BANCO CENTRAL DE CHILE

DETERMINANTS OF THE CHILEAN SOVEREIGN SPREAD: IS IT PURELY FUNDAMENTALS?

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ABSTRACT

In recent years, the Chilean economy has been widely recognized at the international level for the seriousness and soundness of the policies applied, achieving high prestige due to its firm commitment to financial and macroeconomic stability. In the last decade the country has developed a solid position in terms of external liquidity and solvency, and has steadily reduced inflation within the context of a healthy financial system and balanced fiscal accounts. All of these elements have meant that analysts and other market participants have clearly differentiated between Chile and most other emerging economies. This phenomenon can be most clearly seen in the behavior of the sovereign spread of Chilean bonds, which averaged 180 basis points from 1999 to 2002, despite the advent of the crisis in Argentina in 2001/2002. Therefore, a study on the behavior of the Chilean sovereign spread becomes relevant. As such, the purpose of this paper is twofold: the first is to study the time series properties of the sovereign spread, in order to be able to determine whether time series and/or GARCH type of models seem suitable for the purposes of modeling the behavior of the spread in the short term, with the purpose of forecasting the future path of this variable. The second is to study the extent to which fundamental variables of the Chilean economy as well as external factors determine the medium to long term behavior of the Chilean sovereign spread. The main contribution of this paper is to quantify the determinants of the cost of external funding for an emerging market economy such as Chile, which has a rather recent history in terms of issuance of sovereign bonds.

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I.- INTRODUCTION

The study of the behavior of the sovereign spread has been a permanent interest of the Chilean economic authorities since the issuance of the first bond in April 1999. This paper in part fulfills that need through the study of the determinants of the sovereign spread through two approaches. The first one involves the use of time series models to capture the main characteristics of the stochastic processes behind the behavior of Chile's sovereign spread, by making use of daily data. The second approach involves the study of the determinants of the level of the sovereign spread, based on fundamental variables that according to theory and previous empirical literature should influence the sovereign spread. This second approach is more limited than the first, due to the rather short history of the sovereign spread, and the monthly frequency of fundamental variables data.

The paper has the following structure: section II briefly describes the different issuance's of sovereign debt by the Chilean government in terms of amounts and general conditions, and compares the evolution of Chile's sovereign spread with other emerging markets with Investment Grade rating and other countries of the Latin-American region. Section III is dedicated to the time series analysis of the spread, in terms of stationarity, ARCH effects, and the estimation of traditional GARCH models and Asymmetric GARCH models. In section IV we carry on the fundamentals analysis, first identifying the variables that help determine the level of the spread and then estimating the corresponding models. Section V presents the conclusions.

II.- CHILE'S SOVEREIGN DEBT

The Chilean government undertook its first bond issue in early 1999 as a way to achieve several objectives. At first, sound and solid Chilean macroeconomic fundamentals had set a favorable environment for sovereign approaches to international markets. Indeed, one of the factors making suitable to direct public financing toward sovereign debt issuance was the extremely positive Chilean assessment by international investors. Since 1995, Chile has received A- rating by Standard & Poor's (for long term debt in foreign currency) keeping a stable outlook in each review exercise. All of these factors provided favorable bias toward external issues, which involved both a positive scenario for a first sovereign debt issuance and an attractive means for adequate management of public debt structure.

Chilean corporations had begun debt issuance in late 1993 (Compañía Sudamericana de Vapores), process that has continued during the last years. Indeed, local companies have accumulated issues for around US$ 7.5 billion from 1993 up-to-date, which is a significant amount taking into account that total private external debt is around US$ 32 billion. Given these developments of private issues, a relevant factor behind the first sovereign bond issuance was to set a public benchmark for future corporate debt announcements. In fact, corporate debt had not shown a high liquidity in international markets, which made important releasing a reference value to institutional investors, in particular, those focused on investment grade country assets.

Further, most of emerging markets had the chance to compare the private to public cost of external financing, since they had already issued sovereign bonds, presenting a guideline for corporations in order to assess the best timing for a bond issue. Thus, investment banks used to follow the premium paid by local companies over sovereign spreads as a overall measure of financial soundness of an economy.

As mentioned above, the first issue of sovereign bond made by the Chilean government was carried out in April 1999, which involved US$ 500 million, a release spread of 169 basis points, and a ten year maturity. The following issue was done in mid-October 2001, weeks after September 11 events in U.S., amid high overall market volatility. The government had already begun the issue process in the second half of 2001 and decided to continue it, regarding the stage achieved at that moment. In the end, the US$ 650 million bond was allocated to high grade investors and registered a release spread of 256 basis points, considered a well assessed auction. This bond presents a ten year maturity.

Finally, the latest issue was carried out in April 2002, which involved not only a new U.S. dollar denominated bond, but also expanded government issues to European markets. Thus, the authority issued a US$ 600 million five year bond and a € 300 million three year bond. Based upon these new bonds, the Chilean government extended the yield curve for international instruments, covering five, seven and ten years horizon denominated in dollars. Figure 1 presents the sovereign spread series for the US dollar bonds.

EMERGING MARKETS SPREADS: INVESTMENT GRADE ECONOMIES

The Chilean economy has outperformed Latin-American countries since mid 1980's, in particular, its strong macroeconomic fundamentals and institutional stability has meant a stable country-risk rating by both Moody's (Baa1) and Standard & Poor's (A-) during the period under study in this paper.

Indeed, if we compare countries with investment grade category and other Latin-American economies, namely Argentina and Brazil, we find meaningful gaps between sovereign spread levels in the period began in early 1999 up-to-date. As it is shown in Table 1 and Figure 2, Investment Grade countries present spreads in a range of 160–230 basis points, considering the mean column. On the other hand, Latin-American countries show higher levels, moving in a range of 900–1700 basis points in terms of the mean spread. Furthermore, if we look at maximum levels reached by both groups, we realize that the poorest performer from investment grade economies, i.e. Poland, slightly exceeded 300 basis points in this sample, whereas the poorest performer from Latin America, i.e. Argentina, surpassed by far 7000 basis points. However, it has to be kept in mind the fact that Argentina has declared selective default for sovereign debt in late 2001.

Table 1: Sovereign Spreads Investment Grade Economies

COUNTRY	MEAN	STD. DEV.	MEDIAN	MAX.	MIN	OBSERVATIONS (04/99-07/02)
			INVESTMENT GRADE			
Chile	178	32	179	245	86	846
South Korea	175	48	189	272	74	860
Israel	164	20	167	217	99	600
Malaysia	162	15	158	195	121	120
Poland	228	36	234	307	148	860
			LATIN AMERICA			
Argentina	1678	1722	801	7199	515	860
Brazil	865	186	821	1727	626	860

III.- TIME SERIES ANALYSIS OF THE SOVEREIGN SPREAD

The data availability on the sovereign spread for the different sovereign bonds issued by Chile depend on the corresponding issuance date of each bond. For the 2009 bond, the data starts on April 21 1999, and for the 2012 bond the data starts on October 23 2001. Table 2 presents the descriptive statistics for each bond for the corresponding sample period[1].

