Are exports and imports of Chile cointegrated?

DIERK HERZER[*] and FELICITAS NOWAK-LEHMANN D.

Ibero-America Institute for Economic Research, University of Goettingen, Platz der

Goettinger Sieben 3, 37073 Goettingen, Germany

This study examines the long-run relationship between Chilean exports and imports during the 1975-2004 period using unit root tests and cointegration techniques that allow for endogenously determined structural breaks. The results indicate that there exists a long-run equilibrium between exports and imports in Chile, despite the balance-of-payments crisis of 1982-83. This finding implies that Chile's macroeconomic policies have been effective in the long-run and suggests that Chile is not in violation of its international budget constraint.

I. Introduction

Over the past decade, the long-run equilibrium or cointegration relationship between exports and imports has received increasing attention (see, e.g., Bahmani-Oskooee and Rhee (1997), Arize (2002), Irandoust and Ericcsson (2004), Narayan and Narayan (2005)). Cointegration between exports and imports implies that trade deficits are only short-term phenomena and thus sustainable in the long-run. This, in turn, means that countries are not in violation of their international budget constraint, since their

---

[*] Corresponding author. E-mail: DierkHerzer@gmx.de

macroeconomic policies have been effective in bringing exports and imports into a long-run equilibrium.

In this paper, we investigate whether a long-run equilibrium exists between exports and imports in Chile. Chile is an interesting case study because Chilean exports and imports grew very rapidly after 1974, when trade liberalisation was initiated. However, the upward movement in exports was partially reversed in 1979-83 because of the appreciation of the real exchange rate and the slow down of the world economy. The result was a large trade deficit, leading to a balance-of-payments crisis and ensuing recession in 1982-83 due to capital flow reversals and high international interest rates. In 1984, Chile finally recovered as exports increased after the exchange rate had depreciated.

Against this background the question arises whether there exists a long-run equilibrium relationship between Chilean exports and imports after the trade liberalisation of 1974.

Indeed, Arize (2002) in a study for 50 countries provides evidence of cointegration in Chile for the period 1973-1998, but he also provides evidence of parameter instability in the estimated relationship due to unmodelled structural breaks. If, however, the estimated long-run relationship is not stable, conclusions can be highly misleading.

Our study differs from that of Arize (2002) and many other studies in the following ways: First, because standard unit root tests are biased towards the null of a unit root in the presence of structural breaks, we use the Perron (1997) test for a structural break at an unknown change point. Second, since standard cointegration tests tend to spuriously reject the null of no cointegration in the face of structural breaks, we apply the Gregory and Hansen (1996) cointegration technique that allows for an

endogenously determined structural break in the cointegration relationship. Third, we estimate the long-run relation between exports and imports, taking into account the endogenously determined structural break. Fourth, based on this relationship we estimate an error correction model for the trade balance. And finally, we conduct misspecification and structural stability tests.

The rest of the paper is organised as follows. Section II describes the theoretical background, Section III presents the empirical results, and Section IV concludes.

II.   Theoretical Background

Husted (1992) provides a simple framework that implies a long-run relationship between exports and imports. He starts with the individual current-period budget constraint given by

$$C_0 = Y_0 + B_0 - I_0 - (1 + r)B_{-1},$$ (1)

where C_0 is current consumption; Y_0 is output, I_0 is investment, r is the one-period world interest rate, B_0 is the international borrowing, and $(1 + r_0)B_{-1}$ is the historically given initial debt.

Since equation (1) must hold in every time period, the period-by-period budget constraints can be combined to form the country's intertemporal budget constraint which states that the amount a country borrows (lends) in international markets equals the present value of future trade surpluses (deficits). Husted than makes several assumptions to derive a testable model, which is given by

$$X_t = \alpha + \beta M_t + \varepsilon_t,$$ (2)

or alternatively, as suggested by Arize (2002),

$$M_t = a + bM_t + e_t, \tag{3}$$

where M_t is imports of goods and services and X_t is exports of goods and services. The null hypothesis states that the intertemporal international budget constraint of the country is satisfied. This implies that $b = 1$, and e_t is a stationary process, or in other words, if M_t and X_t are integrated, then under the null hypotheses they are cointegrated with the cointegrating vector (1, -1).

