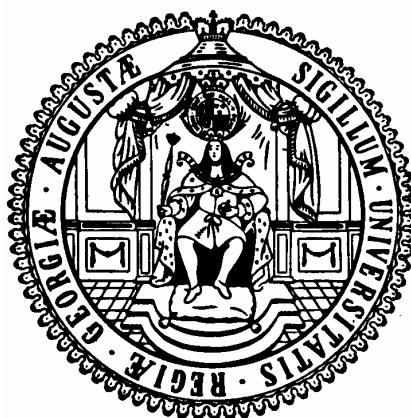


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**Export-Led Growth in Chile:
Assessing the Role of Export**

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Export-Led Growth in Chile: Assessing the Role of Export Composition in Productivity Growth

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This study examines the export-led growth hypothesis using annual time series data from Chile in a production function framework. It addresses the limitations of the existing literature and focuses on the impact of manufactured and primary exports on productivity growth. In order to investigate if and how manufactured and primary exports affect economic growth via increases in productivity, several single-equation and system cointegration techniques are applied. The estimation results can be interpreted as evidence of productivity-enhancing effects of manufactured exports and of productivity-limiting effects of primary exports.

I. INTRODUCTION

The export-led growth (ELG) hypothesis has been the subject of considerable empirical research, though with mixed and questionable results. Earlier studies which use cross-country data can be

criticised for taking positive correlations as evidence of causation without testing for the direction of causality.¹ A statistically significant positive relationship between exports and growth often found in cross-country studies, admittedly, need not necessarily be the result of an impact of exports on economic growth. A positive correlation or coefficient of exports in the growth equation can equally be compatible with causality running from growth to exports [*Abu-Quarn and Abu Bader 2004*]. However, the main criticism directed at cross-country studies is that they implicitly assume a common economic structure and similar production technologies across countries. Significant parametric variations between different countries may therefore lead to highly misleading results [*Shan and Sun 1998*].²

In response to these criticisms, more recent econometric studies use time-series data from individual countries to investigate the causal relationship between exports and growth by means of Granger-type causality tests. The evidence derived from these tests is mixed and often conflicting.³ But the wide variations in empirical results can be attributed to the fact that causality tests are extremely sensitive to omitted variables. Even if exports are found not to cause growth in bivariate models, this same inference does not necessarily hold in the context of larger economic models that include other relevant variables such as capital and labour [*Awokuse 2003*]. Indeed, numerous studies estimate an export-augmented production function, but in many cases they fail to incorporate imports along with exports in their production function estimates.⁴ According to Riezman et al. [1996], omitting the import variable can result in spurious conclusions regarding the ELG hypothesis, because capital goods imports are inputs for export and domestic production. Furthermore, export growth may relieve the foreign exchange constraint, allowing capital goods to be imported to boost economic growth.

Additional problems arise because exports, via the national income accounting identity, are themselves a component of gross domestic product [GDP].⁵ Accordingly, exports are partly endogenous within an output equation. The outcome of this is a strong bias in favour of a

correlation between these two variables, whatever actual causal relationship may exist between them [*Greenaway and Sapsford 1994*]. Finally, it should be pointed out that most of the recent time-series literature focus on ‘aggregate’ exports only.⁶ This may mask important differences between different export categories. Even if there is evidence in favour of the ELG hypothesis relating to certain export categories, this may not be reflected at the aggregate level, and spurious conclusions may be drawn when disaggregated exports are not examined [*Ghatak et al. 1997*].

The objective of this paper is to carefully investigate the long-run relationship between exports and growth. It contributes to the existing literature in the following ways: First, because of the above-mentioned limitations of gross country regressions, we apply time series techniques. Second, in order to tackle the possible specification bias, we go beyond the traditional two-variable causality relationship and estimate an export-augmented production function. Third, we test the ELG hypothesis while controlling for capital goods imports in order to capture the role of exports in financing capital goods imports, which in turn are expected to promote growth. Fourth, we separate the ‘economic influence’ of exports on output from that incorporated into the ‘growth accounting relationship’ by defining the output variable net of exports.⁷ Fifth, we decompose aggregate exports into primary and manufactured exports. Sixth, as far as the econometric methodology is concerned, two types of unit root tests are performed on each series. Both types of tests control endogenously for the possibility of structural breaks. Moreover, we use single equation and system equation techniques to test for cointegration and causality as well as to estimate the parameters of our production function. Finally, misspecification and structural stability tests are conducted for the estimated causal long-run relation between exports and growth.

In order to investigate the growth effects of primary and manufactured exports we use Chilean time series data from 1960 – 2001. Chile is chosen as a case study because (i) Chile experienced a pattern of high long-run growth, which, however, was interrupted by three deep economic crises [the collapse of the Allende government in 1973, the 1975 recession, and the 1982

economic crisis], (ii) Chilean exports grew very rapidly after 1974, when trade liberalisation was initiated,⁸ (iii) Chilean exports rely heavily on primary products, although the share of manufacturing exports in goods exports rose from 7 percent in 1973 to 44 percent in 2001, and (iv) Chile is extremely vulnerable to fluctuating commodity prices, especially copper prices,⁹ since copper still accounts for about 37 percent of total exports of goods in 2001.¹⁰ Moreover, up to now no attempt has been made to examine the separate effects of primary and manufacturing exports on Chilean economic growth.

The rest of the paper is organised as follows. Section II discusses the theoretical background of the ELG hypothesis and the empirical model. The data and the econometric methodology are described in Section III. The estimation results are presented in Section IV. A final section summarises the conclusions.

II. THEORETICAL BACKGROUND AND EMPIRICAL MODEL

Theoretical Background

The ELG hypothesis postulates that export expansion is a key factor in promoting long-run economic growth. Several arguments can be put forward to justify the ELG hypothesis theoretically. From a demand-side perspective, it can be argued that sustained demand growth cannot be maintained in small domestic markets, since any economic impulse based on the expansion of domestic demand is bound to be exhausted quickly. Export markets, in contrast, are almost limitless and hence do not involve growth restrictions on the demand side. Thus, exports can be a catalyst for income growth, as a component of aggregate demand [*Agosin 1999*]. This is the direct impact of exports on economic growth which we do not need to verify econometrically here. Given the fact that Chilean exports increased from about 9 percent of GDP in 1960 to about 33

percent in 2001, it immediately becomes clear that exports have played a significant role in the Chilean growth process – as part of demand for Chilean output. In the empirical analysis, we will control for this huge demand-side effect by defining the GDP variable net of exports.

