PREDICTION OF THE ECONOMIC ACTIVITY FROM THE SHORT AND LONG TERM INTEREST RATE DIFFERENTIAL: NEW EVIDENCES IN CHILE AND THE UNITED STATES OF AMERICA CASES

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Abstract. The purpose of this work is to verify the stability of the relationship between real activity and interest rate spread. The test is based on Chen (1998) and Osorio and Galea (2006). The analysis is applied to Chile and the United States, from 1980 to 1999. In general, in both cases the relationship was statistically significant in early 80s, but a break point is found in both countries during that decades, suggesting that the relationship depends on the monetary rule follow by the Central Bank.

1. Introduction

Various authors have proposed predicting the economic cycle from the short and long term interest rate differential. There are different reasons behind this relationship. For example, the economic agents' expectations affects the interest rate differential. If the agents expect a recession to occur in one year, they will rise the demand on one-year-forward bonds, thus reducing that term interest rate; and they will rise the supply of short term bonds, rising the short rate. Therefore, if we observe a reduction on the slope of the interest rates forward structure today, a recession is expected to occur in one year. Another explanation is that the difference between short and long rates could be reflecting the monetary policy. Diminishing money supply would make the short term interest rates rise, but the reactions of long rates would not be proportional, thus diminishing the differential. Today restrictive monetary policy will have contractive effects in the following quarters. Thus, the decrease on the following year activity will be relate to today observation of the differential decrease.

Empirical works on this area have found consistent evidence of the predictive power of rate differential. With different methodologies, time periods, and countries, the results support the relationship. We can mention, among others, the works by Harvey (1989), Estrella and Hardouvelis (1991), and Estrella and Mishkin (1998) for the American case; Álvarez et al. (1994), and Fernández (2000) for the Chilean case[1].

The first three works analyze the American case, and the last two analyze the Chilean case. Harvey (1989) found that the short and long term interest rate

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[1]Evidence for other countries can be found in McMillan (2002), Estrella and Mishkin (1998).

differential explains a 30% of the variation of the United States' GNP growth rate between 1953 and 1989, while the stock price index explains only a 5%. Estrella and Hardouvelis (1991) conclude that the rate differential predicts the United States' GDP and its components in different periods, the results being more significant in a period shorter than two years. Estrella and Mishkin (1998) found that the rate differential predicts in a good way the economic activity in two-to-four-year periods, and that the performance of this variable is affected in a negative way when other variables are incorporated to the specification. They use data from the 1959-1995 period. The evidence in the Chilean case is limited. Álvarez et al. (1994) followed the methodology used by Harvey (1989) and found that the rate differential predicts the variation of Chilean GDP growth rate in the 1980-1990 period. Recently, Fernández (2000) used VARs and Granger causality test, and found that in a long term period, the relationship between the real interest rate and the economic activity variation is stronger than in a shorter period.

The purpose of this paper is to verify the stability of the relationship between the interest rate differential and the economic activity in the course of time. We wonder whether this relationship is regular with changes in the way the monetary policy is managed. We followed the methodology used and the evidence found by Harvey (1989) and Álvarez et al. (1994), and we analyze the American and Chilean cases between 1980 and 1999. On the one hand, we have a developed country which has a deep and mature capital market, the US; on the other hand we have a developing country: Chile. For the US, Estrella et al. (2002) found, among other things, that the relationship between industrial production and rate differential was not regular in a one-year period, with a breaking point occurring by the end of 1983, when changes in managing the US' monetary policies occurred.

Chile is a developing country where deep reforms were applied in the 70s, such as the liberalization of the capital market. Just by the beginning of the 80s we can talk about the existence of a capital market. Because of this fact, we focus our analysis in the 1980-1999 period. It is worth to mention that in 1982, Chile suffered a deep recession, after which some reforms were applied on the way the monetary policy was managed, including leaving out the exchange parity and focusing monetary policy on inflationary targets through managing the interest rate.

This paper is organized as follows. In Section 2, we present the methodology and data we used for estimating the relationship (Harvey, 1989), as well as for detecting whether a change-point exists (Chen, 1998). The test used for detecting the change point is different from the ones used by Estrella et al. (2002). This gives us the possibility of counting on the results of this work as a control element for the methodology, as well as for the results for the Chilean case. In Section 3, we present the results of the estimations of the relationship between the interest rate differential and the production growth and the tests of change-point for both countries. Finally, in Section 4, we present our conclusions.