III.   Data and Results

Data

We use annual data over the period 1975-2004 to estimate equation (3). They were gathered from the *Indicadores económicos y sociales de Chile 1960 -2000* and the *Boletínes mensuales* published by the Chilean Central Bank. Following Irandoust and Ericsson (2004) imports (M_t) and exports (X_t) are evaluated in local currency at constant prices (Chilean pesos at constant 1996 prices) and expressed in natural logarithms.

Unit root test results

In the first step, we test the variables for unit roots. It is well known that standard unit root tests are biased in favour of identifying data as integrated in the presence of structural breaks. For the two series M_t and X_t there is a strong likelihood that they are subject to structural breaks due to the balance-of payments crisis of 1982-83. 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. Perron presents the following models:

$$y_{1t} = \mu_1 + \theta_1 DU_t + b_1 t + \delta_1 D(TB)_t + a_1 y_{t-1} + \sum_{i=1}^{k} c_{1i} \Delta y_{t-1} + e_{1t}, \tag{4}$$

$$y_{2t} = \mu_2 + b_2 t + \delta_2 DT_t + \hat{y}_t$$
$$\hat{y}_t = a_2 \hat{y}_{t-1} + \sum_{i=1}^{k} c_{2i} \Delta \hat{y}_{t-1} + e_{2t}, \tag{5}$$

where y_{1t} and y_{2t} are the series of interest ($y_{1t} = M_t$, X_t and $y_{2t} = M_t$, X_t), t is the deterministic trend component, $TB \in T$ ($T = 30$, $1 \le t \le 30$) denotes the time at which the break in the trend component 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.

Model (4) allows for a one-time change in the intercept of the trend function, whereas model (5) allows for a change in the slope of the trend function without a change in the level. 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 $a_2 = 1$:

$$t_{\hat{a}}^{*}(i) = Min_{TB} t_{\hat{a}}(i, TB, k), \tag{6}$$

where $t_{\hat{a}}(i, TB, k)$ is the t-statistic for testing $a = 1$ under model $i = 1, 2$ [model (4) and

(5)] with a break date TB and truncation lag parameter k. If $Min_{TB} t_{\hat{a}}(i, TB, k)$ exceeds (in absolute value) the critical value reported by Perron (1997), the hypothesis of difference stationarity and a unit root is rejected.

Table 1 contains the results of the sequential unit root tests for the variables in levels and in first differences. The results show that the null hypothesis of a unit root cannot be rejected for the two time series in levels. Since for the first differences, the

unit root hypothesis can be rejected, it is concluded that M_t and X_t are integrated of order one, $I(1)$.

Cointegration test results

In the next step we use the standard Engle-Granger (1987) approach for testing the null of no cointegration. The null of no cointegration implies that the estimated residuals, $\hat{e}_t$, from equation (3) 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 DW statistic from regression (3), which is commonly denoted as CRDW, and the augmented Dickey-Fuller (ADF) *t*-statistic, which is estimated according to:

$$\Delta \hat{e}_t = \rho \hat{e}_{t-1} + \sum_{j=1}^{k} \beta_j \Delta \hat{e}_{t-j} + v_t . \tag{7}$$

If $t_{\hat{\rho}}$ and CRDW are (in absolute value) greater than the critical values reported by Banerjee et al. (1993), we reject the null. The results of this testing procedure are summarised in Table 2. Both the cointegration regression Durbin-Watson and the ADF test statistics suggest that we cannot reject the null hypothesis of no cointegration.

However, standard cointegration techniques are biased towards rejecting the null of no cointegration, if there is a structural break in the cointegration relationship. Therefore, we therefore apply the Gregory and Hansen (1996) cointegration procedure that allows for an endogenously determined structural break. Gregory and Hansen present the following models:

the level shift model (C)

$$y_{1t} = \mu_1 + \mu_2 \varphi_{tt} + \alpha^T y_{2t} + e_t, \qquad t = 1, ..., n, \tag{8}$$

the slope change model (C/T)

6

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \beta t + \alpha^T y_{2t} + e_t, \qquad t = 1, \ldots, n, \tag{9}$$

and the regime shift model (C/S)

$$y_{1t} = \mu_1 + \mu_2 \varphi_{t\tau} + \alpha_1^T y_{2t} + \alpha_2^T y_{2t} \varphi_{t\tau} + e_t, \qquad t = 1, \ldots, n, \tag{10}$$

where y_{1t} and y_{2t}, in the context of our analysis are M_t and X_t, μ_1 and α_1 are the intercept and the slope coefficients before the shift, μ_2 and α_2 denote the changes to the intercept and slope coefficients at the time of the shift. The dummy variable $\varphi_{t\tau}$ is defined by

$$\varphi_{t\tau} = \begin{cases} 0, & if \quad t \le [\eta \tau], \\ 1, & if \quad t > [\eta \tau], \end{cases} \tag{11}$$

where the unknown parameter $\tau \in (0,1)$ denotes the (relative) timing of the change point, and $[\quad]$ denotes the integer part.

