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THE ADJUSTMENT TO TARGET LEVERAGE OF SPANISH PUBLIC FIRMS: MACROECONOMIC CONDITIONS AND DISTANCE FROM TARGET*

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Our evidence suggests that Spanish public firms adjust slowly toward their capital structure target, with the typical firm closing approximately one-fifth of the gap between its current and target debt ratios each year. This finding is in contrast with previous evidence; however, we employ econometric techniques specially designed for highly persistent dependent variables, like market debt ratios. Moreover, our evidence does not seem to indicate that macroeconomic conditions, at least under the conditions experienced by the Spanish economy during our sample period, affect the speed of adjustment. If anything, our results are consistent with faster adjustments during economic states in which the distance between the current and target leverage is the greatest.

Key words: market debt ratio, dynamic trade-off, target leverage, speed of adjustment, macroeconomics, distance from target.

JEL classification: G32, C33, E30.

Three main theories are recognized in explaining capital structure: market timing, the pecking order approach and the trade-off theory. As expected, the empirical evidence closely follows the development of theoretical models[1]. The first collection of papers studies the determinants of capital structure under the implicit assumption that observed and target market debt ra-

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(1) See, among many others, Shyam-Sunder and Myers (1999), Baker and Wurgler (2002), Welch (2004), Leary and Roberts (2005), Kayhan and Titman (2007), Lemmon, Roberts, and Zender (2008), Byoun (2008), Frank and Goyal (2009), and Berk, Stanton, and Zechner (2010).

tios are basically the same. The second set of papers deals with factors that may cause firms to be over- or underleveraged with respect to their target market debt ratios. The final group of papers closely examines capital structure changes. In particular, they look for factors that may prevent firms from maintaining constant (target) market debt ratios. In this sense, recent literature focuses on analyzing how and when firms dynamically adjust their capital structure. This is the framework in which we conduct our research. Specifically, the main objective of the paper is to study how Spanish public firms adjust toward their capital structure target. This might be relevant because previous empirical evidence from Spanish firms shows a relatively high speed of adjustment toward target debt ratios, at least with respect to the typical U.S. firm. De Miguel and Pindado (2001), using data from 1990 to 1997, report a surprisingly high speed of adjustment of approximately 0.79, which suggests that nearly 80 percent of the debt ratio is adjusted each year toward target leverage. González and González (2008) report that, between 1995 and 2004, the typical Spanish firm closes approximately 54 percent of the gap between its current and target debt ratios each year. These results suggest that presenting additional and more recent evidence of the speed of adjustment of Spanish firms may be appropriate.

Interestingly, the U.S. empirical evidence about how quickly firms move toward their target market debt ratios is controversial because the results seem to be sensitive to the econometric specifications employed as well as to external financing constraints. Flannery and Rangan (2006), in a very influential paper, report a relatively fast speed of adjustment of 35.5 percent per year, which suggests that approximately one-third of the gap between current and target debt ratios is closed each year. However, Faulkender, Flannery, Hankins, and Smith (2008) argue that the speed of adjustment depends heavily on firms having access to external capital markets. The speed of adjustment is 31.3 percent for firms with access to external markets, but only 17.1 percent for firms with more restrictive access to external capital markets[2]. Similarly, Byoun (2008) reports that the adjustment speed is around 33 percent when firms have above-target debt with a financial surplus but about 20 percent when firms have below-target debt with a financial deficit. Huang and Ritter (2009) show how sensitive the results are to the econometric techniques employed in estimating the speed of adjustment. Finally, Lemmon *et al*. (2008) argue that, although observed leverage is consistent with the existence of a target debt ratio, since firms seem to trade off the costs and benefits of leverage, their lagged (up to 15 years) debt ratios are highly significant determinants of their current capital structure[3].

The first contribution of this paper is to provide more recent evidence of the speed of adjustment of Spanish firms and, more importantly, to provide new results under an econometric technique especially appropriate for a dynamic panel when

(2) All these figures are based on market leverage. The evidence using book leverage instead shows slightly lower speeds of adjustment.

(3) Also, recent papers by Chang and Dasgupta (2009) and Iliev and Welch (2010) argue that the empirical evidence in favor of target behavior and dynamic rebalancing of leverage may be caused by mechanical mean reversion.

the dependent variable is highly persistent. Specifically, our results are based on the long difference estimator of Hahn, Hausman, and Kuersteiner (2007). It turns out that, like Huang and Ritter (2009), we show that this estimator is much less biased than the traditional generalized method of moments (GMM) estimators employed in previous papers with Spanish and U.S. data. In particular, our evidence shows that the typical Spanish firm moves much more slowly toward its target leverage than previously reported estimates. The speed of adjustment is found to be 17.5 percent, which indicates that approximately only one-fifth of the gap is closed each year. This seems to contrast not only with the available Spanish evidence discussed above, but also with the higher speed traditionally accepted for U.S. firms where approximately one-third of the gap is supposed to be closed each year[4]. However, it is closer to the most recent U.S. evidence obtained using a long difference estimator[5].

The second and more specific objective of the paper is to analyze how macroeconomic conditions influence the adjustment speed of capital structure of Spanish public firms toward their target leverage levels. Several theoretical papers discuss the influence of macroeconomic conditions on financing patterns. Korajczyk and Levy (2003) argue that the response of firms to cyclical fluctuations depends upon the stringency of financing constraints. Levy and Hennessy (2007) find that relatively less constrained firms exhibit countercyclical variations in market debt ratios. More specifically, the authors find countercyclical variations in outstanding debt and procyclical variations in outstanding equity. They propose a general equilibrium model with agency costs in which firms substitute debt for equity to maintain managerial equity shares during recessions to avoid agency conflicts. The opposite occurs during expansions, because increases in managerial wealth facilitate the substitution of equity for debt. The authors also show that the relatively constrained firms should be characterized by a flat market debt ratio over the economic cycles. Finally, Hackbarth, Miao, and Morellec (2006) formally show that there is a benefit for firms to accommodate their capital structure decisions to the macroeconomic cycle, as long as operating cash flows depend on current economic conditions. Their dynamic model shows that the countercyclical nature of the aggregate market debt ratio is basically due to the relatively higher increase in the present value of operating cash flows during booms. Moreover, this is the only theoretical paper with clear-cut implications about the speed of adjustment toward target leverage. Given that firms' restructuring thresholds are lower in expansions than in recessions, the authors show that firms should adjust their capital structure more often during expansions than in contractions. Hence, the speed of adjustment should be higher in expansions than in recessions.

The empirical evidence about the effects of macroeconomic conditions on the speed of adjustment to target capital structure is rather scarce. Drobetz and Wanzenried (2006) show that the speed of adjustment of Swiss firms is higher when the term spread is higher and economic prospects are good. Drobetz, Pensa, and

(4) See Flannery and Rangan (2006) and Antoniou, Guney, and Paudyal (2008), who report coefficients of 35.5 and 32.2 percent, respectively.

(5) Huang and Ritter (2009) estimate a speed of adjustment of 23.2 percent per year for market leverage.