Table 2: Descriptive Statistics of Chilean Sovereign Bond Spreads

VARIABLE	MEAN	STD. DEV.	SKEWNESS	KURTOSIS	MINIMUM	MAXIMUM	NO. OBS. (TRADING DAYS)
CHI 2009	177.9	31.94	-0.499	3.478	86	245	801
CHI 2009	141.8	32.12	0.061	2.128	86	219	178
CHI 2012	169.5	33.85	0.584	2.530	121	253	178

Throughout the paper, the analysis will be centered in the 2009 bond, since it is the bond for which we have a longer history, and therefore a larger number of observations and time span. The average spread for the 2009 bond was 178 basis points, reaching a maximum of 245 basis points on April 14 2000, and minimum of 86 basis points of March 28 2002. A more complete characterization of the distribution of the data series can be made by calculating the skewness and kurtosis statistics. By computing this statistics we are able to compare the sample values with those of the normal distribution. Under the normal distribution, the value of the skewness statistic[2] should be zero, because the normal distribution is a symmetric distribution, and the kurtosis statistic[3] should be equal to 3. Therefore, if the estimated statistics from the data differ from these values, it is indicative of the fact that the distribution of the sovereign spread departs from the normal distribution. For the 2009 bond we get a value –0.5 for the skewness statistic, which provides evidence of a distribution that has a long left tail, so that the spread is more likely to be far below the 178 basis points mean than above it.

The usefulness of the kurtosis statistic is that it measures the peakedness or flatness of the distribution of implied volatilities. As mentioned above, we need to compare the estimated kurtosis statistic with the value of 3, in order to know if the distribution is leptokurtic (values of kurtosis greater than 3) or platykurtic (values of kurtosis less than 3). For the 2009 bond we observe that the calculated kurtosis statistic is 3.5, above the critical value of 3, which indicates that the distribution

[1] The sample ranges from each starting date up to July 15 2002.

[2] Skewness: $S = \frac{1}{T}\sum_{t=1}^{T}\left(\frac{X_t - \overline{X}}{\hat{\sigma}}\right)^3$

[3] Kurtosis: $K = \frac{1}{T}\sum_{t=1}^{T}\left(\frac{X_t - \overline{X}}{\hat{\sigma}}\right)^4$

is leptokurtotic, so that the distribution has more mass in the tails than a normal distribution, what is usually called fatter tails.

To test whether the series are normally distributed or not, we can make use of the Jarque–Bera test statistic. Under the null hypothesis of a normal distribution the Jarque–Bera statistic[4] is distributed as a χ^2 with 2 degrees of freedom. The advantage of this test is that it is a joint test, since it measures the difference of the skewness and kurtosis of each series of implied volatilities with those from the normal distribution. The estimated value of the Jarque–Bera statistic is 40.9, so we reject the null hypothesis of normality for the 2009 spread at the 1% significance level.

In order to check the time series properties of the data, we need to estimate the autocorrelation (ACF) and partial autocorrelation (PACF) functions to explore the possibility of fitting an traditional time series model to the sovereign spread data. Table 3 below presents the ACF and PACF for the spreads of the 2009 and 2012 bonds.

Table 3: ACF and PACF, Sovereign Spreads

	2009 Bond			2012 Bond	
Lag	ACF	PACF	Lag	ACF	PACF
1	0.990	0.990	1	0.976	0.976
10	0.887	– 0.011	10	0.657	0.059
20	0.774	– 0.007	20	0.355	– 0.002
30	0.646	– 0.072	30	0.187	– 0.074
40	0.491	0.001	40	0.047	0.026
50	0.359	– 0.031	50	– 0.051	0.053
60	0.251	– 0.072	60	– 0.148	– 0.051

The autocorrelation function for the 2009 bond exhibits a very slow decay, which is indicative of some high degree of persistence in the series, since after 60 lags, the effect of a shock to the spread is still present. As a way of complementing the persistence exhibited by the autocorrelation function, we can calculate the half-life [5] of a shock to the sovereign spread. The half life allows us to have an idea about how much time does it take for the sovereign spread to reduce to a half the impact of a shock. A large half–life value means that the process is very persistent, so that any shock to the implied volatility takes a long time to die out (as would be in the random walk case). A low half–life value means that the time it takes for a shock to reach half of its original level is shorter, indicative of lower persistence in the process. For the 2009 bond, we get a half life of 40 trading days, that is, it takes 8 weeks to dissipate half of the original shock. On the other hand, for the 2012 bond, we get a half life of 15 trading days, so it takes 3 weeks to dissipate half of the

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4 Jarque–Bera: $JB = \frac{N-k}{6}\left(S^2 + \frac{1}{4}(K-3)^2\right)$; H_0: Sovereign Spread ~ Normal; Under H_0: JB ~ χ^2_2

5 Half-life was calculated solving the following equation: $(\gamma_1)^h = 0.5$

original shock. It should be noted that the samples for the 2009 and 2012 bonds are different, and therefore it is not surprising to find different half lives.

STATIONARITY

In order to check whether the sovereign spreads are stationary or not, we ran a series of unit root tests to check whether the series were stationary. A series is stationary if the mean, variance and covariances are constants and do not change over time. If this result holds for spreads series, we can say that the series are weakly stationary or covariance stationary[6]. The relevance of checking for stationarity has to do with shock persistence, in the sense that for a stationary series, a shock to the series has no permanent effect. On the other hand, if we have a non-stationary series, we will find that a shock to the series will actually have a permanent effect.

The unit root tests we ran include the Augmented Dickey Fuller (ADF Tests), the ADF–GLS Tests by Elliot, Rothemberg and Stock, and the KPSS Test. Table 4 below presents the results of the ADF, under two different lag selection criteria, namely the AIC and SIC. The traditional unit root tests fail to reject the presence of a unit root in the series, under both lag selection criteria. Given the low power of the traditional unit root tests against the local alternative of a root close to, but below unity, we ran the ADF–GLS test, which is the most powerful invariant test against the local alternative.

Table 4: ADF Unit Root Tests

	MINIMIZES AIC				MINIMIZES SIC			
VARIABLE	ADF (μ)	LAG (P)	ADF (τ)	LAG (P)	ADF (μ)	LAG (P)	ADF (τ)	LAG (P)
CHI 2009	– 2.203	3	– 2.537	3	– 1.977	0	– 2.285	0

† : denotes rejection of hypothesis of a unit root at 1% significance level.
‡ : denotes rejection of hypothesis of a unit root at 5% significance level.
* : denotes rejection of hypothesis of a unit root at 10% significance level.