In addition to the direct demand-side effect, export expansion may indirectly affect growth by providing the foreign exchange that allows for increasing levels of capital goods imports [Riezman 1996]. Increasing capital goods imports in turn stimulate output growth by raising the level of capital formation. Furthermore, recent theoretical work suggests that capital goods imports from technologically advanced countries may increase productivity and thereby growth, since knowledge and technology is embodied in equipment and machinery and therefore transferred through international trade [*Chuang 1998*]. We will control for this indirect effect in the empirical analysis by incorporating capital goods imports into the estimating equation because our real focus in the empirical work will be on examining the effects of exports on productivity growth.

In theory there are a several potential ways in which exports can cause an increase in productivity. First, an expansion in exports may promote specialisation in sectors in which a country has comparative advantage, and lead to a reallocation of resources from the relatively inefficient non-trade sector to the more productive export sector. Second, the growth of exports can increase productivity by offering larger economies of scale [*Helpman and Krugman 1985*]. Third, export growth may affect total factor productivity through dynamic spillover effects on the rest of the economy [*Feder 1982*]. The possible sources of these knowledge externalities include productivity enhancements resulting from increased competitiveness, more efficient management styles, better forms of organisation, labour training, and knowledge about technology and international markets [*Chuang 1998*]. In short, knowledge is generated through a systematic learning process initiated by exports and spilling over to the domestic economy. Thus, the ELG hypothesis implies that export growth will lead to economy-wide productivity growth.

However, several authors hypothesise that primary exports are an obstacle to greater productivity growth. The main arguments advanced in support of this hypothesis are: (i) Primary products offer no sustainable potential for knowledge spillovers, and an increase in primary exports can draw resources away from the externality-generating manufacturing sector [*Matsuyama 1992*]. (ii) Primary exports are subject to extreme price and volume fluctuations. Increasing primary exports may therefore lead to increasing GDP variability and macroeconomic uncertainty. High instability and uncertainty may, in turn, hamper efforts at economic planning and reduce the quantity as well as efficiency of investments [*Dawe 1996*]. Consequently, it is assumed that the effects of exports on productivity and growth differ significantly between primary and manufactured products. In the empirical analysis we will examine how these effects differ.

Empirical model

On the basis of the above-mentioned theoretical and methodological arguments, our empirical model starts with a simple neoclassical production function:

$$Y_t = A_t K_t^\alpha L_t^\beta, \quad (1)$$

where Y_t denotes the aggregate production of the economy at time t , and A_t , K_t , L_t are the level of total factor productivity, the capital stock, and the stock of labour, respectively. Because we want to investigate if and how manufactured and primary exports affect economic growth via increases in productivity, we assume that total factor productivity can be expressed as a function of manufactured exports, IX_t , primary exports, PX_t , and other exogenous factors C_t :

$$A_t = f(CM_t, IX_t, PX_t, C_t) = CM_t^\delta IX_t^\gamma PX_t^\rho C_t, \quad (2)$$

where capital goods imports, CM_t , are also considered to offer potential to boost productivity, since they may include technologically sophisticated items. Moreover, omission of this variable can result in spurious conclusions regarding the ELG hypothesis. We combine equation (2) with equation (1) and obtain

$$Y_t = C_t K_t^\alpha L_t^\beta CM_t^\delta IX_t^\gamma PX_t^\rho, \quad (3)$$

where α , β , δ , γ , and ρ are the elasticities of production with respect to K_t , L_t , CM_t , IX_t , and PX_t . Taking natural logs (L) of both sides of equation (3) gives an estimable linear function:

$$LY_t = c + \alpha LK_t + \beta LL_t + \delta LCM_t + \gamma LIX_t + \rho LPX_t + e_t, \quad (4)$$

in which all coefficients are constant elasticities, c is a constant parameter, and e_t is the usual error term, which reflects the influence of all other factors. Accordingly, the estimates of γ and ρ may serve to measure the productivity effects of manufactured exports and primary exports on economic growth. It is problematic, however, that exports – via the national accounting identity – are themselves a component of output. A positive and statistically significant correlation between manufactured exports, primary exports, and aggregate output is therefore almost inevitable, even if there are no productivity effects.¹¹ To remedy this problem, it is necessary to separate the ‘economic influence’ of exports on output from the influence incorporated into the ‘growth accounting relationship’. Following Ghatak et al. [1997], we deal with this issue by using the aggregate output, net of primary and manufactured exports, NY_t ($NY_t = Y_t - IX_t - PX_t$), instead of total output, Y_t . By replacing Y_t with NY_t , we finally obtain equation (5):

$$LNY_t = c + \alpha LK_t + \beta LL_t + \delta LCM_t + \gamma LIX_t + \rho LPX_t + e_t, \quad (5)$$

This equation is estimated to determine the impact of increasing manufactured exports and primary exports on economic growth via increases in productivity. However, when estimating equation (5), we must take into consideration that higher rates of capital formation, labour force growth, increased capital goods imports, and increased manufactured and primary exports may all be consequences of economic growth. This issue of causality will also be addressed in the empirical analysis.

