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LONG RUN DYNAMICS OF REGIONAL GROWTH IN CHILE*

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Abstract

The aim of this article is to examine the process of economic growth in the regions of Chile using a time series approach. In particular, we test the hypothesis of a unit root in the log of the ratios of per capita product between every possible pair taken from the 13 regions. The 'acceptance' of the null hypothesis means that the ratio of the per capita product doesn't tend to revert to a deterministic constant value and therefore, one of the definitions of convergence in Bernard and Durlauf (1996) is violated. We have found that there are two groups of regions that show within-group convergence but the two groups don't converge. Also, there are two regions that don't converge with any other region of the country.

Resumen

Este trabajo examina al proceso de crecimiento regional en Chile utilizando un enfoque de series de tiempo. En particular, se testea la hipótesis de una raíz unitaria en el (log) de las razones de producto per cápita entre todos los posibles pares de regiones. Se verifica que existen dos grupos de regiones que presentan convergencia entre grupos y dos que no lo hacen.

JEL classification*: O40, C22*

Keywords*: Convergence, Cointegration, Unit root tests, Level Shifts, Chilean Economy.*

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1. Introduction

One of the most important questions for fast growing developing countries is if the growth process involves all of the different regions of the economy. One of the most influential approaches to this kind of question is the Neoclassical Model associated with Robert Solow (1956). The key implication of this theory is the *convergence* hypothesis. This hypothesis predicts that poor economies grow faster than rich economies and so, as the years go by, the initial gap disappears. In the case we are studying, this implies that poor regions would get closer to richer regions if they had a common access to technology, and had similar saving and population growth rates. There are four studies that have tried to assess the validity of the convergence hypothesis in the Chilean case[1] (Anríquez (1996), Fuentes (1996), Morandé *et al.* (1996), and Cuervo and Mella (1998)) and there exists a great amount of literature that has studied this at an international level (see Sala-i-Martin (1996) for a quick review). In general, the evidence supports the hypothesis at both levels. The standard empirical approach of this kind of study relies on cross sectional data analysis. However, as Bernard and Durlauf (1996) point out, this method fails to detect economies inside the sample that are not converging.[2] An alternative approach, applied by Bernard and Durlauf (1995) (BD95 below) to the OECD countries, is based on the long-term properties of GDP per capita series of the different economies. Although this method has its own pitfalls, it is able to detect non-convergent economies. In this paper we apply this approach to the 13 regions of Chile. The main objective is to assess the long run properties of GDP per capita ratios between the different regions for the available data. This data spans from 1960 through to 1996. In particular, we try to answer the following question: is there a unit root in any of series of GDP per capita ratios between regions? If there were a unit root in these series, the ratio of GDP per capita series wouldn't tend to revert to a deterministic constant and some regions wouldn't converge.

This paper is structured as follows; in section 2 we present the methodology and the convergence concepts that we study, also in this section we provide first results. In section 3 we explore structural change effects over the results in section 2. Section 4 summarizes the main conclusions.

2. Convergence and Time Series Analysis

As we pointed out in the introduction, we follow a similar approach in this paper to that used in BD95. The following definition of convergence was used:

Definition 2.1 Convergence as equality of long term forecasts at a fixed time(BD96):
Countries i and j converge if the long-term forecasts of (log) per capita output for both countries are equal at a fixed time t,

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1 Chile is divided into 13 administrative regions comparable to provinces or states.

2 In the Chilean case the only study that is not principally based on cross section data analysis is Morandé *et al.* (1996).

(1) $$Lim\,E(Y_{i,\,t+k} - Y_{j,\,t+k} / I_t) = 0$$

where the log of per capita output of country *i* is denoted by y_i and the limit is taken for $k \rightarrow$. In accordance with this definition if two per capita product series have a unit root, they would converge only if they are cointegrated with cointegrating vector [1,–1].[3]

Although our approach is very similar to that in BD95, there are two noteworthy differences. Firstly, we allow for any constant on the left-hand side of (1). This can be interpreted as allowing economies to converge to a deterministic ratio not necessarily equal to 1.[4] Secondly, we have focused our attention on bivariate cointegration analysis. The main reason for doing this is that the properties of other approaches are not well known when there are regions that converge while other regions don't.[5]

We start our analysis applying the Augmented Dickey and Fuller unit root test (ADF below) to the GDP per capita series of all regions. Table 1 provides the results of this test. As can be seen in this table, there isn't enough evidence to reject the unit root hypothesis for any of these series.