Similar to the above sequential unit root tests, the models (8) - (10) are estimated for each possible break date in the data set (for each τ). We obtain the residuals $\hat{e}_{t\tau}$. Next a unit root test is performed on the estimated residuals, where the smallest values of the unit root test statistics are used to test the null of no cointegration against the alternative hypothesis of cointegration with a structural break.

In this study we use the ADF t-statistics on the residuals (obtained from the C, C/T, C/S models), which are calculated by estimating equation (7). As mentioned above, our statistics of interest are the smallest values of the ADF statistics, across all values of $\tau \in T$ and thus

$$ADF^* = \inf_{\tau \in T} ADF(\tau). \tag{12}$$

Asymptotic critical values are provided by Gregory and Hansen (1996). Table 3 presents the results. As can be seen, all three models show evidence of cointegration

with a structural break in 1982 indicated by the dummy variable *DU*83. Note that the break point detected by the Gregory and Hansen procedure coincides with the balance-of-payments crisis of 1982-83.

Dynamic OLS results

In the next step, we apply the Dynamic OLS (DOLS) procedure developed by Saikkonen (1991) to estimate the cointegration relation given by equation (3). DOLS generates unbiased and asymptotically efficient estimates for variables that cointegrate, even with endogenous regressors (Stock and Watson 1993). Moreover, because the DOLS estimator has good small sample properties, it is often used in small samples like ours (see, e.g., Narayan and Narayan (2004)). The DOLS regression is given by

$$M_t = a + bX_t + \sum_{i=-k}^{i=k} \Phi_i \Delta X_{t+i} + cDU83 + \varepsilon_t, \tag{13}$$

where Φ_i are coefficients of lead and lag differences of the $I(1)$ regressors, which are treated as nuisance parameters. They serve to adjust for possible endogeneity, autocorrelation, and nonnormal residuals. The results are reported in Table 4.

As can be observed, the diagnostic tests underneath the estimated coefficients do not indicate any problems with autocorrelation, heteroscedasticity or nonnormality. Furthermore, the estimated b coefficient is highly significant and close to unity and, moreover, the null hypothesis $b = 1$ cannot be rejected. Restricting the coefficient b to one gives the following long-run relation:

$$M_t = -1.283 + X_t - 0.251DU83 + ec_t. \tag{14}$$

Following Apergis et al. (2000) we interpret this finding as Chile not being in violation of its international budget constraint. However, as expected, the balance-of-payments crises (*DU*83) had strong negative effects on Chilean imports.

8

Error correction model results

In the final step, the lagged error correction term ec_{t-1} from (14) is used to estimate a simple error correction model. The following equation results when the Schwarz criterion is used to choose the lag length (*t*-statistics in parentheses underneath the estimated coefficients. *** denote the 1% level of significance):

$$\Delta M_t = 0.169^{***} - 0.805^{***} ec_{t-1} + 0.014 \Delta X_{t-1} + 0.432^{***} \Delta M_{t-1}$$
$$\qquad (4.198) \qquad (-5.315) \qquad (0.048) \qquad (3.090)$$

$$\tag{15}$$

$\overline{R}^2 = 0.551; \, SE = 0.118; \, JB = 1.823 \, (0.402)$

$ARCH(1) = 0.554 \, (0.463); \, ARCH(2) = 1.442 \, (0.245); \, ARCH(4) = 0.815 \, (0.530)$

$LM(1) = 2.692 \, (0.113) \, LM(2) = 1.442; \, LM(4) = 1.698 \, (0.186)$

Again, the diagnostic tests do not indicate any problems with autocorrelation, heteroscedasticity or nonnormality. Moreover, in Figure 1 stability CUSUM tests are presented, which do not indicate any instability in the estimated equation. As expected, the estimated error correction term is negative and highly significant, which implies cointegration as well as long-run Granger-causality from exports to imports (Granger 1988). Against the background of the above findings, this result suggests that in the long run, an increase of imports by one percent is Granger-caused by an one percent increase of exports, resulting in a long-run trade balance. The insignificant coefficient of ΔX_{t-1} indicates that short-run imbalances do occur, since short-run Granger-causality between exports and imports does not exist.