III. DATA AND ECONOMETRIC METHODOLOGY

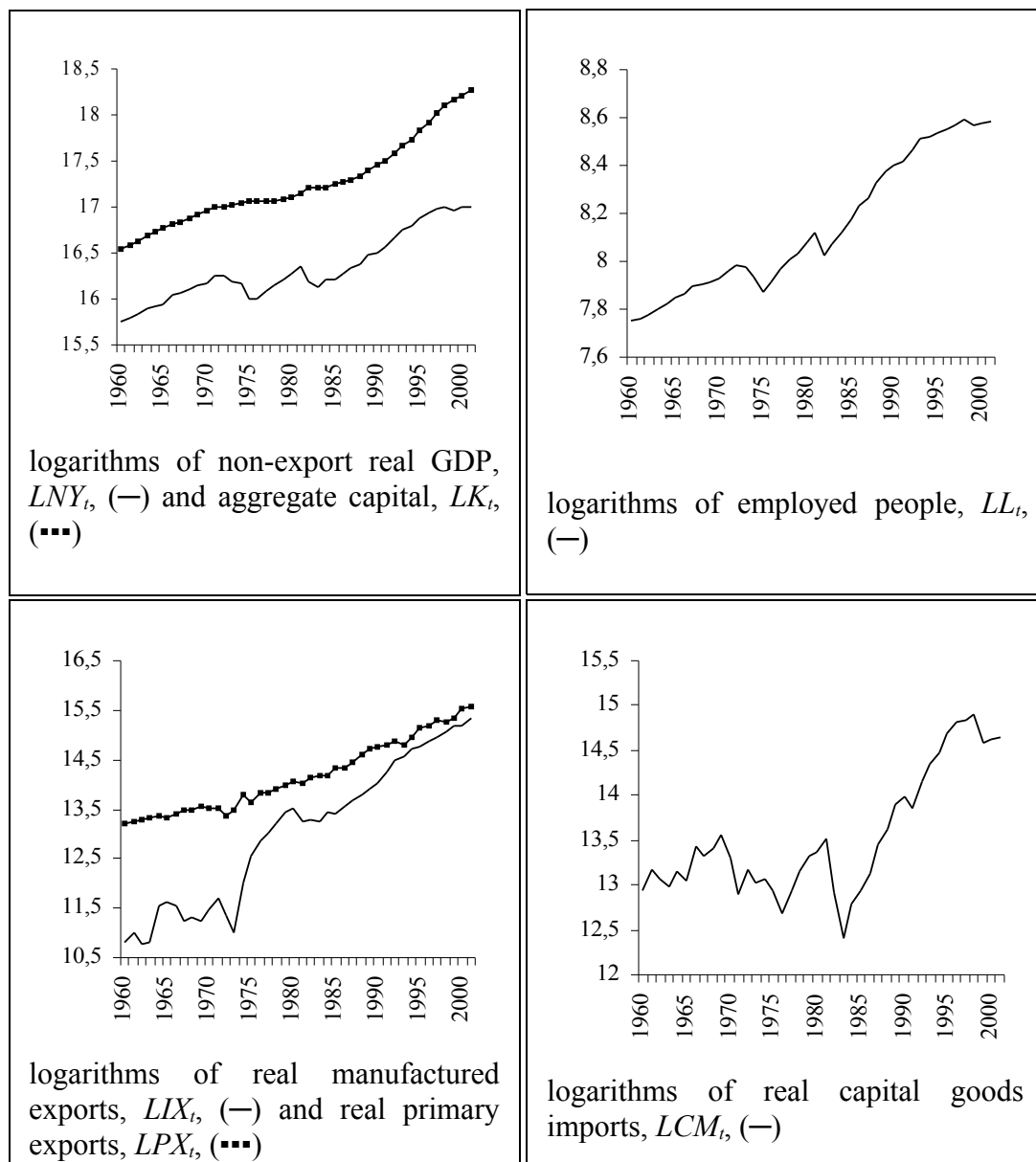
Data

The empirical analysis is based on annual data from 1960 to 2001. They were gathered from the *Indicadores económicos y sociales de Chile 1960-2000* and the *Boletines mensuales* published by the Chilean Central Bank. The variables CM_t , IX_t , and PX_t represent real imports of capital goods, real exports of manufactured goods, and real exports of primary products respectively. The non-export output, NY_t , is measured by real Chilean GDP net of primary and manufactured exports. K_t is the Chilean capital stock in real terms, which was computed on the basis of accumulated capital expenditure using the perpetual inventory method in simple form. Non-export GDP, capital stock, capital goods imports, exports of manufactured products, and primary products are evaluated in Chilean pesos at constant 1996 prices. The labour variable, L_t , is represented by the total number of people employed each year. Figure 1 shows the evolution of the variables in the period under consideration. (All variables are in logarithms).

From Figure 1, it can be inferred that all variables are trending and are thus nonstationary. Nonstationary variables may contain unit roots. Such variables are said to be integrated of order d , $I(d>0)$, because they have to be differenced d times to achieve stationarity [difference stationary

series]. In the case where nonstationary variables are not driven by a unit root process, they are subject to deterministic time trends [trend stationary series]. By removing the deterministic trend, they can be made stationary, $I(0)$. If the variables are $I(0)$, then standard regression methods are applicable. If the variables individually have unit roots, then cointegration analysis is appropriate.

FIGURE 1: TIME SERIES USED



Econometric methodology

In the first step, we test the variables for unit roots to verify their order of integration. It is well known that standard unit root tests are biased in favour of identifying data as integrated if there are structural changes. For all the series there is indeed a strong likelihood that structural discontinuities are present [e.g., the socialist government of President Allende (1970-1973) which pursued a highly inward-oriented economic policy; the 1975 economic crisis; and the deep 1982 recession]. Therefore, we undertake the unit root test developed by Perron [1997]. The Perron procedure permits a formal evaluation of the time series properties in the presence of structural breaks at unknown points in time. It allows the break date to be identified endogenously through the testing procedure itself. A problem might be that the Perron procedure allows only for one possible break point for any single series. To consider the possibility that two break points occurred over the relevant period we apply Kapetanios' [2002] test for the unit root hypothesis against the alternative of trend stationarity with two endogenously determined breaks.

If all variables are found to be $I(1)$, the second step is to test for the existence of a cointegration relationship between them. We apply the standard Engle-Granger [1987] two-step estimation procedure, which involves estimating the static cointegration equation [equation (5)] by OLS and testing the residuals for stationarity. If the residuals are stationary, then the variables are cointegrated. However, the Engle-Granger approach is criticised for several shortcomings, which include the following: (a) the arbitrary normalisation of the cointegrating vector, (b) the assumption of one cointegrating vector in systems with more than two variables and (c) biased OLS estimators.¹² Furthermore, due to non-normality of the distribution of the estimators, no final judgement can be passed on the significance of the estimated coefficients.

Therefore, in the third step, we use the full information maximum likelihood [FIML] cointegration approach developed by Johansen [1995] in addition to the Engle-Granger method. Johansen's system-based procedure treats all variables as potentially endogenous and thus avoids

the problem of normalising the cointegrating vector on one of the variables. Moreover, it allows the empirical determination of the number of cointegrating relations and produces maximum likelihood estimators of the parameters of these relations. These estimators are governed by asymptotic normal distributions, permitting valid statistical inference with conventional test statistics.

After testing for the number of cointegrating vectors and estimating their coefficients, the fourth step is to test for weak exogeneity of the long-run parameters. According to Hall and Milne [1994] a rejection of weak exogeneity implies long-run Granger causality. To detect long-run causality we employ a weak exogeneity testing approach, which has been used similarly, among others, by Lütkepohl and Wolters [1998] and Juselius [2001]. It involves estimating a vector error correction model, reducing the parameter space by imposing zero restrictions on the short-run dynamics and testing the significance of the error correction term.

In the last step, we check the robustness of the cointegration estimates from step three. Since in small sample sizes FIML, estimates are very sensitive to the specification of the statistical model and the choice of the lag length, we additionally apply the Dynamic OLS [DOLS] procedure developed by Saikkonen [1991]. This procedure is asymptotically equivalent to Johansen's maximum likelihood estimator and is known to perform well in small samples. Moreover, DOLS generates unbiased and asymptotically efficient estimates for variables that cointegrate, even with endogenous regressors.