2. Methodology

In this section we briefly describe the model used for studying the stability of the relationship between the output variation and the rate differential.

2.1. **Independent t Regression Model.** The model used by Harvey (1989) establishes the relationship between the GDP growth rate and the interest rate differential. It is a simple regression model given by:

$$y_t = \beta_0 + \beta_1 x_t + \epsilon_t, \tag{2.1}$$

where y_t indicates the GDP growth from the $t+1$ quarter to the $t+5$ quarter, x_t is the difference between short and long term rates of return per year, given by $x_t = \log((1+Tl_t)/(1+Tc_t))$, with Tl and Tc being short and long term rates, respectively.

The random errors $\epsilon_1, \epsilon_2, \ldots, \epsilon_n$ are not correlated with a parameter of location zero and a parameter of scale ϕ. Following Lange et al. (1989) we assume that the random errors ϵ_t are iid $t(0, \phi, \nu)$, $\nu > 0$. The t distribution has its tails heavier than normal distribution and allows to diminish the outliers influence on the maximum likelihood estimator, allowing a robust inference process; see, for instance, Lange et al. (1989), and Taylor (1992). In this paper we will use the model proposed by Harvey (1989) and the t distribution for studying possible changes in the relationship between output growth and the interest rate differential in Chile and the US. That is to say, for estimating the parameters we will use the maximum likelihood method, assuming that $y_t \sim t(\beta_0 + \beta_1 x_t, \phi, \nu)$, $\nu > 0$, independent with density given by:

$$f(y_t) = K(\nu)\frac{1}{\sqrt{\phi}}(1 + z_t^2/\nu)^{-\frac{1}{2}(\nu+1)} \tag{2.2}$$

with, $K(\nu) = \Gamma((\nu+1)/2)/\Gamma(\nu/2)(\pi\nu)^{\frac{1}{2}}$ and $z_t^2 = (y_t - \beta_0 - \beta_1 x_t)^2/\phi$.

We use the algorithm EM to obtain the maximum likelihood estimator (see McLanchlan and Krishnan, 1997).

2.2. **Detection of a change-point.** The regression model (2.1) assumes that the relationship between the variables does not change along all the period under study. However, because of the proper dynamics belonging to the economic processes, and/or because of the economic authorities' decissions, the relationship between the variables can change, improving (or harming) the predictive capacity of the Harvey Model (HM). In this paper we are interested in applying the HM on the data from the Chilean and US economy, and in detecting possible change points in the coefficients of the HM. In order to do so, we will use that methodology proposed by Chen (1998) and Osorio and Galea (2006), based on the Schwarz Information Criterion (SIC). In fact, let $(x_1, y_1), \ldots, (x_n, y_n)$ an observations sequence. Let us assume a simple linear regression model given by, $y_t = \beta_0 + \beta_1 x_t + \epsilon_t$, where random errors, ϵ_t are independent with parameters of location 0 and of scale ϕ. We can pose the problem of the change-point in the context of hypothesis test. We want to test the following null hypothesis:

$$H_0 : y_t = \beta_0 + \beta_1 x_t + \epsilon_t, \quad t = 1, 2, \ldots, n, \tag{2.3}$$

that is to say, there is no change in the regression coefficients, against

$$\begin{aligned} H_1 : y_t &= \beta_0^1 + \beta_1^1 x_t + \epsilon_t, \quad t = 1, \ldots, k, \\ y_t &= \beta_0^2 + \beta_1^2 x_t + \epsilon_t, \quad t = k+1, \ldots, n. \end{aligned} \tag{2.4}$$

That is to say, there is a change (in the regression coefficients) in an unknown position k, called *change-point*.