Wanzenried (2007) extend their previous analysis to a representative sample of European firms with data from 1983 to 2002. These authors also conclude that the adjustment is faster under favorable macroeconomic conditions. In particular, European firms tend to move faster toward their target leverage when interest rates are low and global financial distress is negligible. Surprisingly, to the best of our knowledge, the only evidence analyzing the effects of economic conditions on the speed of adjustment using U.S. data is reported by Cook and Tang (2010). By defining good and bad states of macroeconomic conditions based on the term spread, gross domestic product (*GDP*) growth, default spread and dividend yield, and using data from 1977 to 2006, the authors conclude that U.S. firms adjust their debt ratios toward target leverage faster in good macroeconomic states than in bad states.

Our research employs an alternative specification of the partial adjustment model of Flannery and Rangan (2006) that allows us to explicitly recognize potential macroeconomic effects on the speed of adjustment. Indeed, the second contribution of this paper is to incorporate macroeconomic conditions into the coefficient of the speed of adjustment. In other words, the adjustment speed is endogenized in the partial adjustment model. We also follow Cook and Tang (2010) in defining good and bad macroeconomic states using *GDP* growth, term spread (*TERM*), and the price-earnings ratio (*PER*). In particular, a high contemporaneous *GDP* growth and a high lagged term spread and price-earnings ratio indicate good macroeconomic states.

Our results first show a countercyclical market debt ratio. Regarding the speed of adjustment, our empirical evidence, always based on the long difference estimator, seems to be sensitive to the methodology employed to analyze the impact of macroeconomic conditions. Under the endogenized speed of adjustment, we do not find any significant results. Interestingly, however, when we define either good or bad states as an additional explanatory variable of the model, our results show that Spanish firms adjust their capital structure back to target leverage faster in bad states than in good states. As discussed above, this result contradicts the previous international evidence. However, we must first point out that our sample period does not really include a recession cycle in a formal sense. Indeed, the average *GDP* growth during our bad state definition is as high as 2.8 percent, while the minimum value is 2.4 percent. More importantly, this evidence is consistent with a larger distance between target leverage and market debt ratios in relatively bad economic states. Rather than a financing response to macroeconomic conditions, we are probably observing a faster adjustment during economic states in which the distance between the current and target leverage is larger. Indeed, we report evidence consistent with this interpretation.

The rest of the paper is organized as follows. Section 1 describes the partial adjustment model with macroeconomic conditions and discusses firms' characteristic target determinants and potential macroeconomic determinants of leverage. Section 2 presents the data and summary statistics. Section 3 presents estimates of the speed of adjustment to target leverage obtained under alternative econometric procedures and discusses the evidence contingent on macroeconomic conditions and distance to target leverage. Finally, Section 4 presents our conclusions.

1. The partial adjustment model of firm leverage with macroeconomic conditions

1.1. Model specifications with macroeconomic variables

The market debt ratio is defined as

$$MDR_{jt} = \frac{D_{jt}}{D_{jt} + n_{jt}P_{jt}}, \qquad [1]$$

where D_{jt} is the book value of firm *j's* interest-bearing debt, n_{jt} equals the number of common shares outstanding and P_{jt} denotes the price per share[6]. The basic partial adjustment model of firm leverage proposed by Flannery and Rangan (2006) is given by

$$MDR_{jt} - MDR_{jt-1} = \lambda\left(MDR^{*}_{jt} - MDR_{jt-1}\right) + u_{jt}, \qquad [2]$$

with λ being the adjustment speed toward target leverage and the standard partial adjustment model defining the desired debt ratio as

$$MDR^{*}_{jt} = \beta' X_{jt-1}, \qquad [3]$$

where MDR^{*}_{jt} is firm *j's* target market debt ratio at t, X_{jt-1} is a K-vector of firm characteristics and β is a K-vector of coefficients such that the trade-off hypothesis implies that $\beta \neq 0$[7]. Substituting [3] into [2], we obtain the basic specification that allows an incomplete adjustment of a firm's initial capital structure toward its target within each time period:

$$MDR_{jt} = \left(\lambda\beta\right)' X_{jt-1} + \left(1-\lambda\right)MDR_{jt-1} + u_{jt}, \qquad [4]$$

To analyze the potential effects of macroeconomic conditions on the adjustment speed, we first allow λ to be time-varying by assuming that the adjustment speed toward target leverage is a linear function of macroeconomic conditions. These are defined by a vector of (time-varying) macroeconomic variables that de-

(6) The choice between scaling debt by the market value of equity or the book value of equity is not obvious. The research briefly discussed in the introduction tends to employ both possibilities when presenting the results. As mentioned above, these results are not sufficiently different to modify the economic conclusions. For this reason, and given that scaling by market values is theoretically more sound, in this paper we only report results based on the market value of equity.

(7) Although it has become standard to use the cross-sectional regression market debt ratio, given by expression [3], to proxy for the target or optimal debt ratio, it may be useful to study the robustness of the results to alternative definitions of the target debt ratio. In particular, the evidence reported by D'Mello and Farhat (2008) suggest that we should use the moving average market debt ratio rather than expression [3]. Another possibility would be to employ a target debt ratio based on a moving average leverage of actual leverage portfolios in event time, as in Figures 1 and 2 of Lemmon *et al*. (2008). The target debt ratio of each firm would be the resulting long-term leverage of the portfolio to which the firm belongs in each event time.

scribe the actual (observed) cycle of the economy. In particular, we assume the linear function

$$\lambda_{t-1} = \lambda_0 + \lambda_1 M_{1t-1} + \lambda_2 M_{2t-1} + \ldots + \lambda_L M_{Lt-1}\,, \qquad [5]$$

where λ_0 is a scalar and λ_1, ..., λ_L are the sensitivities of the speed of adjustment to L current macroeconomic variables denoted by M_1, ..., M_L[8]. If macroeconomic conditions do not affect the speed of adjustment toward the desired target market debt ratio, λ_1, ..., λ_L should all be equal to zero and λ_{t-1} would be constant and equal to λ_0. It must be noted that the model assumes that all firms are equally affected by macroeconomic conditions. This may easily be generalized to allow for specific dependence on macroeconomic conditions.

By substituting [5] into [4] and rearranging, we obtain the following specification:

$$MDR_{jt} = \lambda_0 \sum_{k=1}^{K} \beta_k X_{kjt-1} + \left(1-\lambda_0\right) MDR_{jt-1} - \left(\sum_{l=1}^{L} \lambda_l Y_{ljt-1}\right) + \sum_{q=1}^{LxK} \eta_q Z_{qjt-1} + u_{jt}\,, \qquad [6]$$

where $Y_{ljt-1} \equiv M_{lt-1} x MDR_{jt-1}$ for $l = 1, ..., L$; $\eta_q \equiv \lambda_l x \beta_k$ for $l = 1, ..., L, k = 1, ..., K$ and $q = 1, ..., LxK$; and $Z_{qjt-1} \equiv M_{lt-1} x X_{kjt-1}$ for $q = 1, ..., LxK$. This specification can be written more compactly as

$$MDR_{jt} = \left(\lambda_0 \beta\right)' X_{jt-1} + \left(1-\lambda_0\right) MDR_{jt-1} - \lambda' Y_{jt-1} + \eta' Z_{jt-1} + u_{jt}\,, \qquad [7]$$

where λ is the L-vector of sensitivities of the adjustment speed to macroeconomic conditions, Y_{jt-1} is the L-vector of the product of L macroeconomic variables and the market debt ratio of each firm at time $t - 1$, η is the LxK-vector of the product of $\lambda's$ and $\beta's$, and Z_{jt-1} is the LxK-vector of the product of L macroeconomic variables and K firms' j characteristics.