IV. EMPIRICAL ANALYSIS

Time Series Properties

We begin by carrying out unit root tests. Standard unit root tests are not be able to reject the unit root hypothesis if the deterministic trend of a series has a break. The methodology developed by

Perron [1997] can distinguish the unit root hypothesis from that of a trend-stationary series with a single break. In order to test the unit root null hypothesis against the one-break alternative, we estimate two models of the Dickey-Fuller type without any prior knowledge of any potential break dates, i.e.

$$y_{1t} = \mu_1 + \theta_1 DU_t + b_1 t + \delta_1 D(TB)_t + a_1 y_{t-1} + \sum_{i=1}^k c_{1i} \Delta y_{t-1} + e_{1t} \quad (6)$$

$$y_{2t} = \mu_2 + b_2 t + \delta_2 DT_t + \hat{y}_t$$

$$\hat{y}_t = a_2 \hat{y}_{t-1} + \sum_{i=1}^k c_{2i} \Delta \hat{y}_{t-1} + e_{2t} \quad (7)$$

where y_{1t} and y_{2t} are the series of interest, Δ is a difference operator, $TB \in T$ ($T = 42$, $1 \leq t \leq 42$) denotes the time at which the change in the trend function occurs and $DU_t = 1(t > TB)$, $D(TB)_t = 1(t = TB + 1)$, $DT_t = 1(t > TB)(t - TB)$ are indicator dummy variables for the break at time TB .

The regression models (6) and (7) correspond, respectively, to the crash model and the changing growth model proposed by Perron [1989]. Model (6), the *innovational outlier model*, allows for a one-time change in the intercept of the trend function. It involves a one-step regression by estimating the trend function and the dynamics of the process simultaneously. Model (7), the *additive outlier model*, which involves a two-step regression, allows for a change in the slope of the trend function without a change in the level.¹³ For LNY_t , LL_t , LIX_t , LPX_t and LCM_t regression of type (6) is carried out. Regression (7) is applied to LK_t as the capital stock data indicates no ‘crash’ but a change in the slope of the series.

The break point is chosen by estimating the models for each possible break date in the data set, and TB is selected as the value which minimises the t -statistics for testing $a_1 = 1$ and $a_2 = 1$: $t_a^*(i) = \text{Min}_{TB} t_a(i, TB, k)$, where $t_a(i, TB, k)$ is the t -statistic for testing $a = 1$ under model $i = 1, 2$ [model (6) and (7)] with a break date TB and truncation lag parameter k . If $\text{Min}_{TB} t_a(i, TB, k)$ exceeds

(in absolute value) the critical value reported by Perron [1997], the hypothesis of difference stationarity and a unit root is rejected.

Since considerable evidence exists that data-dependent methods of selecting the value of the truncation lag k are superior to choosing a fixed k a priori, we follow Perron [1997] and use the t -sig method. Here, k max is specified to be four. If the last included lag is insignificant, the number of lags is reduced by one and the equation is reestimated until a significant lagged dependent variable is found. If none of the coefficients on the lagged variables are found to be significant (at the 10% level), no lags are utilised in the test. Table 1 contains the results of the sequential unit root tests for the variables in levels and in first differences. The results indicate that LNY_t , LK_t , LL_t , LCM_t , LIX_t and LPX_t are integrated of order one.

TABLE 1
PERRON [1997] UNIT ROOT TEST

Series	Model	Dummy Variables	Test Statistic $t_{\hat{\alpha}}$	Critical Value 5% (1%)	Result
Levels					
LNY_t	(6)	$DU74, D74$	-2.91	-5.23 (-5.92)	$I(1)$
LK_t	(7)	$DT82$	-2.48	-4.83 (-5.45)	$I(1)$
LL_t	(6)	$DU82, D82$	-3.70	-5.23 (-5.92)	$I(1)$
LCM_t	(6)	$DU70, D70$	-2.77	-5.23 (-5.92)	$I(1)$
LIX_t	(6)	$DU73, D73$	-5.16	-5.23 (-5.92)	$I(1)$
LPX_t	(6)	$DU72, D72$	-2.76	-5.23 (-5.92)	$I(1)$
First Differences					
$\Delta(LNY_t)$	(6)	$D74$	-4.44	-3.53 (-4.21)	$I(0)$
$\Delta(LK_t)$	(7)	$D82$	-6.61	-1.95 (-2.63)	$I(0)$
$\Delta(LL_t)$	(6)	$D82$	-4.57	-3.53 (-4.21)	$I(0)$
$\Delta(LCM_t)$	(6)	$D70$	-5.82	-3.53 (-4.21)	$I(0)$
$\Delta(LIX_t)$	(6)	$D73$	-5.79	-3.53 (-4.21)	$I(0)$
$\Delta(LPX_t)$	(6)	$D72$	-9.45	-3.53 (-4.21)	$I(0)$

Notes: The dummy variables are specified as follows: $D70, D72, D73, D74, D82$, are impulse dummy variables with zeros everywhere except for a one in 1970, 1972, 1973, 1974, 1982. $DU70, DU72, DU73, DU74, DU82$ are 1 from 1970, 1972, 1973, 1974, 1982 onwards and 0 otherwise. $DT82$ is 0 before 1982 and t otherwise. Critical values for the levels are provided by Perron [1997]. Critical values for the first differences are from MacKinnon [1991]. For the first differences only impulse dummy variables were included in the regression. Impulse dummy variables, that is those with no long-run effect, do not affect the distribution of the MacKinnon test statistics.

However, we do need to be cautious in interpreting the results. As Lumsdaine and Papell [1997] point out, results of unit root tests are sensitive to the assumed structural breaks. The authors show that the results obtained using one endogenous break are often reversed when a model with two breaks is estimated. This introduces a degree of uncertainty to the analysis. Therefore we check the validity of the results represented in Table 1 by considering the possibility that two break points occurred over the relevant time period. We employ Kapetanios' [2002] test for the null hypothesis of a unit root against the alternative hypothesis of an unspecified number of structural breaks. We estimate two models:

$$y_{1t} = \mu_1 + b_1t + a_1y_{t-1} + \sum_{i=1}^m \delta_{1i}DU_{i,t} + \sum_{i=1}^k c_{1i}\Delta y_{t-1} + e_{1t} \quad (8)$$

$$y_{2t} = \mu_2 + b_2t + a_2y_{t-1} + \sum_{i=1}^m \delta_{2i}DT_{i,t} + \sum_{i=1}^k c_{2i}\Delta y_{t-1} + e_{2t} \quad (9)$$

where y_t is the variable considered, m denotes the number of breaks, and $DU_{i,t}$ and $DT_{i,t}$ are defined as in equations (6) and (7). Setting $m = 2$, model (8) allows for two breaks in the intercept of the trend function.