Following Osorio and Galea (2006), we will assume that random errors ϵ_t are *iid* $t(0,\phi,\nu)$, $\nu > 0$. In our case, SIC under H_0 is given by,

$$\begin{aligned} SIC(n) &= -2\,n\log K(\nu) + n\,\log\widehat{\phi} + 3\log n \\ &\quad + (\nu+1)\sum_{t=1}^{n}\,\log\{1 + z_t^2(\widehat{\boldsymbol{\beta}},\widehat{\phi})/\nu\} \end{aligned} \tag{2.5}$$

and under H_1 (given k), the SIC is given by

$$\begin{aligned} SIC(k) &= -2\,n\log K(\nu) + n\,\log\widehat{\phi} + 5\log n \\ &\quad + (\nu+1)\Big[\sum_{t=1}^{k}\log\{1 + z_t^2(\widehat{\boldsymbol{\beta}}_1,\widehat{\phi})/\nu\} + \sum_{t=k+1}^{n}\log\{1 + z_t^2(\widehat{\boldsymbol{\beta}}_2,\widehat{\phi})/\nu\}\Big], \end{aligned} \tag{2.6}$$

where, $z_t^2(\widehat{\boldsymbol{\beta}},\widehat{\phi}) = (y_t - \widehat{\beta}_0 - \widehat{\beta}_1 x_t)^2/\widehat{\phi}$, and $z_t^2(\widehat{\boldsymbol{\beta}}_j,\widehat{\phi}) = (y_t - \widehat{\beta}_0^j - \widehat{\beta}_1^j x_t)^2/\widehat{\phi}$, $j = 1, 2$. Observe that when $\nu \to \infty$, we obtain the expressions corresponding to the normal case (Chen, 1998).

With implementing the SIC, we change from our proving hypothesis task to a model selection hypothesis. The null hypothesis belongs to a regression model with no change in the coefficients, and the alternative hypothesis represents regression models with a change-point in the k position.

Using SIC on equations (2.5) and (2.6), the hypothesis of no change in the regression coefficients H_0 is rejected if:

$$SIC(n) > \min\{SIC(k) : 2 \le k \le n-2\}.$$

If H_1 is accepted, that is to say, if there is a change-point in the regression coefficients, the position of the change-point, $\widehat{k}$, satisfies:

$$SIC(\widehat{k}) = \min\{SIC(k) : 2 \le k \le n-2\}.$$

For further details, see Osorio and Galea (2006).

3. Application

In this section we apply the Harvey's model to quarterly period data about Chile and America's output growth and the interest rate differential between 1980 and 2000. Data can be obtained on request.

3.1. **Chilean Data.**

Figure 1 shows the scatter plot of Chilean data. We can observe a certain linear tendency and an increase of the output growth volatility depending on the interest rate differential.

We applied SIC in order to detect possible change points in the independent t regression model with known degrees of freedom and also in the normal model. We have that the model that gives the best description of the data is the t with four degrees of freedom ($SIC(n) = -215.3$, $\min SIC(k) = -245.7$, $\widehat{k} = 6$), which is also confirmed by the QQ plots not shown here. Then we have a possible change point on the position $k = 6$ (fourth quarter of 1981). Then, the coefficients change from the place number 7 (first quarter of 1982). Figure 1 shows the scatter plot of data and the fitted regression lines, before $\triangle$ (—) and after $\circ$ (– –) the change point.

Table 1 shows the results of the model fitted with a change on the position 6. The approximate standard errors of the estimators are shown in parenthesis (see Lange et al., 1989). Observe that the detected change-point is highly significant ($c_{0.10} = 4.67599$, $c_{0.05} = 7.92304$ and $c_{0.01} = 16.57939$ see condition (5) in Osorio and Galea (2006)), for both the t model with 4 degrees of freedom and the normal model. Also observe that the slope goes from significant to non-significant. Figure 1 shows fitted regression lines, before and after the change-point.

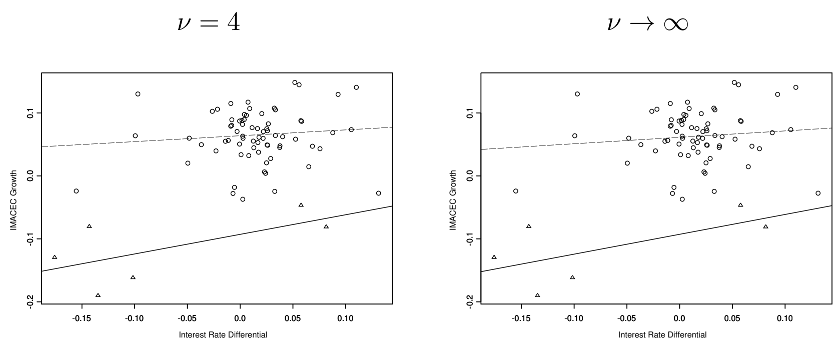

FIGURE 1. Scatter plot and regression models before $\triangle$ (—) and after $\circ$ (– –) change-point. Chilean data.