We are especially concerned with the parameters describing the speed of adjustment. We should pay special attention to λ_0 and the sensitivities λ to alternative macroeconomic variables used in the empirical application. If these sensitivities are all equal to zero, the key parameter would be λ_0 and the interpretation would be similar to that of Flannery and Rangan (2006). Alternatively, if we find evidence of significant sensitivities, we may learn which macroeconomic variable determines the speed of adjustment while the total speed of adjustment would be given by λ_{t-1}. The interpretation of the L-vector of λ coefficients associated with the interaction terms, $Y_{ljt-1} \equiv M_{lt-1} x MDR_{jt-1}$, is especially relevant. Let us consider a particular macroeconomic variable that suggests future good economic prospects. If macroeconomic conditions influence the speed of adjustment, we expect a significant λ_l coefficient. Additionally, if, as suggested by Hackbarth *et al.* (2006), we expect faster adjustments in expansions than in recessions, the interaction term between the lagged market debt ratio and the macroeconomic variable should be positive.

(8) The subindex t - 1 in the lambda coefficient is due to the fact that the macroeconomic variables are only known at time t - 1 but not at time t.

As a second way of incorporating macroeconomic effects into the partial adjustment model, we follow Cook and Tang (2010). We first define the target leverage for each firm as[9]

$$MDR_{jt}^{*} = \beta' X_{jt-1} + \gamma M_{t-1} + \eta M_{t-1} x MDR_{jt-1} \,, \qquad [8]$$

where we allow for an interaction term to recognize that the lagged market debt ratio is a significant determinant of the firm's capital structure[10]. Substituting [8] into [2], we obtain the specification tested in the empirical section of the paper,

$$MDR_{jt} = (\lambda\beta)' X_{jt-1} + (1-\lambda) MDR_{jt-1} + (\lambda\gamma) M_{t-1} + (\lambda\eta) M_{t-1} x MDR_{jt-1} + u_{jt} \,. \qquad [9]$$

It is important to recall that λ is positive. This implies that the coefficient associated with the macroeconomic variable should be negative in order to have a countercyclical market debt ratio as the theoretical models suggest. Moreover, the coefficient of the interaction term provides information about the speed of adjustment, depending upon the level of the macroeconomic variable.

1.2. Firm characteristics and target leverage

To characterize a target debt ratio, we use a set of firm indicators that have appeared regularly in previous empirical papers[11]. In our case, the following variables are used as control variables in equations [4], [7] and [9]:

- Profitability measured as earnings before interest and taxes over total assets (*ROA*). It is generally accepted that firms that have more profits tend to have lower leverage given that high retained earnings reduce the need of external financing with debt. It could also be the case that firms limit the issuance of debt to protect their competitive advantage while producing these high operating profits.
- Market-to-book ratio (*MTB*). Firms that have a high market-to-book ratio tend to have lower leverage since it is generally a sign of future growth opportunities.
- Effective tax rate, defined as taxes paid as a proportion of earnings before taxes (*ETR*). Firms with higher effective tax rates are (theoretically) expected to have more leverage to take advantage of the tax deductibility of interest payments. However, the empirical evidence has proved to be unclear.

(9) This version of the model assumes that there is only one macroeconomic variable that summarizes all the relevant information about the state of the economy.

(10) See Lemmon *et al*. (2008). Cook and Tang (2010) initially propose a model in which the target leverage depends on the lagged macroeconomic variable and the firm's financial characteristics. However, their empirical application includes an interaction term.

(11) The financial characteristics chosen are very common in the previous related literature. They basically coincide with the financial variables employed by Flannery and Rangan (2006) except that we do not explicitly use research and development (*R&D*) expenses as a proportion of total assets but, rather, employ non-tangible assets as a proportion of total assets. Additionally, we also take into account the effective tax rate and interest expense coverage. The effective tax rate is used by Huang and Ritter (2009) and the interest expense coverage is included as a measure of financial distress. Our chosen financial variables are also similar to those used by De Miguel and Pindado (2001), although our variables are more disaggregated.

- Non-debt tax shields calculated as depreciation expenses as a proportion of total assets (*NDTS*). Firms with more depreciation expenses may be under less pressure to increase their debt to take advantage of the deductibility of interest payments.
- The natural logarithm of total assets (*SIZE*). Large firms tend to have higher leverage because they have fewer restrictions in their access to financial markets, lower cash flow volatility and less financial distress.
- Fixed assets as a proportion of total assets, or tangibility (*TANG*). Firms functioning with greater tangible assets, which are potentially collateralized, tend to have higher debt capacity.
- Non-tangible fixed assets as a proportion of total assets (*NTANG*). Firms with more intangible assets, especially if associated with R&D expenses, tend to have lower debt to protect themselves from higher bankruptcy costs.
- Interests paid as a proportion of earnings before interest and taxes, or interest expense coverage (*IC*). This variable may capture financial distress, although it may just be mechanically positively related to higher amounts of debt.
- Median market debt-to-equity ratio of the company's industry (*IDR*). Firms in industries in which the median company has high debt tend to have higher leverage.

1.3. Macroeconomic conditions, speed of adjustment and target leverage

To measure the macroeconomic conditions of the Spanish economy, we employ three variables, including real and financial aggregate variables. Specifically we use the lagged price-earnings ratio of the Spanish stock exchange (*PER*), the real growth rate of the gross domestic product (*GDP*) and the lagged term spread, defined as the difference between the ten-year and one-year government bonds (*TERM*). We find *PER* to be a predictor of both economic and stock market cycles[12], while changes in the term spread are a well-known predictor of future business cycles[13]. A flattening of the slope of the term structure is known to anticipate a recession, while an increase in the slope of the yield curve anticipates an improvement in the business cycle. The growth in *GDP* measures the current state of the economy. Following Hackbarth *et al*. (2006), we expect faster adjustments in good macroeconomic conditions. This suggests that a higher lagged *TERM*, a higher lagged *PER* and a higher contemporaneous real *GDP* growth should be associated with a faster speed of adjustment toward target leverage.

2. Data

Our sample consists of Spanish traded non-financial firms (including utilities). Panel A of Table 1 reports the number of companies in the sample and the number of annual observations per firm. The final sample contains 101 firms with incomplete information for the 13-year period between 1995 and 2007. This im-

(12) See, for example, references in Bollerslev, Tauchen, and Zhou (2009).
(13) See, among many others, Cochrane and Piazzesi (2005).

plies that we have unbalanced panel data with a total of 1,112 observations. Panel B of Table 1 contains the firm distribution by industry and the specific weight of each firm sector over the total sample. Three sectors account for almost 75 percent of the total number of firms in the sample: manufacturing, mining and quarrying; construction; and wholesale and retail trade, transportation and accommodation.