In model (9) the two breaks are restricted to the slope of the trend function. Since visual inspection of the capital stock data suggests only possible changes in the slope regression (9) is applied to LK_t . For LNY_t , LL_t , LCM_t , LIX_t , and LPX_t we carry out a regression of type (8), where both breaks in the trend function are restricted to the intercept. Running the regressions for all indicator dummy variables, we choose the date of the first structural break such that the sum of squared residuals is smallest among all possible break points in the data set. Imposing the estimated break date on the sample, we start looking for the second break. Again, the second break point is associated with the minimum of squared residuals.

The results of testing the unit root null against the two-break alternative are reported in Table 2. Except for some break points [dummy variables], they do not differ from the results of the Perron [1997] procedure. The results of both the Perron and the Kapetanios unit root test show that the null hypothesis of a unit root cannot be rejected for all time series in levels. Since for the first differences, the unit root hypothesis can be rejected, it is concluded that LNY_t , LK_t , LL_t , LCM_t , LIX_t and LPX_t are integrated of order one, $I(1)$. Therefore, the next step in our analysis is an investigation of the cointegration properties of the variables.

TABLE 2:
KAPETANIOS [2002] UNIT ROOT TEST

Series	Model	Dummy Variables	Test Statistic $t_{\hat{\alpha}}$	Critical Value 5% (1%)	Result
Levels					
LNY_t	(8)	$DU74, DU82$	-4.77	-5.69 (-6.16)	$I(1)$
LK_t	(9)	$DT75, DT82$	-2.95	-6.11 (-6.59)	$I(1)$
LL_t	(8)	$DU74, DU82$	-2.73	-5.69 (-6.16)	$I(1)$
LCM_t	(8)	$DU71, DU81$	-3.69	-5.69 (-6.16)	$I(1)$
LIX_t	(8)	$DU74, DU81$	-3.92	-5.69 (-6.16)	$I(1)$
LPX_t	(8)	$DU71, DU81$	-4.01	-5.69 (-6.16)	$I(1)$
First Differences					
$\Delta(LNY_t)$	(8)	$D75, D82$	-4.95	-3.53 (-4.23)	$I(0)$
$\Delta(LK_t)$	(9)	$D75, D82$	-3.54	-3.53 (-4.23)	$I(0)$
$\Delta(LL_t)$	(8)	$D74, D82$	-4.90	-3.53 (-4.23)	$I(0)$
$\Delta(LCM_t)$	(8)	$D71, D81$	-6.82	-3.53 (-4.23)	$I(0)$
$\Delta(LIX_t)$	(8)	$D74, D81$	-5.36	-3.53 (-4.23)	$I(0)$
$\Delta(LPX_t)$	(8)	$D71, D81$	-8.37	-3.53 (-4.23)	$I(0)$

Notes: The dummy variables are specified as follows: $D71, D74, D75, D81, D82$ are impulse dummy variables with zeros everywhere except for a one in 1971, 1974, 1974, 1981, 1982. $DU71, DU74, DU75, DU81, DU82$ are 1 from 1971, 1974, 1974, 1981, 1982 onwards and 0 otherwise. $DT82$ ($DT75$) is 0 before 1982 (1975) and t otherwise. Critical values for the levels are provided by Kapetanios [2002]. Critical values for the first differences are from MacKinnon [1991]. For the first differences, only impulse dummy variables were included in the regression. Impulse dummy variables, that is, those with no long-run effect, do not affect the distribution of the MacKinnon Test statistics.

Testing for Cointegration: The Engle-Granger Method

We use the Engle-Granger [1987] approach for testing the null of no cointegration. The null of no cointegration implies that the estimated residuals, $\hat{\epsilon}_t$, from equation (5) are $I(1)$, whereas the

alternative hypothesis of cointegration implies that the estimated residuals are $I(0)$. Two test statistics are computed to test for no cointegration: The first one is the DW statistic from regression (5), which is commonly denoted as CRDW. The second one is the augmented Dickey-Fuller [ADF] t -statistic, which is estimated according to:¹⁴

$$\Delta \hat{e}_t = \rho \hat{e}_{t-1} + \sum_{j=1}^k \beta_j \Delta \hat{e}_{t-j} + v_t. \quad (10)$$

If $t_{\hat{\rho}}$ and CRDW are (in absolute value) greater than the critical values, we reject the null. The critical values are reported in Banerjee et al. [1993]. The results of this testing procedure are summarised in Table 3.¹⁵

TABLE 3
ENGLE-GRANGER COINTEGRATION TESTS

CRDW	Critical Value (5%)	$t_{\hat{\rho}}$ (ADF)	Critical Value (1%)
1.58	1.19	-5.10	-4.97

Notes: Banerjee et al [1993], Table 7.1, generated (only) 5% Critical Values for CRDW ($T=50$). Critical Values for the residual-based ADF are from Banerjee et al [1993], Table 7.2.

As can be seen, both the cointegration regression Durbin-Watson and the ADF test statistics suggest that we can reject the null hypothesis of no cointegration at least at the 5% significance level. Thus, equation (5) can be regarded as a long-run equilibrium relationship.

Testing for Cointegration: The Johansen Method

We provide additional evidence regarding cointegration by applying the multivariate cointegration technique developed by Johansen [1995]. The Johansen approach estimates cointegration relationships between $I(1)$ series using a maximum likelihood procedure, which tests for the number

of cointegration relationships. The method is based on the unrestricted vector autoregression [VAR(p)] model represented by the following equation:

$$y_t = \mu + \sum_{k=1}^p \Pi_k y_{t-k} + \varepsilon_t, \quad (11)$$

where y_t is an $(n \times 1)$ column vector of n $I(1)$ variables, Π_k is a coefficient matrix, μ represents a $(1 \times n)$ vector of constants, p denotes the lag length, and ε_t is a disturbance term independently and identically distributed with zero mean and constant variance.

Since $y_t = [LNY_t, LK_t, LL_t, LCM_t, LIX_t, LPX_t]'$ is assumed to be $I(1)$, letting $\Delta y_t = y_t - y_{t-1}$, equation (11) can be rewritten in first difference notation, reformulated in vector error correction [VECM] form as:

$$\Delta y_t = \mu + \sum_{k=1}^{p-1} \Gamma_k \Delta y_{t-k} + \Pi y_{t-1} + \varepsilon_t, \quad (12)$$

where Γ_k and Π represent coefficient matrices and the rank r of matrix Π determines the number of cointegration relations in the system.