IV.   Conclusions

The study investigates the long-run relationship between Chilean exports and imports after the trade liberalisation of the mid-1970s using unit root tests and cointegration techniques that allow for structural breaks. Our concern was whether the balance-of-payments crisis of 1982-83 was sustainable in the long-run. We found that Chilean exports and imports are cointegrating indicating that the balance-of-payments crisis was indeed sustainable. This suggests that Chile's macroeconomic policies have been effective in bringing exports and imports into a long-run equilibrium.

References

Apergis, N., K.P. Katrakilidis, and N.M. Tabakis (2000) Current account deficit sustainability: The case of Greece, *Applied Economic Letters*, **7**, 599–603.

Arize, A.C. (2002) Imports and exports in 50 countries: tests of cointegration and structural breaks, *International Review of Economics and Finance*, **11**, 101–115.

Bahmani-Oskooee, M. and H.-J. Rhee (1997) Are Imports and Exports of Korea Cointegrated? *International Economic Journal*, **11**, 109–114.

Banerjee, A., J. Dolado, J.W. Gailbraith, and D.F. Hendry (1993) *Co-integration, error-correction, and the econometric analysis of nonstationary data*. Oxford University Press, New York.

Engle. R.E. and W.J. Granger (1987) Cointegration and error-correction: representation, estimation, and testing, *Econometrica*, **55**, 251–76.

Granger, C.W.J. (1988) Some recent developments in a concept of causality, *Journal of Econometrics*, **39**, 199–211.

Gregory, A. and Hansen, B. (1996) Residual-based tests for cointegration in models with regime shifts, *Journal of Econometrics*, **70**, 99–126.

Husted, S. (1993) The emerging US current account deficit in the 80s: a cointegration analysis, *The Review of Economics and Statistics*, **74**, 159–66.

Irandoust, M. and M. Ericsson (2004) Are imports and exports cointegrated? an international comparison, *Metroeconomia*, **55**, 49–64.

MacKinnon, J. G. (1996) Numeral distribution functions for unit root and cointegration tests, *Journal of Applied Econometrics*, **11**, 602–18.

Narayan, P.K. and Narayan, S. (2005) Are exports and imports cointegrated? Evidence from 22 least developed countries, *Applied Economics Letters*, **12**, 375-378.

Narayan, S. and Narayan, P. K. (2004) Determinants of demand for Fiji's exports: an empirical investigation, *The Developing Economies*, **42**, 95–112.

Perron, P. (1997) Further evidence on breaking trend functions in macroeconomic variables, *Journal of Econometrics*, **80**, 355–385.

Saikkonen, P. (1991) Asymptotically efficient estimation of cointegrating regressions. *Econometric Theory*, **7**, 1–21.

Stock, J.H. and M.W. Watson (1993) Simple estimator of cointegrating vectors in higher order integrated systems. *Econometrica*, **61**, 783–820.
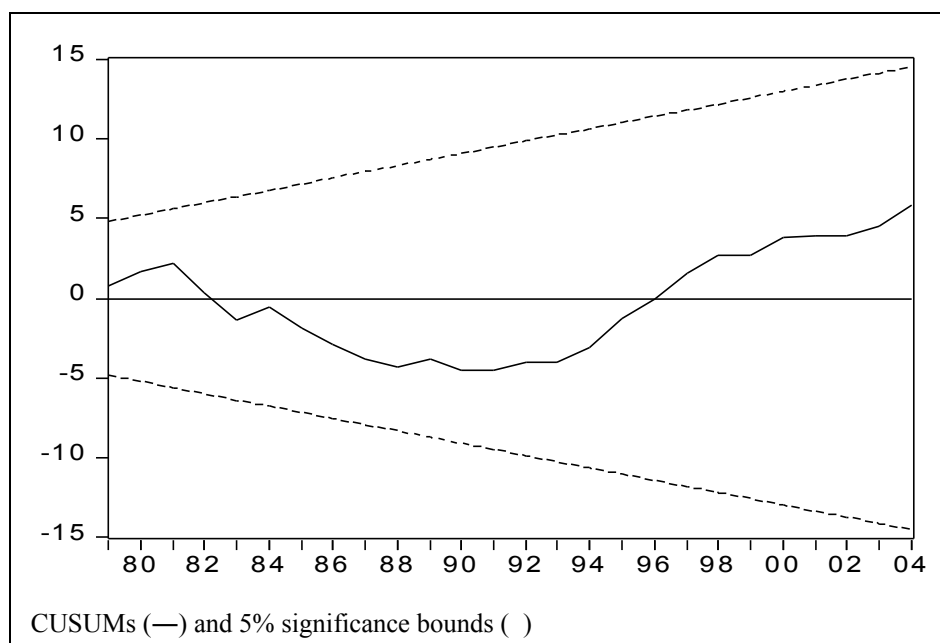
Fig. 1. *Stability CUSUM tests*