From the results given in Table 5 below, we can see that the ADF–GLS gives us some evidence that the spread of the 2009 bond exhibits level stationarity, since the tests rejects the presence of a unit root at the 5% significance level.

Table 5: ADF–GLS Unit Root Tests

	MINIMIZES AIC				MINIMIZES SIC			
VARIABLE	ADF (μ)	LAG (P)	ADF (τ)	LAG (P)	ADF (μ)	LAG (P)	ADF (τ)	LAG (P)
CHI 2009	– 2.139‡	3	– 2.158	3	– 2.139‡	3	– 2.158	3

† : denotes rejection of hypothesis of a unit root at 1% significance level.
‡ : denotes rejection of hypothesis of a unit root at 5% significance level.
* : denotes rejection of hypothesis of a unit root at 10% significance level.

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6 Formally, for a time series X_t to be coavariance stationary, we need that the following conditions hold: $E(X_t)=\mu, \forall \text{ t}$, $V(X_t)=\sigma^2, \forall \text{ t}$ and $COV(X_t, X_{t-k}) = f(k) \;\; \forall \text{ t,k}$

In terms stationarity, the last test corresponds to the Kwiatowski, Phillips, Shin and Schmidt (KPSS) test, which is one of the few tests that has a null of stationarity. The results presented in Table 6 presents evidence in favor of level stationarity for the spread series for a lag truncation parameter above 40 trading days.

Table 6: KPSS Tests

		Lag Truncation Parameter (l)								
Variable	Test	1	10	20	30	40	50	60	70	80
CHI 2009	η_μ	15.72†	1.59†	0.84†	0.59‡	0.47‡	0.39	0.35	0.32	0.30
	η_τ	6.81†	0.69†	0.37†	0.26†	0.21‡	0.18‡	0.16‡	0.15‡	0.14

η_μ : corresponds to the test under the null that the series is level stationary

η_τ : corresponds to the test under the null that the series is trend stationary.

† : denotes rejection of hypothesis of stationarity at 1% significance level.
‡ : denotes rejection of hypothesis of stationarity at 5% significance level.

In summary, the traditional unit root tests fail to reject the presence of a unit root for the 2009 sovereign bond spread. However, the ADF–GLS provides evidence that the series is stationary around a certain level, result that is corroborated by the KPSS test, once a we consider a long lag truncation parameter for the purposes of estimating the long run variance of the series.

ARCH Effects

The results obtained from the previous section, justify the need for a more parsimonious model in order to explain the behavior of the sovereign spread. Therefore, the following section will be devoted to examine the existence of ARCH effects in the 2009 bond spreads, since from the observation of the series, we can see that it exhibits several episodes where positive (negative) shocks seem to be followed by positive (negative) shocks for some periods of time, generating several clusters of up or downswings in the series that might be better captured by ARCH type models.

In order to test for the presence of ARCH effects, we use the Lagrange Multiplier test of Engle (1982). Under the null hypothesis, there is no ARCH effect in spreads, and the alternative hypothesis is that ARCH effects are present. The test involves a two step procedure: the first step involves estimating the mean regression, which will consist of a simple AR(1) specification of the form: $Spread_t = \delta + \phi_1 Spread_{t-1} + \varepsilon_t$. From this regression we need to recover the series of estimated errors ($\hat{\varepsilon}_t$), which will be used in the second stage. The second stage involves regressing the square of the estimated error terms on a constant and q-lags of the square of the estimated error terms. From this second stage equation, we can test for the presence of ARCH effects by constructing the statistic T^*R^2 (numbers of observations (T) times the coefficient of determination (R^2)) which has a Chi-square distribution with q degrees of freedom. The results are presented in Table 7.

Table 7: Test for ARCH Effects on Sovereign Spreads

VARIABLE	T^*R^2 STATISTIC	LAGS (Q)	P-VALUE
CHI 2009	39.718	3	0.000000
CHI 2012	13.726	2	0.001046

† : denotes rejection of hypothesis of ARCH effects at 1% significance level.
‡ : denotes rejection of hypothesis of ARCH effects at 5% significance level.
* : denotes rejection of hypothesis of ARCH effects at 10% significance level.

GARCH(1,1) MODEL

In order to capture the ARCH structure of the errors, we will estimate an AR(1) model for the spread (mean equation) of the form:

$$Spread_t = \delta + \phi_1 \, Spread_{t-1} + \varepsilon_t$$

Given this specification, the unconditional mean spread will be:

$$E[Spread_t] = \frac{\delta}{1-\phi_1}$$

whereas the conditional mean spread will be $E[Spread_t|\Omega_{t-1}] = \delta + \phi_1 Spread_{t-1}$, where Ω_{t-1} corresponds to the information set at time t–1. The unconditional variance of the spread under the AR(1) specification corresponds to:

$$Var[Spread_t] = \frac{\sigma^2}{1-\phi_1^2}$$

However, under GARCH type model, we relax the previous assumption of constant conditional variance, and allow the conditional variance to vary over time, so that the conditional variance takes the following general form:

$$Var[Spread_t|\Omega_{t-1}] = \frac{h_t}{1-\phi_1^2}$$

Under GARCH(1,1) models the conditional volatility takes the following functional form:

$$h_t = \omega + \alpha_1 \varepsilon_{t-1}^2 + \beta_1 h_{t-1}$$

so we see that the conditional volatility depends on the square of the previous error term and on the previous conditional volatility. The interpretation of the ε_{t-1}^2 term corresponds to the news that have an impact on the conditional volatility. In terms of the sovereign spread series, good news correspond to negative shocks $(\varepsilon_{t-1}^2 < 0)$, since they would reduce conditional volatility, while bad news correspond to positive shocks $(\varepsilon_{t-1}^2 > 0)$, since they would increase conditional volatility. It should be noted that in standard GARCH models, the effect of a shock on conditional volatility depends only on its size, since the sign of the shock is irrelevant. As such, positive and negative shocks will impact conditional volatility in the same way. The parameters in this model should satisfy $\omega>0$, $\alpha_1>0$ and $\beta_1\geq 0$ to guarantee that $h_t\geq 0$. The GARCH(1,1) model is covariance–stationary if and only if $\alpha_1+\beta_1<1$. In this case, the unconditional variance of the errors is equal to:

$$\sigma^2 = \frac{\omega}{1-\alpha_1-\beta_1}$$

In order to better capture the existence of ARCH effects, we estimated a model consisting of an AR(1) specification for the spread, and a GARCH(1,1) specification for the error terms. The results presented in Table 8 show that the values of the parameters satisfy the requirements of covariance-stationarity, since $\alpha_1+\beta_1=0.1132+0.8098=0.923<1$. The value of the unconditional variance of the errors is equal to $\sigma^2 = \frac{\omega}{1-\alpha_1-\beta_1} = \frac{1.5563}{1-0.1132-0.8098} = 20.21$ basis points.