As Δy_t and Δy_{t-1} variables are $I(0)$ and y_{t-1} variables are $I(1)$, equation (12) will be balanced if the left-hand side and the right hand-side have the same degree of integration. This will either occur if $r = 0$, so that $\Pi = 0$, in which case the variables in y_t are not cointegrated, or if the parameters of Π are such that Πy_{t-1} is also $I(0)$. In the first case ($r = 0$; $\Pi = 0$), equation (12) is just a traditional VAR model in first differences. The second case applies when the rank of Π is greater than zero, indicating that there will exist $r < n$ cointegration relations, meaning r possible stationary linear

combinations of y_t . If $0 < r < n$, the reduced-rank matrix Π can be decomposed into two matrices α and β [each $n \times r$], such that

$$\Pi y_{t-1} = \alpha (\beta' y_{t-1}). \quad (13)$$

Here the loading matrix α contains the error correction coefficients measuring the speed of adjustment toward equilibrium. The second term on the right-hand side ($\beta' y_{t-1}$) represents the cointegration relations. The cointegrating vectors β have the property that $\beta' y_t$ is stationary even though y_t itself is nonstationary.

The number of cointegrating vectors [the cointegration rank], r , can be formally tested with the trace and the maximum-eigenvalue statistics. The trace statistic tests the null hypothesis that the number of distinct cointegration vectors is less than or equal to r against the general alternative of n cointegrating vectors. The maximum-eigenvalue test evaluates the null hypothesis of r cointegration vectors against the alternative of $r+1$ cointegration vectors.

To determine the optimal lag length, p , the Schwarz information criterion is used. The Schwarz criterion has been shown to choose the correct lag length more often than other information criteria in the VAR process [Lütkepohl 1985]. This criterion suggests one lag for our VAR model.

Table 4 reports the trace and the maximum-eigenvalue statistics from the cointegration tests based on the VAR(1). Comparing both the trace and the maximum-eigenvalue statistics with the corresponding critical values, it can be seen that the null hypothesis of no cointegration, $r = 0$, can be rejected at the 5% and the 1% significance level, but not the null of at most one cointegrating vector.

TABLE 4

JOHANSEN'S COINTEGRATION TEST

Hypothesised no. of cointegrating vectors	Trace statistics		Maximum-eigenvalue statistics	
	Statistics	Critical Value 0.01% (0.05%)	Statistics	Critical Value 0.01% (0.05%)
None, $r = 0$	119.77***	103.18 (94.15)	51.33***	45.10 (39.37)
At most 1	68.44	76.07 (68.52)	32.16	38.77 (33.46)

Notes: The *** indicate a rejection at the 1% level. Critical values are taken from Osterwald-Lenum [1992]. The model includes an unrestricted constant.

We therefore conclude that there is a single cointegrating vector. This stationary vector is estimated after normalising on $LN Y_t$, so that we obtain the following long-run relation:¹⁶

$$LN Y_t = 0.742^{***} LK_t + 0.914^{***} LL_t + 0.078^{***} LCM_t + 0.033^{***} LIX_t - 0.428^{***} LPX_t + 0.673 + ec_t \quad (14)$$

(20.15) (14.27) (3.94) (3.74) (11.94)

From equation (14), it can be inferred that Chilean non-export GDP increases by 0.033 percent in response to a one percent increase in manufactured exports. In contrast, a one percent increase in primary exports leads to a 0.428 percent decrease in non-export GDP. This result suggests that manufactured exports promote economic growth via increases in productivity. In contrast, primary exports seem to have a negative impact on total factor productivity.

Testing for Long-Run Causality and Weak Exogeneity

However, up to now we have implicitly assumed that long-run causality runs from LK_t , LL_t , LCM_t , LIX_t , and LPX_t to $LN Y_t$. This assumption will hold if the error correction coefficient α_1 of the lagged error correction term ec_t ,

$$ec_t = LN Y_t - (0.742LK_t + 0.914LL_t + 0.078LCM_t + 0.033LIX_t - 0.428LPX_t + 0.673), \quad (15)$$

in the VECM representation

$$\begin{bmatrix} \Delta LNY_t \\ \Delta LK_t \\ \Delta LL_t \\ \Delta LCM_t \\ \Delta LIX_t \\ \Delta LPX_t \end{bmatrix} = \begin{bmatrix} \mu_1 \\ \mu_2 \\ \mu_3 \\ \mu_4 \\ \mu_5 \\ \mu_6 \end{bmatrix} + \sum_{k=1}^{p-1} \Gamma_k \begin{bmatrix} \Delta LNY_{t-k} \\ \Delta LK_{t-k} \\ \Delta LL_{t-k} \\ \Delta LCM_{t-k} \\ \Delta LIX_{t-k} \\ \Delta LPX_{t-k} \end{bmatrix} + \begin{bmatrix} \alpha_1 \\ \alpha_2 \\ \alpha_3 \\ \alpha_4 \\ \alpha_5 \\ \alpha_6 \end{bmatrix} ec_{t-1} + \begin{bmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \\ \varepsilon_{3t} \\ \varepsilon_{4t} \\ \varepsilon_{5t} \\ \varepsilon_{6t} \end{bmatrix} \quad (16)$$

is significantly different from zero. A significant error correction term indicates long-run Granger causality from the explanatory to the dependent variables [Granger 1988], where “long-run Granger causality/Granger non-causality” and “endogeneity/weak exogeneity” can be regarded as equivalent [Hall and Milne 1994].

Similar to Lütkepohl and Wolters [1998] and Juselius [2001], we test for weak exogeneity by imposing zero restrictions on the insignificant short-run parameters (Γ_k) and then we decide on the significance of the α s. In doing so, we reduce the number of parameters [according to Hendry’s general-to-specific methodology] and thereby we increase the precision of the weak exogeneity tests on the α coefficients. Since all variables in (16), including ec_{t-1} , are $I(0)$ variables, conventional t - and F -tests can be used. Given the low frequency of the data [annual] and the small sample size, we start with two lags in the VECM. After applying the general-to-specific model reduction procedure, we obtain the following results [Table 5]:

TABLE 5

WEAK EXOGENEITY TESTS ON THE ERROR CORRECTION COEFFICIENTS / LONG-RUN CAUSALITY TESTS

	α_1 (LNY_t)	α_2 (LK_t)	α_3 (LL_t)	α_4 (LCM_t)	α_5 (LIX_t)	α_6 (LPX_t)
t -Value of α	-4.11***	-2.63**	-1.30	-2.05**	-4.29***	-9.10***

Notes: ** and *** denote the 5% and 1% level of significance. Corresponding variables which were tested for weak exogeneity are in parentheses.

According to the t -statistics in Table 5 all error correction coefficients are significantly different from zero except α_3 . Thus, the weak exogeneity tests suggest a long-run feedback relationship between non-export GDP, capital stock, capital goods imports, exports of manufactured products and primary products since all these variables can be regarded as endogenous to the system. In other words, long-run causality runs from LK_t , LL_t , LCM_t , LIX_t and LPX_t to LNY_t but LK_t , LCM_t , LIX_t and LPX_t are not weakly exogenous with respect to the long-run parameters.¹⁷ The only variable which is weakly exogenous to the long-run relationship is labour.