Table 1: Sample Characteristics

Panel A: Annual Availability of Firms in the Sample

Annual Observations per Firm	Number of Firms	Total Number of Observations
13	37	481
12	20	240
11	9	99
10	11	110
9	7	63
8	6	48
7	5	35
6	6	36
Total	101	1,112

Panel B: Distribution of Firms by Economic Sector

	Economic Sectors	Firms	Percentage
Sector 1	Agriculture, forestry and fishing	2	1.98
Sector 2	Manufacturing, mining and quarrying	45	44.55
Sector 3	Electricity, gas and water	8	7.92
Sector 4	Construction	16	15.84
Sector 5	Wholesale and retail trade, transportation and accommodation	14	13.86
Sector 6	Information and communication	6	5.94
Sector 7	Real estate activities	3	2.97
Sector 8	Professional, scientific and support service activities	6	5.94
Sector 9	Other services	1	0.99
Total	101	100	

Source: Own elaboration.

Given the usual requirements of dynamic panel data models, we construct a sample of firms that includes all non-financial companies listed on the Spanish stock exchange with at least six consecutive years of observations. The account-ing information of these firms is obtained from the Sistema de Análisis de Balances Ibéricos (SABI), a database managed by Bureau Van Dyck and Grupo Informa, S.A., while financial market information was provided by the quotation bulletins of the Spanish stock exchange.

Table 2 reports summary statistics for alternative measures of aggregate leverage and the total value of equity for our sample of firms. The market debt ratio is the long-term book value of debt of all firms in our sample divided by the sum of the aggregate market equity and the long-term book value of debt of all firms. The total market debt ratio is the total book value of debt of all firms divided by the sum of the aggregate market equity and the total book value of debt of all firms in our sample. Equity is the total market capitalization of all available firms. The last column of Table 2 shows the time structure of the available sample. Our sample ranges from 38 firms in 1995 to 101 companies in 2002 and 2003. The number of firms in the sample remains relatively stable after 2000.

Figure 1 displays the evolution of the (long-term) market debt ratio from 1995 to 2007. The sample is made up of two basic periods: an increasing pattern from 1997 to 2002 and a decreasing market debt ratio from 2003 to 2006. The latter is clearly influenced by the extremely good performance of the Spanish equity market during those four years. The highest ratio is reached in 1996, while there is slight change in tendency between 2006 and 2007. Figure 1 also shows the evolution of the three macroeconomic variables: the contemporaneous real *GDP* growth, the lagged *PER* and the lagged *TERM*. The three business cycle variables behave similarly over time, with positive correlation coefficients between them and an expected negative correlation between each of them and the market debt ratio. This confirms the countercyclical nature of the leverage ratio for Spanish firms.

Table 3 shows the main descriptive statistics of the dependent variable, the financial characteristics defining the target debt ratio and the three macroeconomic indicators. The financial characteristics are winsorized at the first and 99th percentiles to avoid the influence of very extreme observations. It is important to note that the average real *GDP* growth during our sample period is 3.65 percent, while the maximum and minimum growth rates are 5.3 percent and 2.4 percent, respectively. These statistics cast doubt on the ability of our sample period to capture sufficient fluctuations of the economy to classify the available years as either booms or recessions. This clearly introduces an interpretation problem about the effects of macroeconomic variables on the speed of adjustment of Spanish public firms.

Table 4 reports the correlation matrix for our financial and macroeconomic variables. According to most of the previous evidence, market debt ratios are negatively correlated with return on assets and the market-to-book ratio and positively correlated with firm size. We also find a negative correlation between market-to-book ratios and non-debt tax shields and between tangible assets and both return on assets and non-debt tax shields. Finally, the correlation between nontangible assets and non-debt tax shields is positive. The only potentially problematic correlation for the empirical tests reported below is the large negative correlation coefficient between size and the market-to-book ratio.

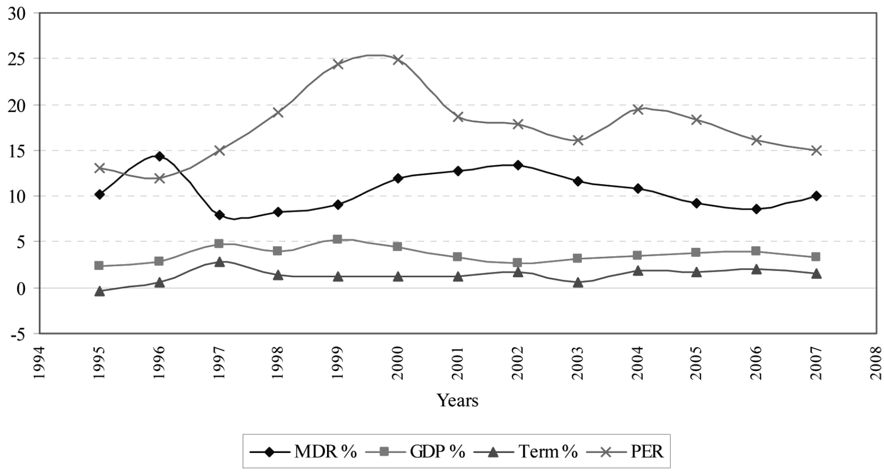
Table 2: Time evolution of the aggregate market debt ratio

Year	MDR Mean	MDR Median	TMDR Mean	TMDR Median	Equity Mean (millions €)	Equity Median (millions €)	Number Available Firms per Year
1995	0.1013	0.0023	0.2562	0.0991	1330	167	38
1996	0.1440	0.0344	0.2931	0.1479	1370	101	59
1997	0.0800	0.0151	0.2055	0.1044	1530	175	68
1998	0.0832	0.0172	0.2037	0.1320	1890	191	79
1999	0.0910	0.0231	0.2205	0.1746	2480	164	86
2000	0.1187	0.0261	0.2531	0.1574	2200	148	90
2001	0.1267	0.0213	0.2495	0.1952	2200	176	95
2002	0.1339	0.0363	0.2631	0.1919	1650	168	101
2003	0.1168	0.0208	0.2303	0.1569	2080	225	101
2004	0.1088	0.0115	0.2154	0.1149	2600	264	100
2005	0.0920	0.0075	0.1901	0.1169	3140	445	99
2006	0.0866	0.0063	0.1698	0.0925	4230	751	99
2007	0.1006	0.0050	0.1965	0.1159	4810	685	97

The market debt ratio (MDR) is the long-term book debt divided by the sum of the long-term book debt and market equity. The total market debt ratio (TMDR) is the total book debt divided by the sum of the total book debt and market equity. Equity is the total market capitalization in millions of euros.

Source: Own elaboration.

Figure 1: AGGREGATE MARKET DEBT RATIO AND MACROECONOMIC VARIABLES 1995-2007



Source: Own elaboration.

Table 3: SUMMARY STATISTICS OF DEPENDENT, EXPLANATORY, AND MACROECONOMIC VARIABLES

Variable	Mean	Median	St. Dev.	Minimum	Maximum
MDR	0.1068	0.0150	0.1717	0	0.9229
ROA	0.0524	0.0432	0.0599	-0.2058	0.4060
MTB	2.8547	1.1186	3.6164	-3.5923	10.2654
ETR	0.1848	0.1695	5.4862	-21.2503	187.5909
NDTS	0.0228	0.0163	0.0221	0	0.1814
SIZE	17.2132	17.7222	2.1779	11.3319	21.3704
TANG	0.5923	0.6204	0.2440	0.0507	0.9919
NTANG	0.0290	0.0063	0.0604	0	0.7235
IC	0.5622	0.1787	5.3246	-19.1171	151.2183
IDR	0.0435	0.0599	0.0864	0.0003	0.6120
PER	17.6908	17.8100	3.6912	11.9800	24.9300
GDP	0.0365	0.0340	0.0080	0.0240	0.0530
TERM	0.0135	0.0133	0.0074	-0.0033	0.0276

MDR is the market debt ratio, *ROA* the return on assets, *MTB* the market-to-book ratio, *ETR* the effective tax rate, *NDTS* the non-debt tax shields, *SIZE* firm size, *TANG* tangibility, *NTANG* non-tangibility, *IC* interest expense coverage, *IDR* the industry debt ratio, *PER* the aggregate price-earnings ratio, *GDP* the GDP growth, and *TERM* term spread.