Table 1. *Perron unit root test*

Series	Model	Break Point	Dummy Variables	Test Statistic $t_{\hat{a}}$	Critical Value 5% (1%)	Result
Levels						
M_t	(4)	1981	*DU*82, *D*82	-4.182 [3]	-5.23 (-5.92)	*I*(1)
X_t	(4)	1980	*DU*81, *D*81	-3.495 [0]	-5.23 (-5.92)	*I*(1)
M_t	(5)	1982	*DT*83	-3.783 [3]	-4.83 (-5.45)	*I*(1)
X_t	(5)	1982	*DT*83	-3.706 [3]	-4.83 (-5.45)	*I*(1)
First Differences						
$\Delta(M_t)$	(4)		*D*82	-5.160 [0]	-3.595 (-4.356)	*I*(0)
$\Delta(X_t)$	(4)		*D*81	-5.648 [0]	-3.595 (-4.356)	*I*(0)

Notes: The dummy variables are specified as follows: *D*81, *D*82 are impulse dummy variables with zeros everywhere except for a one in 1980, 1982. *DU*81, *DU*82 are 1 from 1981, 1982 onwards and 0 otherwise. *DT*82 is 0 before 1983 and *t* otherwise. Number inside the brackets are the number of lags selected by the Schwarz criterion. Critical values for the levels are provided by Perron (1997). Critical values for the first differences are from MacKinnon (1996). 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.

Table 2. *Engle-Granger cointegration tests*

CRDW	Critical Value (5%)	$t_{\hat{\rho}}$ (ADF)	Critical Value (5%)
0.62	0.72	-2.62	-3.55

Notes: Critical values from Banerjee et al (1993, p. 209, 213). The lag length k was chosen according to Schwarz criterion ($k = 3$).

Table 3. *Gregory-Hansen cointegration tests*

Model	Break Point	Dummy Variables	ADF^*	Critical Value 1% (5%)
C	1982	*DU*83	-4.82**	-5.13 (-4.61)
C/T	1982	*DU*83	-5.46***	-5.45 (-4.99)
C/S	1982	*DU*83	-5.71***	-5.47 (-4.95)

Notes: Critical values from Gregory and Hansen (1996, p. 109); *** (**) significant at 1% (5%). *DU*83 is 1 from 1983 onwards and 0 otherwise. The lag length k was chosen according to the Schwarz criterion ($k = 0$).

Table 4. *DOLS procedure results*

a	b	c	$H_0 : b = 1$ F statistic (*p*-value)
-1.283	1.090	-0.251	2.58 (0.13)
(-1.339)	(16.486)***	(-2.163)**	

$\overline{R}^2 = 0.998$, $SE = 0.089$; $JB = 0.217 \, (0.897)$

$ARCH(1) = 0.496 \, (0.489)$; $ARCH(2) = 0.381 \, (0.688)$; $ARCH(4) = 1.624 \, (0.214)$

$LM(1) = 1.713 \, (0.21)$; $LM(2) = 2.529 \, (0.12)$; $LM(4) = 1.553 \, (0.26)$

Notes: t-statistics in parentheses underneath the estimated coefficients. *** (**) significant at 1% (5%). The number in parenthesis behind the values of the diagnostic tests statistics are the corresponding *p*-values. *JB* is the Jarque-Bera test for normality, *LM* (*k*), k = 1,3, are *LM* tests for autocorrelation based on 1 and 3 lags, respectively and *ARCH* (*k*) is an *LM* test for autoregressive conditional heteroscedasticity of order $k = 1, 2, 4$. To correct for autocorrelation we added up to two leads and lags of the differenced regressors.