Reestimation of the Long-Run Elasticities: Dynamic OLS Results

Before drawing some conclusions about the impact of increasing manufactured and primary exports on productivity, we check the robustness of the cointegration estimates. We reestimate equation (5) by means of the DOLS procedure developed by Saikkonen [1991] because FIML estimations are very sensitive to the choice of the lag length and the specification of the statistical model.

These criticisms do not apply to the single-equation-based DOLS regression, which has been shown to provide unbiased and asymptotically efficient estimates, even in the presence of endogenous regressors [Stock and Watson 1993]. The DOLS regression in our case is given by equation (17) below:

$$\begin{aligned}
 LNY_t = & c + \alpha LK_t + \beta LL_t + \delta LCM_t + \gamma LIX_t + \rho LPX_t \\
 & + \sum_{i=-k}^{i=k} \Phi_1 \Delta LK_{t+i} + \sum_{i=-k}^{i=k} \Phi_2 \Delta LL_{t+i} + \sum_{i=-k}^{i=k} \Phi_3 \Delta LCM_{t+i} + \sum_{i=-k}^{i=k} \Phi_4 \Delta LIX_{t+i} + \sum_{i=-k}^{i=k} \Phi_5 \Delta LPX_{t+i} + \varepsilon_t
 \end{aligned} \tag{17}$$

where α , β , δ , γ , and ρ are the long-run elasticities, and Φ_1 , Φ_2 , Φ_3 , Φ_4 , Φ_5 are coefficients of lead and lag differences, which are treated as nuisance parameters. These serve to adjust for possible endogeneity, autocorrelation, and non-normal residuals and result in consistent estimates of α , β , δ , γ , and ρ . Similar to model (16), the DOLS equation is estimated with up to two leads and lags ($k=2$).

The following equation results by applying the general-to-specific modelling approach where the least significant variables are successively eliminated [*t*-statistics are given in parentheses beneath the estimated coefficients]:¹⁸

$$\begin{aligned}
 LNY_t = & 0.811^{***} LK_t + 0.865^{***} LL_t + 0.058^{***} LCM_t + 0.042^{***} LIX_t - 0.459^{***} LPX_t \\
 & (20.67) \quad (13.01) \quad (3.299) \quad (4.635) \quad (-13.41) \\
 & + 0.480 + 1.326^{***} \Delta LK_t - 0.491^{***} \Delta LL_t - 0.056^{***} \Delta LIX_t \\
 & (1.377) \quad (3.114) \quad (-3.559) \quad (-4.287) \\
 & - 0.860^{***} \Delta LK_{t-1} + 0.554^{**} \Delta LK_{t+1} - 0.069^{***} \Delta LXI_{t-1} - 0.297^{***} \Delta LXP_{t-1} \\
 & (-2.841) \quad (2.177) \quad (-5.609) \quad (-7.912)
 \end{aligned} \tag{18}$$

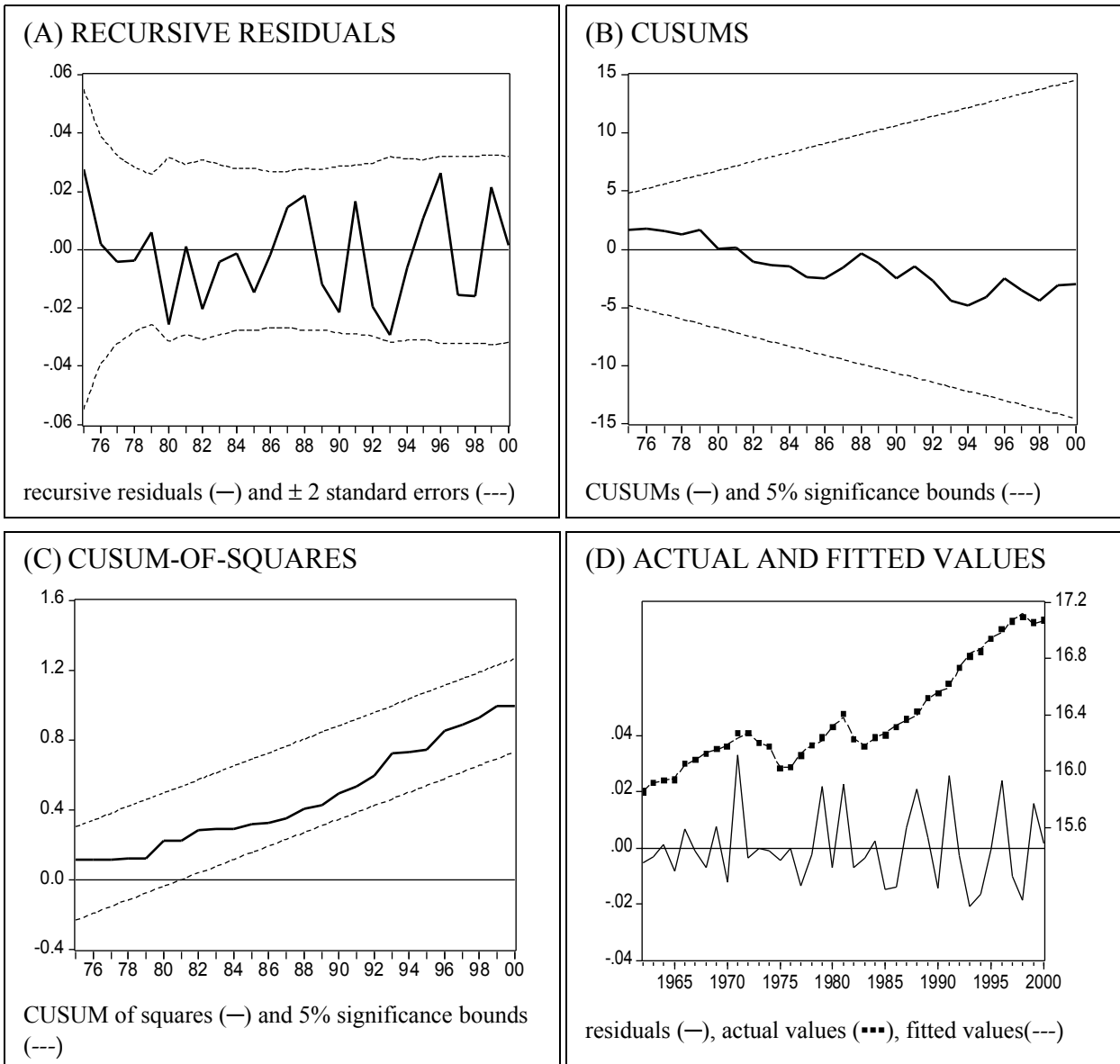
$$\begin{aligned}
 \bar{R}^2 = 0.998 \quad SE = 0.016 \quad JB = 3.819(0.15) \\
 ARCH(1) = 0.932(0.34) \quad ARCH(2) = 0.570(0.57) \quad ARCH(4) = 0.880(0.487) \quad , \\
 LM(1) = 1.452(0.23) \quad LM(3) = 1.876(0.16) \quad LM(5) = 1.175(0.35)
 \end{aligned}$$

where the numbers in parentheses behind the values of the diagnostic test statistics are the corresponding *p*-values. All these test statistics suggest that the model is well specified: The assumption of normally distributed residuals cannot be rejected [*JB*] and the Lagrange multiplier [*LM*] tests for autocorrelation based on 1, 3 and 5 lags, respectively, do not indicate any problems concerning autocorrelated residuals. The model also passes the LM tests for autoregressive conditional heteroscedasticity [*ARCH*(*k*)] of order *k* = 1, 2, 4.