Source: Own elaboration.

Table 4: Correlation matrix

	MDR	ROA	MTB	ETR	NDTS	SIZE	TANG	NTANG	IC	IDR	PER	GDP	TERM
MDR	1.000												
ROA	-.284	1.000											
MTB	-.530	.122	1.000										
ETR	-.020	-.003	.041	1.000									
NDTS	.094	.099	-.214	-.014	1.000								
SIZE	.377	.084	-.758	-.035	.124	1.000							
TANG	-.007	-.119	.154	.006	-.160	-.079	1.000						
NTANG	-.023	.071	-.046	.007	.274	.056	-.223	1.000					
IC	.071	.003	-.082	.029	.043	.021	-.054	.027	1.000				
IDR	.372	-.143	-.329	-.015	.033	.085	.123	-.031	.004	1.000			
PER	-.003	.026	-.010	-.014	-.020	-.021	-.004	.022	.008	.001	1.000		
GDP	-.066	.001	.001	-.007	.019	-.033	.002	-.022	.022	.017	.626	1.000	
TERM	-.052	-.023	.026	-.045	-.047	.027	.011	.019	-.016	.012	.039	.338	1.000

MDR is the market debt ratio, *ROA* the return on assets, *MTB* the market-to-book ratio, *ETR* the effective tax rate, *NDTS* the non-debt tax shields, *SIZE* firm size, *TANG* tangibility, *NTANG* non-tangibility, *IC* interest expense coverage, *IDR* the industry debt ratio, *PER* the aggregate price-earnings ratio, *GDP* the GDP growth, and *TERM* term spread.

Source: Own elaboration.

3. Empirical Evidence

3.1. The Long Difference Estimator

It is well known that the estimation of dynamic panel data as in expressions [4], [7], and [9] is non-trivial because the combination of fixed effects and a lagged highly persistent dependent variable can severely bias coefficient estimates, particularly when the panel length is small[14]. The Appendix at the end of this paper provides a detailed explanation of these biases and discusses potential solutions. Unfortunately, applying the traditional "within" transformation or first difference estimator to remove the time-invariant fixed effects generates a correlation between the new lagged dependent variable and the new error term that leads to biased and inconsistent results. Since the instrumental variable solution is difficult to implement because reasonable instrumental variables are difficult to obtain, many papers employ the system GMM estimation of Arellano and Bover (1995) and Blundell and Bond (1998). This procedure instruments for the first difference of predetermined variables with lags of their own levels and differences. However, as recently shown by Hahn *et al*. (2007) and Huang and Ritter (2009), the system GMM use of the full set of moment conditions does not provide proper guidance in dynamic panel data models when the dependent variable is highly persistent, so the autoregressive parameter is close to one[15]. Moreover, in this case, the GMM system estimators are downward biased, so the estimate of the speed of adjustment is upward biased.

We now briefly discuss the long difference estimator procedure. Consider a dynamic panel data equation at the end of year *t* with fixed effects:

$$MDR_{jt} = (\lambda\beta)' X_{jt-1} + (1-\lambda) MDR_{jt-1} + \left(\mu_j + u_{jt}\right). \qquad [10]$$

The dynamic equation at the end of year *t - k* with fixed effects is given by

$$MDR_{jt-k} = (\lambda\beta)' X_{jt-k-1} + (1-\lambda) MDR_{jt-k-1} + \left(\mu_j + u_{jt-k}\right). \qquad [11]$$

Then, subtracting [11] from [10], we get the equation to be estimated under the long difference estimator:

$$MDR_{jt} - MDR_{jt-k} = (\lambda\beta)' \left(X_{jt-1} - X_{jt-k-1}\right) + (1-\lambda)\left(MDR_{jt-1} - MDR_{jt-k-1}\right) + \\ + \left(u_{jt} - u_{jt-k}\right) \qquad [12]$$

We first estimate equation [12] by the system GMM to obtain the initial parameter estimators, using MDR_{jt-k-1} and X_{jt-k-1} as valid instruments. Then we obtain the residuals of equation [11], $\left(MDR_{jt-1} - \hat{\lambda}\hat{\beta}X_{jt-2} - \left(1-\hat{\lambda}\right)MDR_{jt-2}\right), \ldots, \left(MDR_{jt-k} - \right.$

(14) This is usually the case with Spanish data.

(15) This is, of course, the case with market debt ratios.

$\hat{\lambda}\hat{\beta}X_{jt-k-1} - \left(1-\hat{\lambda}\right)MDR_{jt-k-1}\Big)$ and, finally, we employ MDR_{jt-k-1}, X_{jt-k-1}, and those residuals as new instruments. This is the first iteration. Both Hahn *et al*. (2007) and Huang and Ritter (2009) use three iterations to obtain the final estimators. Additionally, using data from 1972 to 2001, Huang and Ritter (2009) try four alternative lag values for k and show that the coefficient of the speed of adjustment varies from 22.3 percent when $k = 4$ to 17.6 for $k = 28$.

When recognizing the potential effects of macroeconomic variables, we follow the same procedure, using equations [7] and [9] with fixed effects instead of [10].

3.2. The speed of adjustment of Spanish public firms

This paper estimates the speed of adjustment using four alternative econometric procedures described in the Appendix and in Section 3.1[16]. We first employ the demeaned Fama-MacBeth (hereafter FM, 1973) and the difference GMM. Given the previous discussion and the analysis presented in the Appendix, we also report results for the system GMM of Arellano and Bover (1995) and Blundell and Bond (1998), and the long difference estimator of Hahn *et al*. (2007). All procedures apply to the dynamic panel data equations with and without macroeconomic conditions.

Table 5 presents the results of the estimation of the partial adjustment model given by expression [4]. There are four lagged financial explanatory variables with the correct theoretical sign that are estimated with a relatively low standard error. Profitability, the market-to-book ratio, the industry market debt ratio and interest coverage are all statistically significant, at least for the long difference and the system GMM estimators given in the first and second columns of Table 5. Firms with high profitability and market-to-book ratios have less leverage, while firms with higher interest coverage and those that belong to an industry with high leverage present higher market debt ratios. Tangible assets tend to be positively and significantly associated with leverage under traditional panel data methods, but this characteristic is estimated with a large standard error under the long difference estimator.

The specification tests under the long difference estimator and the GMM methodologies suggest that we cannot reject the model at conventional significance levels. As expected, given the discussion on the econometric issues, the coefficient of the speed of adjustment tends to vary quite significantly across estimation methodologies. The demeaned FM two-stage estimation presents a very high speed of adjustment. The GMM estimator procedures obtain a speed of adjustment of 24.9 percent and 30.8 percent for the system and difference GMM, respectively. These values are similar to, although slightly lower than, the speed of adjustment found for U.S. companies by Flannery and Rangan (2006) and Antoniou *et al*. (2008).