Table 8: AR(1), GARCH(1,1) Model for Sovereign Spread

VARIABLE	COEFFICIENT	Z-STATISTIC
	MEAN EQUATION	
Constant	170.1479	12.55
$Spread_{t-1}$	0.9898	225.79
	VARIANCE EQUATION	
Constant	1.5563	6.295
ε^2_{t-1}	0.1132	6.843
h_{t-1}	0.8098	31.449

As mentioned before, one of the most interesting aspects of GARCH type models is the fact that we can say something about the effect of the size of shocks on conditional volatility, which is captured by the News Impact Curve (NIC), introduced by Pagan and Schwert (1990) and popularized by Engle and Ng (1993). It basically measures how new information is incorporated into volatility. More precisely, it shows the relationship between the current shock or news ε_t and conditional volatility 1 period ahead h_{t+1}, holding constant all other past and current information. For the GARCH(1,1) model we see that the effect of either positive or negative shocks is symmetric, and so the sign of the shock does not affect the NIC. Figure 3 presents the NIC for the GARCH(1,1) model. We can see that a shock in t–1 of 10 basis points will have an impact of 29.4 basis points on the conditional variance of next period, whereas a 20 basis points shock in t–1 will have an impact of 63.2 basis points on the conditional variance of next period.

One of the caveats of the GARCH(1,1) model is that it does not allow for different responses of the conditional variance in terms of the sign of the shock, so that positive shocks will have a different impact on the conditional variance than negative shocks. In order to check whether positive and negative shocks have a different impact on the conditional variance, we ran the Engle and Ng (1993) test for asymmetric effects. In order to conduct the test, we let S^-_{t-1} denote a dummy variable that takes the value of 1 when $\hat{\varepsilon}_{t-1}$ is negative and 0 otherwise, where $\hat{\varepsilon}_t$ are the residuals from estimating a model for the conditional mean of the sovereign spread under the assumption of homocedasticity. The tests examine whether the squared residual $\hat{\varepsilon}_t^2$ can be predicted by S_{t-1}^-, $S_{t-1}^-\hat{\varepsilon}_{t-1}$ and/or $S_{t-1}^+\hat{\varepsilon}_{t-1}$, where $S_{t-1}^+ = 1 - S_{t-1}^-$ [7].

The results of the Engle and Ng tests are presented in Table 9. The first column corresponds to the Sign Bias test, which simply tests whether the magnitude of the square of the current shock (ε_t) depends on the sign of the lagged shock (ε_{t-1}). We can see that the sign bias is not significant, which means that the sign of the lagged shock has no significant impact on the magnitude of the shock. The second and third columns present the Negative Sign Bias and the Positive Sign Bias, respectively. These tests examine whether the effect of negative or positive shocks on the conditional variance also depend on their size. The tests show substantial evidence of asymmetric ARCH effects, since the negative sing bias and the positive sign bias tests show that the size of the either negative or positive shocks do affect the conditional variance differently. This result is corroborated by the last column, the general test, which consists of a joint test of the three previous measures of asymmetry.

Table 9 Test for Asymmetric ARCH Effects on Sovereign Spreads

VARIABLE	SIGN BIAS		NEGATIVE SIGN BIAS		POSITIVE SIGN BIAS		GENERAL TEST	
	TEST	P-VALUE	TEST	P-VALUE	TEST	P-VALUE	TEST	P-VALUE
CHI 2009	– 0.196	0.422	–14.99	0.000	34.173	0.000	665.06	0.000

The tests are applied to residuals from an AR(k) model, with k determined by the AIC.

NONLINEAR ASYMMETRIC GARCH MODELS

Given the evidence of the asymmetric ARCH effects, we need to make use of GARCH models that capture this asymmetry. There are several models that are able to capture this asymmetry, such as the Threshold ARCH (TARCH) model by Zakonian (1990), the Exponential GARCH model by Nelson (1991) and the GJR-GARCH model by Glosten, Jaganathan and Runkle (1993). Of the three models mentioned above, we will estimate the GJR-GARCH model in order to capture the asymmetric effects of positive and negative shocks on the conditional variance.

Under the GJR–GARCH model, the conditional variance takes the following functional form, that is obtained from the previous GARCH(1,1) model, but it assumes that the parameter of $\hat{\varepsilon}_{t-1}$ depends on the sign of the shock, that is:

$$h_t = \omega + \alpha_1 \varepsilon_{t-1}^2 (1 - I[\varepsilon_{t-1} > 0]) + \gamma_1 \varepsilon_{t-1}^2 I[\varepsilon_{t-1} > 0] + \beta_1 h_{t-1}$$

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[7] The test statistic are computed as the t-ratio of the parameter ϕ_1 in the regression: $\hat{\varepsilon}_t^2 = \phi_0 + \phi_1 \hat{w}_t - 1 + \xi_t$ where $\hat{w}_t$ is one of the three measures of asymmetry, so that $\hat{w}_t = \begin{cases} S_{t-1}^- \\ S_{t-1}^- \hat{\varepsilon}_{t-1} \\ S_{t-1}^+ \hat{\varepsilon}_{t-1} \end{cases}$

where $I[\cdot]$ is an indicator function. Under this specification, the conditions for nonnegativeness of the conditional variance (h_t) are $\omega>0$, $(\alpha_1+\gamma 1)/2\geq 0$ and $\beta_1>0$. The condition for covariance–stationarity is $(\alpha_1+\gamma 1)/2+\beta_1<1$[8]. Table 10 presents the results of estimating the GJR–GARCH model.

Table 10: AR(1), GJR–GARCH Model for Sovereign Spread

VARIABLE	COEFFICIENT	Z-STATISTIC
	MEAN EQUATION	
Constant	168.56	7.00
$Spread_{t-1}$	0.995	274.3
	VARIANCE EQUATION	
Constant	0.723	6.62
$\varepsilon_{t-1}^2(1-I[\varepsilon_{t-1}>0])$	-0.0075	–1.35
$\varepsilon_{t-1}^2 I[\varepsilon_{t-1}>0]$	0.1259	6.93
h_{t-1}	0.9055	76.20

From the values of the coefficients, we get that the unconditional variance under the GJR–GARCH model is equal to $\sigma^2=\dfrac{\omega}{\left(1-\dfrac{(\alpha_1+\gamma 1)}{2}-\beta_1\right)}=\dfrac{0.723}{1-\dfrac{(-0.0075+0.1259)}{2}-0.9055}=20.48$ basis points, slightly higher than the unconditional variance under the GARCH(1,1).