Moreover, in Figure 2 (A)-(C) recursive residuals (A), CUSUM (B) and CUSUM of square-tests (C) are presented, which overall support a stable relation for the period of interest. Accordingly, the model does a good job even in the Chilean ‘breakdown periods’ [1973, 1975, 1982]. Furthermore, figure x (D) shows that equation (18) fits the actual data very well and the plot of the residuals indicates that equation (18) is stationary. Thus, statistically valid inferences can be drawn from the estimated long-run elasticities:

FIGURES 2 (A)-(D)

STABILITY TESTS, ACTUAL AND FITTED VALUES



As in the FIML estimation, the effect of manufactured exports on non-export GDP is significantly positive. According to equation (18) a one percent increase in manufactured exports leads to a 0.042 percent increase in non-export GDP. The effect of primary exports on non-export GDP is again found to be strong and negative, implying that Chilean non-export GDP decreases by 0.459 percent in response to a one percent increase in primary exports. The magnitude of the coefficients in equation (18) does not differ substantially from equation (14). From this, it follows that the

coefficient estimates are fairly robust to different estimation techniques. Both the DOLS and the FIML estimation results can be interpreted as evidence of productivity-enhancing effects of manufactured exports and of productivity-limiting effects of primary exports. This finding is in line with the results of Ghatak et al. [1997], who demonstrated a negative effect of primary exports and a positive effect of manufactured exports on real GDP and non-export real GDP in Malaysia. However, further studies are needed to establish the role of primary and manufactured exports in the economic growth of developing countries.

V. SUMMARY AND CONCLUSIONS

This paper has used single-equation and system cointegration techniques to examine the productivity effects of manufactured and primary exports in the context of the export-led growth hypothesis. The examination was based upon Chilean time series data for 1960 – 2001. To overcome the problem of specification bias under which previous studies have suffered, an augmented neoclassical production function was developed. In the production function framework, total factor productivity was assumed to be a function of primary, manufactured exports and capital goods imports. The output variable of the function was defined net of exports to separate the influence of exports on output from that incorporated into the national income identity. The results of the production function estimation suggest that (1) there exists a long-run relationship between capital, labour, capital goods imports, manufactured exports, primary exports and non-export GDP; (2) the results indicate long-run Granger causality running from capital stock, aggregate employment, capital goods imports, and exports of manufactured products and primary products to non-export GDP, where capital stock, capital goods imports, exports of manufactured products and primary products are also endogenous; (3) However, primary-product exports were found to have a

statistically negative impact, whereas manufactured-product exports have a statistically positive impact on non-export GDP. This result is robust to different estimation techniques. In connection with the theoretical foundations underpinning our model, the estimation results can be interpreted as evidence of productivity-enhancing effects of manufactured exports and of productivity-limiting effects of primary exports. The latter may be due to the problem of fluctuating commodity export prices and earnings, especially copper prices, which is well known in the Chilean literature.¹⁹ Additionally, manufactured exports might offer greater potential for knowledge spillovers and other externalities than primary exports. Accordingly, the primary conclusion that emerges from this study is that while primary and manufactured export earnings certainly contributed to the Chilean national income, exports of manufactured products have been especially important for productivity and thus for long-run economic growth. This conclusion has crucial policy implications. It is particularly important to promote exports of manufacturing goods – by avoiding trade-distorting measures that would counteract the comparative advantages, and building new comparative advantages and export opportunities in the Chilean manufacturing sector.

NOTES

1. See Giles and Williams [2000] for a comprehensive survey of the empirical literature.
2. For a critical review of cross-country studies, see Giles and Williams [2000].
3. See Giles and Williams [2000].
4. See for example the studies of Ghatak et al. [1997], Agosin [1999], Awukese [2003].
5. This problem is often ignored in the recent literature, e.g., Agosin [1999], Lee and Huang [2002], Abu-Qarn and Abu-Bader [2004].
6. See for example the studies by Shan and Sun [1998], Agosin [1999], Lee and Huang [2002], Awukese [2003], Abual-Foul [2004].
7. Greenaway and Sapsford [1994], among others, recommend the use of the national product net of exports.

8. Since 1974, the growth of exports has been very rapid. In the seven years from 1974 to 1980, the annual growth rate of exports was 17.8 percent. However, the export growth rate became negative in the period 1981-1985, with an average annual decrease of 1.5 percent due to the appreciation of the real exchange rate and the slowdown of the world economy. The second phase of high export growth rates began in 1985 after the real exchange rate had been sharply devaluated. Exports grew at an average rate of 10 percent per year between 1985 and 2001. See, also for example, Agosin [1999].
9. See, for example, Romaguera and Contreras [1995]. In their paper, the authors showed that Chilean activity responds strongly to the price of copper.
10. In 1971-1973 the share of copper represented almost 80% of total exports of goods, and the share of minerals as a whole announced to almost 90%. See, for example, Agosin [1999].
11. As already mentioned, exports make up a large part of Chilean GDP [Agosin 1999]. In that, rapid increases in exports automatically have an impact on GDP growth.
12. Biased OLS estimators may be due to the exclusion of short run dynamics and the presence of endogenous explanatory variables.
13. The *additive outlier model* implies that the change in the trend function is sudden. The *innovational outlier model* implies that the break in the series does occur gradually.
14. The lag length k is chosen to minimise the Schwarz criterion ($k = 0$).
15. The estimated coefficients from the cointegrating regression are not reported in Table 3, since standard regression interpretation of the coefficients is not valid.
16. t -ratios in parentheses underneath the estimated coefficients; *** denote the 1% level of significance.
17. Note: We are not interested in the short-run Granger causality, but in the long-run effects. Therefore we do not test for the joint significance of the lagged variables in equation (16).
18. ** and *** denote the 5% and 1% level of significance.
19. Romaguera and Contreras [1995], for example, find that copper price volatility had negative effects on Chilean GDP growth.

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