As in Huang and Ritter (2009), we find that the long difference estimator generates a speed of adjustment lower than that under the system GMM methodology. The long difference estimator is obtained with three iterations and for $k = 2$. Our database is much smaller than that available to Huang and Ritter (2009), which clearly limits our possibilities for checking our results for alternative lags. The

(16) All models are estimated with both time and sector dummies.

Table 5: ESTIMATION RESULTS OF THE PARTIAL ADJUSTMENT MODEL TOWARD TARGET CAPITAL STRUCTURES

Explanatory Variable	Long Difference	System GMM	Difference GMM	Demeaned FM
MDR	0.8246 (0.0339)	0.7513 (0.0574)	0.6920 (0.0603)	0.5268 (0.1590)
ROA	–0.2218 (0.0669)	–0.2027 (0.0609)	–0.1918 (0.0771)	–0.0403 (0.1116)
MTB	–0.0046 (0.0017)	–0.0060 (0.0022)	0.0123 (0.0084)	0.0090 (0.0080)
ETR	0.0001 (0.0001)	0.0001 (0.0001)	0.0001 (0.0004)	0.0049 (0.0050)
NDTS	0.1207 (0.1667)	0.2311 (0.1819)	0.3776 (0.4081)	0.4035 (0.5896)
SIZE	–0.0007 (0.0032)	–0.0005 (0.0035)	–0.0080 (0.0144)	0.0336 (0.0190)
TANG	0.0159 (0.0129)	0.0229 (0.0142)	0.0278 (0.0351)	0.0755 (0.0422)
NTANG	–0.0088 (0.0860)	–0.0328 (0.0817)	–0.1027 (0.1139)	0.3533 (0.1962)
IC	0.0027 (0.0011)	0.0024 (0.0011)	0.0024 (0.0008)	0.0054 (0.0017)
IDR	0.6854 (0.2634)	0.7985 (0.3371)	0.0233 (0.0000)	
Number of firms	101	101	101	101
Number of observations	853	853	794	955
Adjusted R^2				0.2500
1st-order autocorrelation	–4.61 (0.0000)	–4.09 (0.0000)	–4.64 (0.0000)	
2nd-order autocorrelation	–1.55 (0.2990)	–1.60 (0.3874)	–1.62 (0.2752)	
Hansen–Sargan test	32.36 (0.357)	148.64 (0.215)	147.67 (0.236)	

Panel regression coefficients estimated from equation [4] with standard errors in parentheses. All the models include both time and sector dummies. The first of the estimations is carried out by GMM using the long difference estimator of Hahn *et al*. (2007). The System GMM column includes Arellano and Bover's (1995) estimation procedure. The Difference GMM column provides Arellano and Bond's (1991) estimator, the robust version, taking the model in first differences. The demeaned FM column shows the estimation of the within transformation of the model. The first –and second-order autocorrelations test the null hypothesis of no first– or second-order autocorrelation, respectively, in the residuals. The Hansen-Sargan test is a test of the overidentifying restrictions under the null hypothesis of instrument validity.

Source: Own elaboration.

speed of adjustment for Spanish public firms under the long difference estimator is 17.5 percent. This implies that the typical Spanish firm closes approximately one-fifth of the gap between its current and target debt ratios each year. Interestingly, the estimate of the speed of adjustment is clearly lower than those reported by De Miguel and Pindado (2001) and González and González (2008)[17], which suggests that Spanish firms adjust slowly toward their target leverage. In fact, that estimate is even lower than the speed coefficient reported by Huang and Ritter (2009) for U.S. data, also using the long difference estimator. Under the system GMM procedure, typical firms close one-fourth of the gap each year. It would be interesting to investigate whether this relatively low speed of adjustment of Spanish firms can be explained by the particularly narrow market for corporate debt in Spain. It may just be the case that restrictions to external market debt financing impact the potentially desired changes of the capital structure of Spanish firms.

3.3. The Time-Varying Speed of Adjustment of Spanish Public Firms

Table 6 contains evidence about the influence of macroeconomic conditions on the speed of adjustment toward the target debt ratio. We now estimate equation [7] using the long difference estimator, again with three iterations and $k = 2$. We estimate this equation separately for each of our three macroeconomic variables described in subsection 1.3[18]. As before, temporal data limitations can reduce the possibility of finding significant results. The low frequency of data, available exclusively on a year-by-year basis, can also bias our results. None of the macroeconomic variables show any significant impact on the speed of adjustment. It seems that, contrary to the U.S. evidence, Spanish firms do not move faster toward target leverage during good economic periods. As already mentioned, this conclusion should be made with caution, because our sample period does not present significantly economic and contractions.

In any case, an interesting characteristic of the first procedure proposed to study the impact of macroeconomic conditions is that we are allowed to estimate a time-varying speed of adjustment as given by equation [5], once we have the corresponding λ coefficients[19]. Figure 2 shows the time-varying speeds of adjustment for each year in the sample period and the real *GDP* growth rate.

It seems clear that the speed of adjustment remains constant for all years when we use either lagged *PER* or lagged *TERM* to describe the business cycle. However, the behaviour of the time-varying speed of adjustment when we employ the real *GDP* growth seems to change over time and countercyclically with respect to *GDP*

(17) We must recognize that the sample period of De Miguel and Pindado (2001) is very different from ours. Moreover, González and González (2008) use a much more similar period, but they use a large database of 39 countries and do not report the number of Spanish firms employed when they report the speed of adjustment in their Table 5.

(18) From Table 4, we know that the correlation coefficient between *GDP* growth and lagged *PER* is as high as 0.63, while the correlation of *GDP* growth with lagged *TERM* is 0.34.

(19) It is also the case that this procedure, in which the speed of adjustment is endogenized, introduces a problem of nonlinearity in the coefficients. This is why we report the results from an alternative methodology below.

Table 6: ESTIMATION RESULTS OF THE PARTIAL ADJUSTMENT MODEL TOWARD TARGET CAPITAL STRUCTURES WITH MACROECONOMIC VARIABLES

Explanatory Variable	Long Difference Estimator		
MDR	0.8097 (0.2595)	0.6969 (0.2008)	0.8194 (0.0844)
ROA	–0.1949 (0.0616)	–0.2299 (0.0684)	–0.2075 (0.0641)
MTB	0.0056 (0.0052)	0.0043 (0.0091)	0.0058 (0.0026)
ETR	–0.0138 (0.0154)	–0.0064 (0.0097)	–0.0011 (0.0009)
NDTS	–3.2528 (0.8837)	–1.7666 (0.9113)	–0.0572 (0.3975)
SIZE	–0.0085 (0.0048)	–0.0022 (0.0122)	–0.0001 (0.0031)
TANG	0.1647 (0.0739)	0.0789 (0.0578)	0.0011 (0.0314)
NTANG	0.9457 (0.3692)	0.5156 (0.3802)	0.0795 (0.1117)
IC	0.0001 (0.0001)	0.0126 (0.0056)	0.0001 (0.0001)
IDR	1.4043 (0.4078)	1.3017 (0.3707)	0.3463 (0.2731)
PER x MDR	0.0004 (0.0126)		
GDP x MDR		3.3940 (4.9473)	
TERM x MDR			0.2635 (5.4883)
Number of firms	101	101	101
Number of observations	853	853	853
1st-order autocorrelation	–4.57 (0.0000)	–4.82 (0.0000)	–4.60 (0.0000)
2nd-order autocorrelation	–0.84 (0.4010)	–1.27 (0.2050)	–1.18 (0.2370)
Hansen–Sargan test	132.55 (0.234)	133.50 (0.246)	131.35 (0.241)