One of the most interesting aspects of GJR–GARCH model is the fact that we can differentiate the effect of positive and negative shocks on conditional volatility. For the GARCH(1,1) model we saw that the effect of either positive or negative shocks is symmetric, and so the sign of the shock does not affect the NIC. However, for the GJR–GARCH model negative shocks have quite a different effect on conditional volatility. In fact, the news impact curve from this model differs significantly form the one obtained for the GARCH(1,1) model. Figure 5 presents the NIC of the GJR–Model, where we see that negative shocks , i.e. that the spread in the current period is below the spread of the previous period, will tend to reduce conditional volatility, while positive shocks, i.e. that the spread in the current period is above the spread in the previous period, will tend to increase conditional volatility, at a faster rate than that predicted by the GARCH(1,1) model. So when the Chilean sovereign spread is rising, it basically becomes a more risky asset, since it will have a higher conditional variance, whereas a reduction in sovereign spread has a very significant effect in actually reducing the conditional volatility. This implies that under times of turbulence in the region, where the Chilean spread might follow upward trends of the sovereign spreads of either

[8] If this condition is satisfied, the unconditional variance of ε_t is: $\sigma^2=\dfrac{\omega}{\left(1-\dfrac{(\alpha_1+\gamma_1)}{2}-\beta_1\right)}$

Brazil or Argentina due to contagion effects, necessarily resulted in higher volatility, but once contagion effects passed, the Chilean sovereign spread quickly became less volatile.

As a way of quantifying these effects, Table 11 below presents the asymmetrical effects of positive and negative shocks on the conditional variance. We can see that a negative shock in t–1 of 20 basis points will reduce next period's conditional variance to 17.1 basis points, while a positive 20 basis points shock in t–1 will have an impact of 70.5 basis points on the conditional variance of next period. The same result can be seen by comparing Figure 4 and Figure 5, which present the conditional variance series for the GARCH(1,1) model and the GJR Model.

Table 11: Effect on Conditional Variance (h_t) of Shocks of Different Sign under GJR-GARCH Model

NEGATIVE SHOCK ($\varepsilon_{t-1} < 0$)		POSITIVE SHOCK ($\varepsilon_{t-1} > 0$)	
BASIS POINTS	IMPACT ON H_T	BASIS POINTS	IMPACT ON H_T
–20	17.103	20	70.462
–15	18.415	15	63.202
–10	19.352	10	32.692
–5	19.915	5	23.250

From the time series analysis of the spread series we can say that shocks seem to be very persistent, so that it takes several trading days for them to dissipate completely. However, the series seems to be stationary around a certain level, once we control for the local to unity unit root of the process. As several other financial time series data, the Chilean sovereign spread series exhibits excess kurtosis, which can be better captured by GARCH type models. In particular, asymmetric GARCH models, such as the GJR-GARCH have led us to the result that downward movements in the spread are followed by significantly lower volatilities than upward movements of the same magnitude.

The previous analysis has been based on the daily data of the sovereign spread, and involved mainly the use of time series tools. However, the level of the spread is mostly determined by macroeconomic fundamentals of each country. In the next section, we estimate a model that explicitly takes into account the influence of such factors on the level of the spread.

IV.- SOVEREIGN SPREAD FUNDAMENTALS ANALYSIS

Regardless of the benefits of making use of high frequency data in terms of modeling the behavior of sovereign spreads in the very short run, the main driving forces of the medium to long run spread levels charged to emerging market economies is essentially determined by macroeconomic fundamental variables. It is therefore necessary to analyze how these variables have determined the path followed by the Chilean sovereign spread. This section will be dedicated to explore the relationship between a set of relevant macroeconomic fundamental variables and the sovereign spread.

FUNDAMENTAL VARIABLES

The variables to be included as determinants of the sovereign spread can be grouped into three broad categories. The first category corresponds to variables related to Chile's external financial position. The second category includes measures of external and domestic performance, and the third category corresponds to international interest rates.

The first category of variables corresponds to Chile's external financial position, and the purpose of including these variables is to capture external investor's assessment on the country's position. It is important to differentiate between solvency problems and liquidity problems. Our focus will be on liquidity, rather than on solvency problems. In order to have a measure of potential liquidity problems, we need to take into account elements such as outstanding external debt, both public and private, external liquidity measured by the level of international reserves and debt composition in terms of its maturity. Countries with higher overall levels of external debt face higher spreads, while countries with lower levels of outstanding debt face lower spreads. The hypothesis behind this relationship is that an increase of external commitments involves higher pressure over external liquidity available in an economy. Increases in external debt should mean higher risk on the assets issued by the country increasing its debt, thereby forcing investors to require a higher yield on sovereign and corporate bonds.

But not only the overall level of external debt is relevant for the spread charged to a certain country. As important as the overall level of external debt is its maturity structure, since a higher concentration of short term debt is viewed as seriously compromising the country's international liquidity. Therefore, even though an economy might present an overall stable relative level of external debt, such as the one measured by the Total Debt to GDP ratio or Total Debt to international reserves ratio, a higher concentration in short-term debt would necessarily translate into a higher spread.

The second set of variables corresponds to variables related to the economic performance of the Chilean economy. Regarding external performance, the behavior of exports becomes a key indicator, since it reflects the country's ability to generate international resources, that in part might be used to serve the external debt. Chile's performance, as a small open economy, is highly dependent on the evolution of its export base, so that higher performance in terms of exports should lower the spread, as the country is able to generate a higher level of international resources. Regarding domestic performance, domestic growth should reduce sovereign spread levels due to several reasons. First, a higher level of domestic growth reflects a higher level of productivity relative to other emerging market economies, impacting positively the price of sovereign bonds, and reducing spread levels. Second, a higher level of domestic growth should also translate into higher revenues by the government, increasing the resources that might be used to service the debt, and also reducing the level of the sovereign spread.

The third set of fundamental variables correspond to international interest rates, and we will focus on the effects of U.S. interest rates[9]. In addition to the direct impact of changes in U.S. interest rates on rates in developing countries, sovereign spreads have tended to move in the same direction as the changes in U.S. interest rates. This effect on developing country spreads was seen clearly in 1994 when a tightening of U.S. monetary policy was reflected in a substantial widening of spreads, and in 1998, when an easing of U.S. monetary policy in response to the flight to quality

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[9] Part of this section was taken from Arora and Cerisola (2001).

and the concerns about a U.S. credit crunch associated with the Russian default and the near demise of Long-Term Capital Management (LTCM) helped to restore global liquidity conditions and to reduce sovereign spreads somewhat. Most of these analyses have tended to explore the role of global liquidity conditions, as proxied by a specific yield on a U.S. treasury security, on sovereign bond spreads.