TABLE 1. Regression model with change-point in $k = 6$. Chilean data

t Model	quarter	Estimate β_0	Estimate β_1
$\nu = 4$	1 to 6	-0.09288	0.31134
		(0.02437)	(0.19917)
	7 to 77	0.06363	0.09338
		(0.00617)	(0.13139)
$\nu \to \infty$	1 to 6	-0.09280	0.31530
		(0.02842)	(0.23227)
	7 to 77	0.06109	0.10204
		(0.00720)	(0.15323)

3.2. **US Data.** Figure 2 shows the scatter plot of US data. Observe that, the same as the Chilean case, certain linear tendency is observed with an increase of the product growth variability depending on the interest rate differential.

We applied SIC in order to detect possible change points in the independent t regression model with known degrees of freedom and also in the normal model. In this case, the model that gives the best description of the data is the t with two degrees of freedom ($SIC(n) = -397.4$, $\min SIC(k) = -418.5$, $\widehat{k} = 15$), which is also confirmed in the quantile plots (not shows here).

Table 2 shows the results of the model fits with a change on the position 15. The approximate standard errors of the estimators are shown in parenthesis. We noticed that the t model with $\nu = 2$ degrees of freedom detects a change in a highly significant way; on the other hand, the change detected in position 15, third quarter

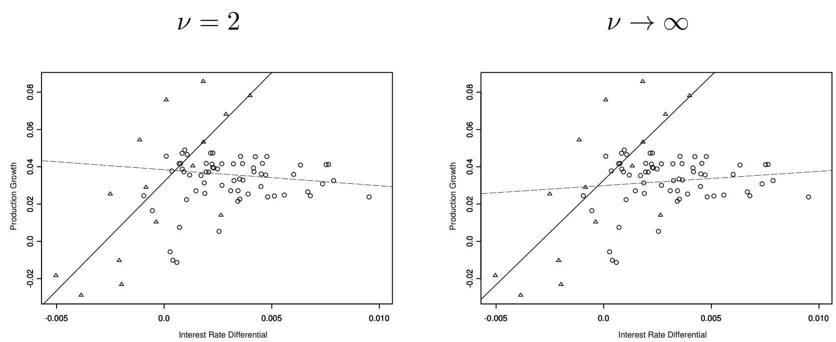

FIGURE 2. Scatter graph and regression models before △ (—) and after ○ (– –) change-point. US data.

1983, for the normal model is significant at 5%, ($c_{0.10} = 4:67599$, $c_{0.05} = 7.92304$ and $c_{0.01} = 16.57939$). Besides, the slope (β_1) goes from highly significant to statistically non-significant. Figure 2 shows fitted regression lines, before △ (—) and after ○ (– –) change-point.

TABLE 2. Regression model with change-point in $k = 15$. US data.

t Model	quarter	Estimate	
		β_0	β_1
$\nu = 2$	1 to 15	0.03206	11.70656
		(0.00353)	(1.39209)
	16 to 79	0.03840	-0.85910
		(0.00288)	(0.74937)
$\nu \to \infty$	1 to 15	0.03284	11.21219
		(0.00493)	(1.94253)
	16 to 79	0.02988	0.76351
		(0.00402)	(1.04567)

4. CONCLUSIONS

From this analysis, we have two conclusions. First of all, we noticed that the relationship between interest rate differential and output growth is not regular through the time. Specifically, in both countries there is at least one change point, from which the model does not predict any more ($\widehat{\beta}_1$ tends to zero). Therefore, the HM has lost its capacity to predict the growth rate.

Second, it is not possible to completely associate the lost of anticipating the economic cycle from the rate differential with the change in managing the monetary policy, in a context in which the Central Bank directly influence the interest rates. In the Chilean case, the breaking point of the relationship (fourth quarter of 1981) occurs four years before the Central Bank began to manage the monetary policy with the interest rate indexed by the CPI. In the American case, we found some relationship. While the breaking point occurred in the fourth quater of 1983, the policy change occurred in October 1982, exactly one year before the statistic break that gathers the anticipation of the interest rate differential. However, in no case this analysis allows to attribute the weakness in the theoretical relationship to the change in the monetary rule.

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