TABLE 1

ADF TEST FOR REGIONS GDP PER CAPITA SERIES

	I	II	III	IV	V	VI	VII	VIII	IX	X	XI	XII	XIII
t_α	–0.27	–2.09	0.44	–1.26	–0.84	–0.66	–0.50	–1.03	–1.01	0.20	–1.67	–1.49	–0.30
10%	–3.20	–3.20	–3.20	–3.20	–3.20	–3.21	–3.20	–3.20	–3.20	–3.20	–3.20	–3.20	–0.20

Figures reported in the first row are *t* tests. Lag truncation procedure is described in Hamilton (1994), chapter 17, with a maximum of 4 lags. The second row provides Mackinnon critical values. The data generating process includes a constant and a trend.

Since the evidence against the unit root hypothesis for regional GDP per capita series isn't strong enough, it is possible to suspect the presence of a unit root in them. In this case, convergence in *definition 2.1* requires series being cointegrated with cointegrating vector [1,–1]. To test if the vector [1,–1] is a cointegrating vector, we defined a new variable $z_t^{i,j}$ as $z_t^{i,j} \equiv y_{i,\,t} - y_{j,\,t}$ for all possible pairs of regions. As the cointegrating vector is known under the null (unit root) hypothesis, cointegration only requires $z_t^{i,j}$ to be stationary. To assess this property we proceed as we did with the GDP per capita series. Table 2

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3 When series are trend stationary, convergence requires equal trends. The relation between *definition 2.1* and the neoclassical model is described in BD96 *Proposition 6.*

4 This could be explained for differences in weather or particular activities in each economy. Mankiw, Romer and Weil (1992) use the same assumption. Also, we point out that *definition 2.1* doesn't imply that the long-term forecast for the ratio of per capita products be equal 1.

5 See Evans (1998).

provides the results of cointegration tests for all possible pairings of regions. As can be seen in this table, there is enough evidence of cointegration for many pairs of regions. On applying the commutative property of cointegration relations we found that there were three groups of regions: a first group of cointegrating regions consisting of Regions II, IV, VI, VII, VIII, IX and X; a second group of cointegrating regions composed for Regions V and XI; and a third group of regions that don't cointegrate with any other region in the country consisting of regions I, III, XII and XIII.[6]

TABLE 2
ADF TEST FOR COINTEGRATION RELATIONS

	I	II	III	IV	V	VI	VII	VIII	IX	X	XI	XII
II	–1.94											
III	1.82	–1.63										
IV	–2.30	–2.45	–0.38									
V	–1.81	–1.28	0.38	–1.78								
VI	–1.76	–0.75	–0.56	–2.88*	–1.77							
VII	–1.26	–3.02*	–1.56	–2.15	–0.74	1.30						
VIII	–2.13	–1.18	–1.62	–2.17	–0.76	0.53	–2.92*					
IX	–1.52	–3.15*	–1.09	–4.38*	0.98	–2.08	–2.79*	–2.47				
X	–1.43	–2.62*	–1.15	–4.81*	–0.39	–1.73	–2.64*	–2.20	–2.96*			
XI	–2.25	–1.76	–0.65	–1.85	–2.66*	–2.44	–1.02	–1.18	–1.66	–1.31		
XII	–0.01	–0.66	1.05	–1.16	–0.77	–1.13	–0.48	–0.21	–0.33	0.04	–1.60	
XIII	–2.13	–1.17	–0.44	–1.64	–1.35	–1.48	–1.05	–0.84	–0.99	–0.94	–1.95	–0.48

Figures reported are *t* tests. Lag truncation procedure is described in Hamilton (1994), chapter 17, with a maximum of 4 lags. The second row provides Mackinnon critical values. The data generating process includes a constant. Rejection at 90% (or higher) confidence level using Mackinnon crtical values is denoted by * (the critical value is –2.62).

From the results above, there are two *cointegrating groups* of regions following a common within-group trend but different between groups. Also there is a third group of *non-cointegrating* regions that follow their own trend, different to all the regions in the country. From this result, we find that there are many ratio series that don't tend to revert to a deterministic constant by the time, and so, the regions involved don't converge.