From a theoretical perspective, a rise in U.S. policy interest rates could lead to an increase in emerging market spreads for several reasons. To the extent that emerging market bonds are risky (there is a probability of default), the yield on emerging market bonds would have to rise by more than any rise in the risk-free rate. To illustrate, if r and i represent the interest rate on the risk-free asset and the risky asset, respectively, and p is the probability of repayment on the risky asset, then the equilibrium condition is:

$$(1+r) = p \times (1+i) + (1-p) \times 0$$

The interest rate spread, S, defined as the difference between the rate on the risky asset and on the risk-free asset, in equilibrium is then:

$$S = (1+r) \times \frac{(1-p)}{p}$$

and its derivative with respect to r is $(1 - p) / p$, which is positive as long as $p < 1$. This says that as long as there is some risk of default, the rate on the risky asset will have to rise by more than any rise in the risk-free rate in order to compensate investors for the risk. A rise in U.S. rates could also raise emerging market spreads through its effects on the ability of debtor countries to repay loans. A rise in U.S. rates would tend to increase debt-service burdens in borrowing countries, which would reduce their ability to repay loans. In addition, as noted by Kamin and Kleist (1999), a rise in U.S. rates could reduce investors' appetite for risk, leading them to reduce their exposure in risky markets, in turn reducing available financial resources in borrowing countries. In terms of the above illustration, if the probability of repayment is a negative function of the risk-free rate ($p = p(r)$, with $p' < 0$), then the first derivative of S with respect to r is:

$$\frac{dS}{dr} = \left(\frac{(1-p)}{p}\right) - \left((1+r) \times \frac{p'}{p^2}\right)$$

which is positive (since $p < 1$ and $p' < 0$). This says that a rise in the risk-free rate raises the spread both because of the risk of default (the first term) and because that risk rises as the risk-free rate goes up (the second term). From a theoretical point of view, changes in U.S. interest rates, or likewise in global liquidity conditions, would be expected to influence positively country risk and sovereign spreads in developing countries.

In order to measure the effect of U.S. monetary policy we will make use of the federal funds rate, instead of the yield on a U.S. treasury security. Most of the specifications adopted so far in the literature have proxied U.S. monetary policy by the yield on U.S. treasury securities. However, shocks to U.S. treasury yields are not necessarily the result of changes in U.S. monetary policy.

Finally, in order to capture the potential effects of the new issuance's, and the effect it might have in the path followed by the spread, we included two dummies for each new issue, namely, 2012, 2007 and 2005 maturities. The purpose of this dummy variables is to capture an increase in the spread due to portfolio balance considerations, so that a larger amount of outstanding debt should necessarily imply a jump in the sovereign spread, since investors would be willing to hold a larger amount of debt only if they are compensated through a larger premium.

FUNDAMENTALS MODEL AND RESULTS

The model specification to be estimated can be summarized as follows:

$$Spread_t = \alpha_0 + \alpha_1(External\ Financial\ Position\ Variables)_t + \alpha_2(Performance\ Variables)_t$$
$$+\alpha_3(U.S.\ Interest\ Rates)_t + \alpha_4 Dummies + \varepsilon_t$$

The estimation process considered a number of alternative specifications, in particular, for international liquidity and external financial position. Several indicators related to the international financial position of Chile were included in the estimated equations. These included external debt interest payments, portfolio investment flows from non-resident and alternative international financial prices. However, since they were not statistically significant, they were are not reported. Table 12 presents the results of the estimation of four alternative specifications, which include different measurements for external financial position of the Chilean economy.

In terms of the variables related to Chile's external position, we see that the levels of either the total short term debt, or the total external debt are not statistically significant in determining the level of the Chilean sovereign spread. However, if we use the ratio of short term debt to international reserves, which corresponds to the most used indicator to measure international liquidity, we can see that it becomes significant in explaining the level of the sovereign spread, so that an increase in the short term debt to reserves ratio will necessarily imply a rise in the sovereign spread, due to the reduced international liquidity. The same result holds if we use the ratio of total external debt to international reserves as a measure of domestic international liquidity, since the parameter is now positive and statistically significant, but of a lower magnitude than the parameter obtained for the short term debt to reserves ratio, as shown by equations 2 and 4.

Regarding the effect of performance measures, external performance measured by exports has had an inverse impact on sovereign spread, so that a positive trend in exports causes a reduction in sovereign spread. This result is consistent within the four specifications of the empirical model, in terms of statistical significance, parameter signs and parameter values. Turning to domestic performance, domestic growth, measured by the monthly index of economic activity (Imacec), the estimated coefficients are consistently negative for the four estimated models, but the estimated coefficient is only significant in equation 4. Higher domestic growth should then results in lower sovereign spreads, since the likelihood of timely repayment increases.

Turning to international interest rates, the federal funds rate shows a positive and statistically significant coefficient in all equations, as predicted by the theoretical considerations described above. A rise in the fed funds rate raises the spread both because of the risk of default and because that risk rises as the risk-free rate goes up. Therefore, changes in U.S. interest rates, or likewise in global liquidity conditions, would be expected to influence positively the Chilean sovereign spreads.

Table 12: Determinants of Chilean Sovereign Spread

VARIABLES	EQUATION 1	EQUATION 2	EQUATION 3	EQUATION 4
Short Term Debt	1.42			
	(1.62)			
Short Term Debt/Reserves		76.47		
		(2.04)		
Total External Debt			0.95	
			(0.16)	
Total External Debt/Reserves				63.64
				(2.07)
Exports	– 1.50	– 1.31	– 1.32	– 1.45
	(– 2.35)	(– 2.13)	(– 2.02)	(– 2.36)
IMACEC	– 13.02	– 12.44	– 13.28	– 17.56
	(– 1.69)	(– 1.65)	(– 1.42)	(– 2.29)
Fed Funds Rate	13.45	16.67	15.12	17.47
	(5.63)	(7.61)	(6.59)	(7.48)
Dummy 2012	77.63	76.07	77.78	73.85
	(3.86)	(3.86)	(3.65)	(3.75)
Dummy 2007	– 47.93	– 46.91	– 45.31	– 50.51
	(– 2.35)	(– 2.36)	(– 2.12)	(– 2.53)
R^2	0.73	0.74	0.71	0.74
Adjusted R^2	0.68	0.69	0.65	0.69
N. Obs.	38	38	38	38

Note: t-stats in parenthesis.

A final word on the 2012 and 2007 dummies, which take into account the most recent issues undertaken by the Chilean government. The coefficient on the 2012 dummy shows a direct positive and statistically significant impact on the sovereign spread. Such a jump in the spread could in part be explained by portfolio balance considerations, but it should also be noted that the 2012 bond was also issued right after the 9/11 events, and so in part it reflects the prevailing market turbulence prevailing in those days, and not purely the portfolio balance effects on the new issuance. On the other hand, the 2007 dummy presents an inverse effect on sovereign spread of the 2009 bond. In fact, the issue spread was influenced by the very favorable conditions regarding the prevailing market prices and a high yield oriented demand could have caused the change in sign of the 2007 dummy, thus showing a decreasing one time effect over the 2009 bond spread.