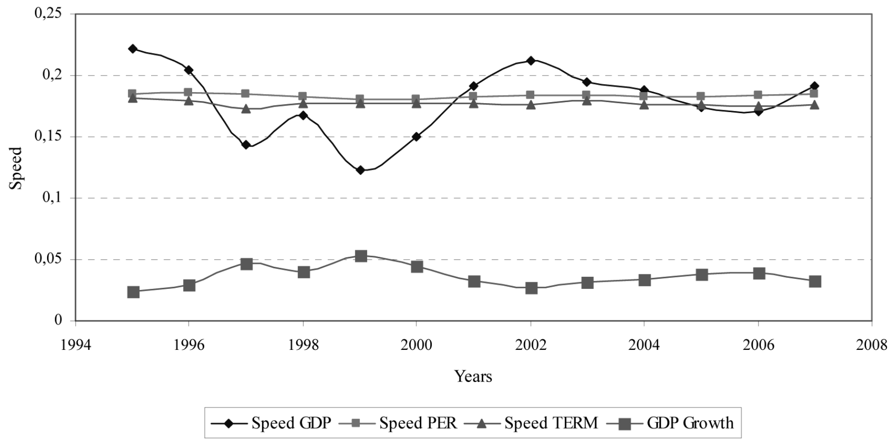
Panel regression coefficients estimated from equation [7], with standard errors in parentheses. All the models include sector dummies and all the estimations are carried out by GMM using the long difference estimator of Hahn *et al*. (2007). The first- and second-order autocorrelations test the null hypothesis of no first- or second-order autocorrelation, respectively, in the residuals. The Hansen-Sargan test is a test of the overidentifying restrictions under the null hypothesis of instrument validity.

Source: Own elaboration.

growth. Indeed, if we separate the 13 years of data into three groups, where the four years of higher *GDP* growth are classified as a good economic state while the four years of lower *GDP* growth are the contraction states, and then calculate the mean of the time-varying speed of adjustment for either good or bad states, we find that the average speed of adjustment is 0.15 during expansions and as high as 0.21 during contractions. Although the formal test of Table 6 rejects the hypothesis that macroeconomic variables significantly affect the speed of adjustment, this countercyclical behaviour of the speed of adjustment seems to contradict the previous evidence reported by Drobetz *et al*. (2007) and Cook and Tang (2010).

The mean of the sample market debt ratio reported in Table 3, which amounts to 10.68 percent, is obtained using exclusively interest-bearing long-term debt. This seems to be a rather low financing ratio. An alternative leverage measure

Figure 2: TIME-VARYING SPEED OF ADJUSTMENT TOWARD TARGET LEVERAGE 1995-2007



Source: Own elaboration.

employs total debt, which also includes total short-term financial and commercial debt. The empirical results reported in Table 7 are based on the total debt and suggest very similar conclusions. Although the signs of the interaction terms associated with *GDP* growth and *TERM* are now negative, the coefficients are not statistically significantly different from zero. As before, macroeconomic conditions do not significantly impact the time-varying behavior of the speed of adjustment.

3.4. Speed of adjustment and state economic conditions

As an alternative way of estimating the impact of macroeconomic conditions on the speed of adjustment, we now follow the procedure suggested by Cook and Tang (2010). Specifically, we employ equation [8] to define the target leverage and then estimate equation [9] using the long difference estimator. This expression contains an interaction term that allows us to compare the distinguishing speed of capital structure toward target leverage between expansions and recessions. As pointed out in the preceding subsection, we classify economic states based on lagged *PER*, contemporaneous *GDP* growth and lagged *TERM* by equating good macroeconomic states with the four highest factor years, moderate states with the five middle years and bad macroeconomic states with the lowest four factor years. We estimate equation [9] for each of the macroeconomic variables separately with a good state dummy variable that takes the value of one if the firm-year observation belongs to a good state, and zero otherwise.

The results reported in Table 8 show that, independently of the macroeconomic variable employed in the dynamic panel estimation, the overall speed of adjustment is now slightly faster than that reported in Table 5. As expected, we

Table 7: ESTIMATION RESULTS OF THE PARTIAL ADJUSTMENT MODEL TOWARD TARGET CAPITAL STRUCTURES WITH TOTAL MARKET DEBT RATIO AND MACROECONOMIC VARIABLES

Explanatory Variable	Long Difference Estimator		
TMDR	0.7920 (0.1851)	0.9158 (0.1386)	0.9074 (0.0593)
ROA	–0.1851 (0.0696)	–0.1963 (0.0759)	–0.2056 (0.0731)
MTB	0.0056 (0.0071)	0.0124 (0.0057)	0.0067 (0.0025)
ETR	–0.0008 (0.0203)	–0.0065 (0.0102)	–0.0004 (0.0011)
NDTS	–1.5443 (0.8317)	–0.6130 (0.7022)	–0.1918 (0.3739)
SIZE	–0.0072 (0.0061)	–0.0091 (0.0050)	–0.0008 (0.0037)
TANG	0.0437 (0.0792)	0.0354 (0.0818)	0.0226 (0.0301)
NTANG	0.7397 (0.4038)	0.2246 (0.4181)	–0.0890 (0.1250)
IC	0.0002 (0.0001)	0.0002 (0.0001)	0.0001 (0.0001)
IDR	1.5960 (0.4288)	1.3793 (0.3518)	0.7720 (0.3774)
PER x TMDR	0.0015 (0.0089)		
GDP x TMDR	–2.0046 (3.5919)		
TERM x TMDR	–0.2909 (3.6781)		
Number of firms	101	101	101
Number of observations	853	853	853
1st order autocorrelation	–3.95 (0.0000)	–3.95 (0.0000)	–3.54 (0.0000
2nd order autocorrelation	0.83 (0.4050)	0.32 (0.7480)	0.67 (0.5030)
Hansen/Sargan test	132.94 (0.251)	131.09 (0.231)	130.03 (0.237)

Panel regression coefficients estimated from equation [7] with total market debt ratio and standard errors in parentheses. All the models include sector dummies and all the estimations are carried out by GMM using the long difference estimator of Hahn *et al*. (2007). The 1st order and 2nd order autocorrelation tests the null hypothesis of no first and second-order, respectively, autocorrelation in the residuals. The Hansen/Sargan test is a test of the overidentifying restrictions under the null hypothesis of instruments validity.