However, these unit root tests usually have low power. As has been emphasized by Perron (1990), this problem is more serious when there are salient structural changes in the series, like, for instance, a level shift. For developing economies, like Chile, this issue is very important because over long periods of time important political and economic reforms may affect different regions in different ways. The effect that these structural changes may cause in our results is studied in next section.

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6 The fact that XIII region doesn't cointegrate with any other region was first reported by Morandé *et al.* (1996).

3. Structural Change Analysis

During the period that our data set spans (1960-1996), the Chilean economy has been exposed to many important structural reforms such as privatization, the opening of trade and capital, and the application of different exchange rate regimes, amongst others. These reforms may have affected the different regions in different ways. In particular, different regions may have had different level or trend shifts. When one or both of these shifts happen, the unit root tests that we reported in section 2 have even lower power than in the case without shifts. This problem has been addressed in a series of papers. The main references are Perron (1989) and (1990); Zivot and Andrews (1992); and Perron and Vogelsang (1992a) and (1992b). This literature has studied how to test the unit root hypothesis in the presence of level and/or trend shifts.

For the reasons discussed above, we applied the unit root tests developed by Zivot and Andrews (1992) to the GDP per capita series. In this test, the null hypothesis continues to be the presence of a unit root and a trend, but the alternative hypothesis includes the possibility of both a level and trend shift. The alternative hypothesis can be represented by the following process:

$$y_t = \mu_1 + \beta_1 t + \left(\mu_2 - \mu_1\right) DU_t + \left(\beta_2 - \beta_1\right) DT_t + e \quad (2)$$

where $DU_t = 1$ if $t > T_B$ and 0 otherwise, $DT_t = t - T_B$ if $t > T_B$ and 0 otherwise, and e_t a follows the process $A(L)e_t = B(L)v_t$, with $v_t iid(0, \sigma^2)$. In this process, the level and trend shift occurs in $T_B + 1$. Details about the test procedure can be found in Zivot and Andrews (1992) or in Chumacero and Quiroz (1996). The main results of these tests are reported in Table 3. In this table we can see that there are some regions for which we could reject the unit root hypothesis. These are the regions IV, V, IX, X and XII.

TABLE 3

ZIVOT AND ANDREWS UNIT ROOT TESTS FOR REGIONAL GDP PER CAPITA

	I	II	III	IV	V	VI	VII	VIII	IX	X	XI	XII	XIII
$inf\ t_\alpha$	–2.26	–4.49	–2.43	–6.61	–5.86	–4.25	–3.52	–3.57	–5.23	–5.71	–2.94	–7.31	4.43

The asymptotical critical value at the 95% confidence level is –5.08 and at the 99% confidence level is –5.57. The lag truncation procedure is as described in Zivot and Andrews (1992).

However, there are others regions where the evidence is not strong, and also, it is possible that the asymptotic critical values that we use have size distortions with respect to the corresponding finite sample values. For these reasons we proceed to apply the cointegration tests to all possible pairs of regions. Now, however we consider the possibility of level shifts in the ratios of GDP per capita between regions as the structural change arguments above suggest. To test for cointegration we used the testing procedures described by Perron and Vogelsang (1992a). In these tests the null hypothesis is that $z_t^{i,j}$ follows this process:

(3) $$z_t^{i,j} = \gamma D(TB)_t + z_{t-1}^{i,j} + e_t$$

with $D(TB)_t = 1$ if $t = T_B + 1$ and 0 otherwise; and e_t following the process $ARMA(p, q)A(L)e_t = B(L)w_t$, with $w_t\, iid(0, \sigma^2)$. Under the alternative hypothesis, $z_t^{i,j}$ is described by:

(4) $$z_t^{i,j} = \mu + \gamma DU_t + v_t$$

where $DU_t = 1$ if $t > T_B$ and 0 otherwise; v_t follows a process $ARMA(p+1, q)$ which can be written as $(1 - \alpha L)A(L)v_t = B(L)w_t$. The null hypothesis is a special case of the alternative with $\alpha = 1$ (see details in Perron (1990), Perron and Vogelsang (1992a) and (1992b)). Perron and Vogelsang (1992a) propose a two-step testing procedure: firstly, (4) is estimated by ordinary least squares to obtain the estimated residuals v_t^*, then, in the second step, an ADF type unit root test is performed based on the stadigraph $inf_{TB\in(p+2,T)}(t_\alpha)$, where the null hypothesis is that $\alpha = 1$ in the following regression:

(5) $$v_t^* = \sum{}^k_{i=0}\left(\varpi_i D(TB)_{t-i}\right) + \alpha v_{t-1}^* + \sum{}^k_{i=1}\left(c_i \Delta v_{t-i}^*\right) + u_t$$

The asymptotical critical values for $inf_{TB\in(p+2,T)}(t_\alpha)$ under the null hypothesis have been obtained by Perron and Vogelsang (1992a). The main results of the application of this test are reported in Table 4[7]. As can be seen in this table, the cointegration results are similar to those found with the simple ADF tests. However, now we find that Region XIII belongs to the first *cointegrating group* and that Region I belongs to the second *cointegrating group*. The differences in the results can be explained by the higher power of Perron and Vogelsang (1992a) tests in the presence of level shifts. As can be seen in Figure 1, the pairs of regions that didn't show cointegration with the ADF tests, but that showed cointegration with Perron and Vogelsang (1992a) tests, exhibit a level shift in their ratios. The procedure that we follow for testing unit roots suggests the T_B period that minimizes t_α as a natural estimator for the date of the shift. Using this estimator, the structural-change date estimators are the year 1967 for the ratio between Region I and Region XI; 1974 for Region XIII and Region II; 1973 for Region XIII and Region IV; and 1974 for Region XIII and Region VII. However, it is important to emphasize that the above methodology attempts to test for unit roots rather than estimate date shifts. In particular, neither the asymptotical nor the finite sample statistical properties of the estimators of shift dates are known.

Finally, it is worth noting that we again found that none of the regions of the first group cointegrated with any region of the second group and that regions III and XII don't cointegrate with any other region.

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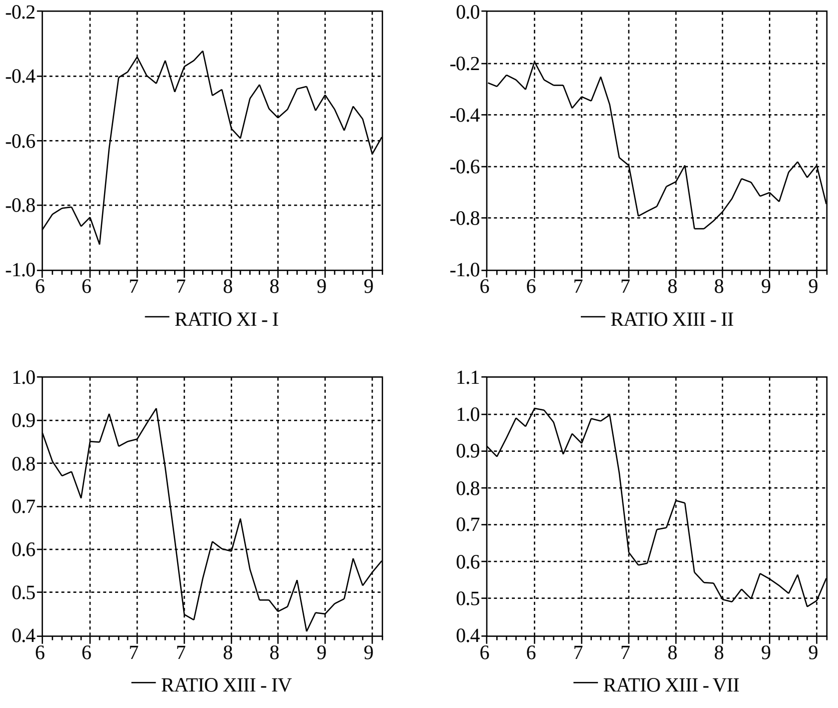
[7] All details about the estimations are available upon request.