V.- CONCLUSIONS

The Chilean sovereign bond has had a rather short but quite interesting history, marked by world and regional turbulence since the first bond issuance in April 1999. Not only the Asian crisis affected global markets, but also Brazil and Argentina. So a closer look at either the daily behavior of the spread series through time series models, or at the medium to long term determinants of the sovereign spread series is granted, since the sovereign spread is nowadays the most clear indicator of the cost of external financing for the Chilean economy as a whole.

Time series analysis of the spread show that shocks seem to be very persistent. However, the series seems to be stationary around a certain level, and it exhibits excess kurtosis, which can be better captured by GARCH type models. In particular, asymmetric GARCH models, such as the GJR-GARCH have led us to the result that downward movements in the spread are followed by significantly lower volatilities than upward movements of the same magnitude.

In terms of fundamentals analysis, the estimated models provide a very reduced set of variables that might explain the medium to long term level of the spread. These variables include liquidity indicators (short term debt/reserves ratio), economic performance variables (external and domestic), and U.S. interest rates. A higher short term debt to reserves ratio, i.e. lower international liquidity, should increase the sovereign spread. Improvements in either domestic or external performance should also reduce the spread of the sovereign bond. And finally, an increase in the Fed Funds rate, i.e. a tightening of the U.S. monetary stance, should increase the sovereign spread, as global liquidity is reduced.

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APPENDIX 1: VARIABLE DEFINITIONS

VARIABLE	DEFINITION
Short-term Debt	Monthly change, 3 month moving average
Short-term Debt/Reserves	Ratio of short-term external debt to international reserves
Total External Debt	Monthly change, 3 month moving average
Total External Debt/Reserves	Ratio of total external debt to international reserves
Exports	Monthly change, 3 month moving average
IMACEC	Monthly change, 3 month moving average
Fed Funds Rate	Federal Funds Rate, Monthly Average
Dummy 2012	Takes value 1 for October 2001
Dummy 2007	Takes value 1 for March 2002

Sources: Central Bank of Chile, Bloomberg.

Figure 1: Chilean Sovereign Spread

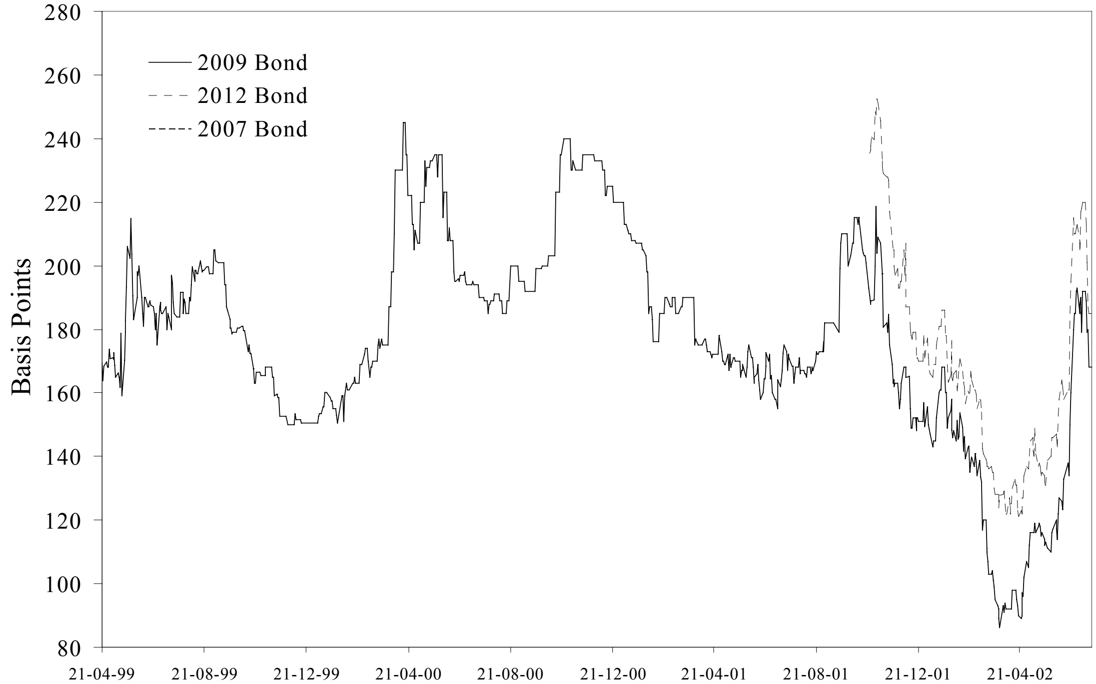

Figure 2: Investment Grade Economies Sovereign Spreads

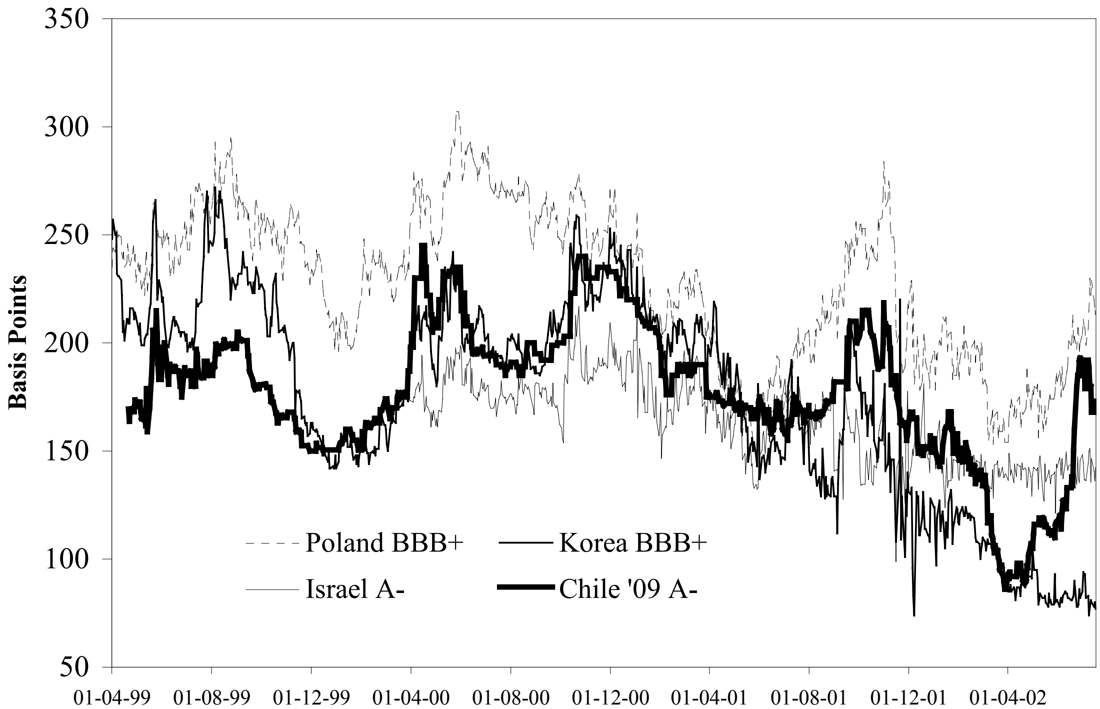

Figure 3: News Impact Curve GARCH(1,1)

