:				
Level	-2.522	-1.609	-2.095	-2.200
(1-L)	-74.770**	-31.450**	-23.788**	-600.153**

Single Equation Methods

To discuss the three methods mentioned above, let us introduce the following equation:

$$y_t = \beta' x_t + \mu_t^{(1)} \tag{A1}$$

$$\Delta x_t = \mu_t^{(2)} \tag{A2}$$

Equation (A1) is the cointegrating equation and (A2) is a vector of k regressors included in (A1). Let $\mu = [\mu^{(1)}, \mu^{(2)}]'$ be the (k+1) vector of residuals in the system A1-A2 and let its covariance matrix be:

$$\Sigma = E[\mu \cdot \mu'] = \begin{bmatrix} \sigma_{11} & \sigma'_{21} \\ \sigma_{21} & \Sigma_{22} \end{bmatrix}$$

partitioned to conform with A1-A2.

P-H note that for time series:

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

Their non-parametric correction for serial correlation adjusts OLS estimates obtained from (8a) by adding: $-[x'x]^{-1} T \hat{\sigma}_{21}$, where $\hat{\sigma}_{21}$ is a consistent estimator of σ_{21}. This adjustment purges the OLS estimates from the nuisance parameters due to serial correlation.

Their "fully modified" estimator requires two corrections that are accomplished as follows. First the LHS variable in (8a) is purged from endogeneity by the following transformation: $y_t^* = y_t - \hat{\sigma}'_{21} \hat{\Sigma}_{22}^{-1} \Delta x_t$. OLS is performed with this transformed variable, and in turn corrected for serial correlation by adding to it $-[x'x]^{-1} T \hat{\delta}$, where $\hat{\delta} = \hat{\Phi} \cdot [1, -\hat{\Sigma}_{22}^{-1} \hat{\sigma}_{21}]'$ and $\hat{\Phi}$ is a consistent estimate of $\Phi = \sum_{s=0}^{\infty} E[\mu_0^{(2)} \mu'_s]$.

S-W 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 $E[\mu_t^{(1)} / \{\Delta x_t\}] = E[\mu_t^{(1)} / \{\mu_t^{(2)}\}] = d_1(L) \Delta x_t$, where $d_1(L)$ is a two sided lag polynomial. It should be noted that $d_1(L) = \sum_{i=-\infty}^{\infty} d_{1,i} \cdot L^i$ in practice is truncated. By adding and subtracting this term to (A1):

$$y_t = \beta' x_t + d_1(L) \Delta x_t + c_{22}(L) \tilde{\mu}_t^{(2)} \tag{A3}$$

where $\tilde{\mu}^{(2)} = \mu^{(2)} - E[\mu^{(2)} / \{\mu_t^{(2)}\}]$ is independent of innovations from the LHS variable by construction. S-W suggest using OLS on the dynamic equation (A3). They call this estimator dynamic OLS (DOLS).

This parametric correction for endogeneity is shared by P-L, and is based upon the work of Sims (1972) on causality tests. Recall that when a variable y_t Granger-causes x_t, then y_t can be expressed as a linear combination of past, future and present values of x_t. Thus future values of x_t will provide information that helps in the prediction of y_t. These future values of x_t are in essence Sims' causality test. Significant values for future x_t, provides evidence that x_t is not weakly exogenous.

Equation (A3) still contains serial correlation. S-W propose to deal with the serial correlation by correcting the covariance matrix used in the estimation of (A3). The covariance matrix should be estimated using non-parametric 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 processses, suggesting dynamic GLS. P-H suggest a parametric correction to deal with serial correlation. They propose to add to equation (A3) the term $d_2(L)(y_t-\beta' x_t)$, where $d_2(L)$ is a one sided lag polynomial defined as $\sum_{i=1}^{\infty} d_{2,i} \cdot L^i$.

Their estimator implies that the cointegrating vector enters nonlinearly, thus it is estimated using nonlinear least squares.

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