TABLE 4
PERRON AND VOGELSANG (1992) TEST FOR COINTEGRATION

	II	III	IV	V	VI	VII	VIII	IX	X	XI	XII	XIII
I	–3.18	–2.79	–3.54	–2.38	–3.12	–2.66	–2.02	–3.58	–2.86	–5.13*	–2.27	–2.82
II	//	–2.64	–3.45	–3.93	–1.78	–3.67	–2.08	–3.52	–2.94	–3.84	–1.65	–4.71*
III		//	–2.59	–1.62	–2.75	–2.67	–2.28	–2.35	–2.94	–2.87	–1.72	–1.90
IV			//	–3.02	–5.29*	–4.53*	–4.41*	–5.23*	–5.18*	–3.61	–1.96	–4.86*
V				//	–3.29	–3.01	–3.99	–2.56	–2.65	–5.83*	–3.27	–2.41
VI					//	–4.00	–2.77	–3.99	–3.90	–3.45	–3.26	–3.43
VII						//	–5.21*	–3.72	–3.80	–2.55	–2.47	–4.42*
VIII							//	–3.40	–3.57	–3.01	–1.61	–3.26
IX								//	–4.78*	–3.09	–1.87	–3.28
X									//	–2.78	–2.09	–3.87
XI										//	–3.87	–4.10
XII											//	–2.57

Rejection of the unit root hypothesis at 90% (or higher) confidence level is denoted by *. The asymptotical critical value for this level of confidence is –4.19 and was obtained from Table 1 in Perron y Vogelsang (1992a).

FIGURE 1
RATIOS AND LEVEL SHIFTS FOR REGIONS COINTEGRATING AFTER A SHIFT



4. Conclusions

The main objective of this paper is to characterize the statistical properties of the series of GDP per capita ratios between the regions of Chile. In particular, we were interested in knowing whether or not these ratios have a unit root, and therefore whether they tend to revert to any deterministic constant. These properties have been interpreted in the literature as evidence in favor (if series are stationary), or against the neoclassical model of growth (see Bernard and Durlauf (1996)).

When we applied ADF unit root tests we found that there were two groups of regions cointegrating within themselves, but not cointegrating with regions of the other *cointegrating group*. Also we found that there is a third group of regions, consisting of Regions I, III, XII and XIII, that didn't cointegrated with any other region.

When we incorporated the shift level alternatives into our testing procedures, as described by Perron and Vogelsang (1992a), we found that Region I belonged to the first *cointegrating group* and Region XIII cointegrated with the second *cointegrating group.* However, we continue to find that the two groups don't cointegrate between each other and Regions III and XII don't cointegrate with any other region. The structural change related to the testing procedures would have occurred in the mid 1960s in the case of Region XI and in the early 1970s for the case of Region XIII.

However, there are other interpretations for the above results. Firstly, it is still possible to explain the no-cointegration results as a consequence of the low power of the testing procedures that we applied. However, we think that regions that don't cointegrate show a very different pattern with respect to the others regions. This is particularly obvious for Region XII that had a negative average rate of growth in all the period 1960-1996. Appendix 1 contains the main statistics of regional growth in the period. Secondly, the no cointegration result can be explained, particularly for the case of Region XII with the other regions, by the initial conditions of Region XII that in 1960 was the richest region of the country. In the neoclassical framework it could be interpreted that Region XII was over the steady state at the beginning of the period. This interpretation would explain the differences between the results above and the conclusions in Fuentes (1997). As explained in detail in Proposition 6 of Bernard and Durlauf (1996), the results of cross country convergence tests are not compatible with the results of time series tests of convergence under certain conditions. This is because cross country type tests require richer economies growing slower than poor economies and time series tests require similar growth rates.[8] However, the negative average rate of growth of region XII for a thirty-year period is puzzling in the neoclassical framework even for relatively richer economies.

8 See details in Bernard and Durlauf (1996).

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APPENDIX 1

REGIONAL GROWTH OF CHILE STATISTICS 1960-1996

Table A1 reports the main statistics of regional growth in Chile for 1960-1996 period.

TABLE A1
GDP PER CAPITA RATE OF GROWTH (%) 1960-1996

	I	II	III	IV	V	VI	VII	VIII	IX	X	XI	XII	XIII
Mean	1.91	3.69	3.72	3.07	1.65	2.27	3.19	3.46	2.84	2.83	2.73	–0.002	2.29
σ	6.56	7.79	8.49	5.80	5.36	6.10	4.74	6.52	5.42	4.36	7.07	7.38	6.60