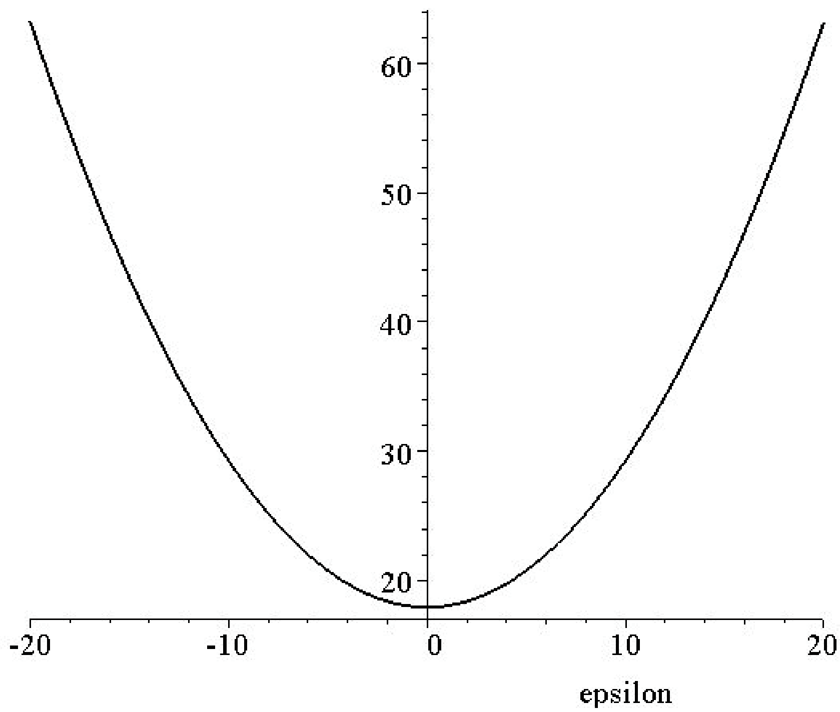

Figure 5: GARCH(1,1) Model, Conditional Variance

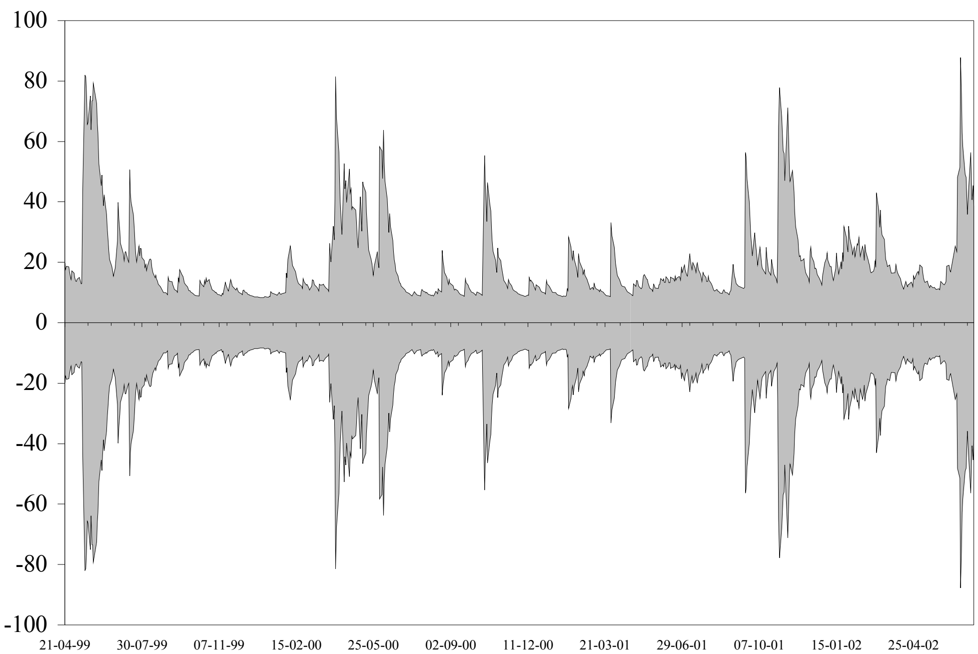

Figure 5: News Impact Curve GJR-GARCH Model

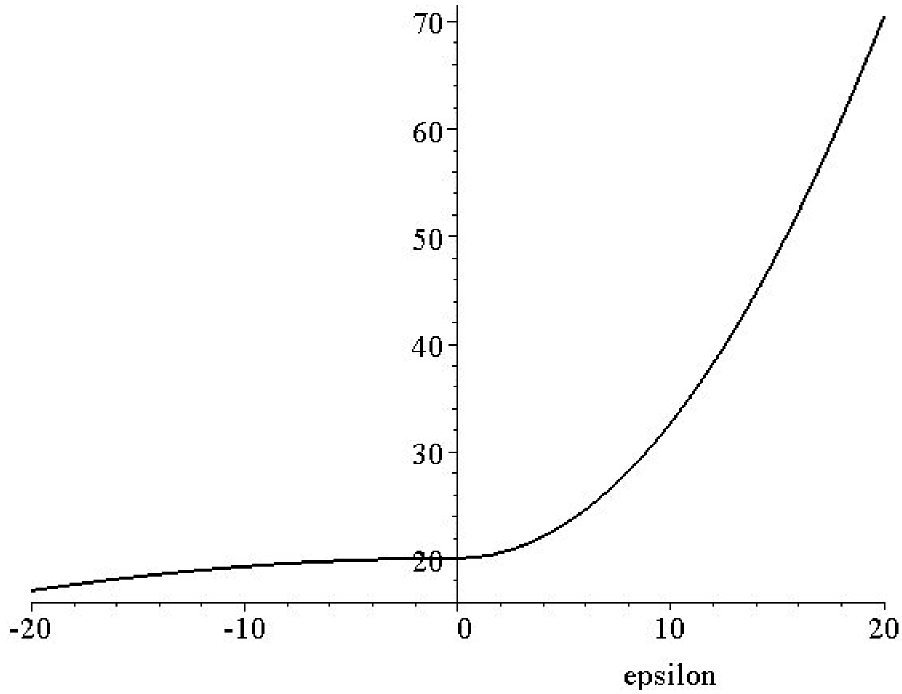

Figure 6: GJR-GARCH(1,1) Model, Conditional Variance