Source: Own elaboration.

also find that the estimators associated with each of the macroeconomic variables are all negative and statistically different from zero. Given that the speed of adjustment is positive by construction, this result indicates that the market debt ratio of Spanish public companies is countercyclical, as suggested by the theory. We finally find that firms adjust their leverage faster in bad states than in good states. This is a direct consequence of the positive and statistically significant coefficient estimate for the interaction term between the lagged market debt ratio and the good state dummy for all three macroeconomic factors. Given the use of a dummy variable in equation [9] to define the macroeconomic variable, when this coefficient is positive, as in Table 8, we add a positive number δ to $(1 - \lambda)$, resulting in a new speed equal to (1 – (λ + δ)). This implies that we report a faster speed toward the target during contractions. This result contradicts both the available empirical evidence and the theoretical model of Hackbarth *et al*. (2006). Table 9 replicates the estimation of equation [9] using total market debt ratio instead of the leverage ratio based only on long-term debt. The results are practically identical to the evidence reported in Table 8.

Table 8: Estimation results of the partial adjustment model toward target capital structures under expansion states with interaction effects and macroeconomic variables

Explanatory Variable	Long Difference Estimator		
MDR	0.7563 (0.0360)	0.7868 (0.0358)	0.7600 (0.0394)
ROA	–0.1908 (0.0613)	–0.1864 (0.0612)	–0.2158 (0.0612)
MTB	–0.0041 (0.0017)	–0.0036 (0.0017)	–0.0043 (0.0015)
ETR	0.0002 (0.0001)	0.0001 (0.0001)	0.0001 (0.0001)
NDTS	0.1186 (0.1538)	0.1576 (0.1751)	0.0443 (0.1554)
SIZE	–0.0013 (0.0029)	–0.0010 (0.0031)	–0.0005 (0.0026)
TANG	0.0091 (0.0136)	0.0105 (0.0139)	0.0153 (0.0130)
NTANG	–0.0168 (0.0806)	–0.0290 (0.0893)	–0.0101 (0.0859)
IC	–0.0001 (0.0001)	–0.0001 (0.0001)	–0.0001 (0.0001)
IDR	0.5574 (0.2849)	0.6768 (0.3023)	0.5099 (0.2411)
Good State PER	–0.0299 (0.0063)		
Good State PER x MDR	0.3037 (0.0584)		
Good State GDP		–0.0322 (0.0084)	
Good State GDP x MDR		0.3601 (0.0671)	
Good State TERM			–0.0281 (0.0059)
Good State TERM x MDR			0.2047 (0.0413)
Number of firms	101	101	101
Number of observations	854	854	854
1st-order autocorrelation	–4.46 (0.0000)	–4.39 (0.0000)	–4.97 (0.0000)
2nd-order autocorrelation	–1.27 (0.2040)	–1.55 (0.1200)	–0.58 (0.5590)
Hansen–Sargan test	132.00 (0.241)	132.48 (0.248)	131.17 (0.224)

Panel regression coefficients estimated from equation [9], with standard errors in parentheses. All the models include sector dummies and all the estimations are carried out by GMM using the long difference estimator of Hahn *et al*. (2007). The first- and second-order autocorrelations test the null hypothesis of no first- or second-order autocorrelation, respectively, in the residuals. The Hansen-Sargan test is a test of the overidentifying restrictions under the null hypothesis of instrument validity.

Source: Own elaboration.

Table 9: ESTIMATION RESULTS OF THE PARTIAL ADJUSTMENT MODEL TOWARD TARGET CAPITAL STRUCTURES WITH TOTAL MARKET DEBT RATIO UNDER EXPANSION STATES WITH INTERACTION EFFECTS AND MACROECONOMIC VARIABLES

Explanatory Variable	Long Difference Estimator		
TMDR	0.7833 (0.0334)	0.7999 (0.0324)	0.7191 (0.0475)
ROA	–0.2087 (0.0702)	–0.2082 (0.0722)	–0.2271 (0.0690)
MTB	–0.0071 (0.0021)	–0.0066 (0.0021)	–0.0117 (0.0024)
ETR	0.0001 (0.0001)	0.0001 (0.0001)	0.0001 (0.0001)
NDTS	0.1704 (0.1805)	0.1853 (0.1913)	0.1353 (0.2071)
SIZE	–0.0004 (0.0025)	–0.0001 (0.0025)	–0.0010 (0.0032)
TANG	0.0080 (0.0179)	0.0123 (0.0181)	0.0034 (0.0202)
NTANG	–0.0227 (0.1003)	–0.0220 (0.1115)	–0.0342 (0.1063)
IC	0.0001 (0.0001)	0.0001 (0.0001)	0.0001 (0.0001)
IDR	0.8902 (0.3259)	0.9401 (0.3284)	1.0774 (0.3885)
Good State PER	–0.0275 (0.0067)		
Good State PER x TMDR	0.1828 (0.0326)		
Good State GDP		–0.0275 (0.0086)	
Good State GDP x TMDR		0.2004 (0.0389)	
Good State TERM			–0.0390 (0.0073)
Good State TERM x TMDR			0.1439 (0.0273)
Number of firms	101	101	101
Number of observations	854	854	854
1st-order autocorrelation	–3.49 (0.0000)	–3.66 (0.0000)	–3.57 (0.0000)
2nd-order autocorrelation	0.30 (0.7640)	0.43 (0.6650)	1.26 (0.2070)
Hansen–Sargan test	126.95 (0.203)	127.42 (0.229)	134.15 (0.258)

Panel regression coefficients estimated from equation [9], with total market debt ratios and standard errors in parentheses. All the models include sector dummies and all the estimations are carried out by GMM using the long difference estimator of Hahn *et al*. (2007). The first –and second-order autocorrelations test the null hypothesis of no first– or second-order autocorrelation, respectively, in the residuals. The Hansen-Sargan test is a test of the overidentifying restrictions under the null hypothesis of instrument validity.

Source: Own elaboration.

As already discussed, our data do not really capture strong fluctuations in the business cycle. Other reasons may easily explain why the speed of adjustment is faster during states identified as contractions. One possibility is that firms farther away from target leverage adjust faster toward their optimal market debt ratio. We follow Cook and Tang (2010) to examine the differences in the mean absolute values of deviations from the target leverage between good and bad states to test whether Spanish companies deviate farther in bad states than in good states. If this is the case, the results reported in Table 8 can be attributed to the distance between current and target market debt ratios rather than to macroeconomic conditions. This case would also be consistent with the results reported in Tables 6 and 7.

The distance between target and actual leverage is defined as

$$DIS_{jt} = \left| MDR^{*}_{jt} - MDR_{jt-1} \right| , \qquad [13]$$

where, $MDR^{*}_{jt} = \beta' X_{jt-1} + \gamma M_{t-1} + \eta M_{t-1} x MDR_{jt-1}$ and the parameters are obtained from the estimates reported in Table 8. The results are contained in Table 10. At least with respect to *GDP* growth and *TERM*, firms are statistically farther away from their target market debt ratio in bad states than in good states. This suggests that the countercyclical nature of the speed of adjustment reported in this paper should not be attributed to the response of Spanish firms to macroeconomic conditions but, rather, associated with how far away Spanish firms are from their target leverage.

Table 10: Distance between target and actual market debt ratios

Economic States	PER	GDP	TERM
Good	0.192	0.186	0.165
Bad	0.197	0.232	0.187
Good–Bad	-0.005	-0.046	-0.022
p-Value	0.194	< 0.0001	< 0.0001

This table reports the distance between target and actual market debt ratios across good and bad economic states over the sample period 1995-2007. We measure the distance as the difference between the absolute value of the target leverage and the actual leverage, where target leverage is defined by equation [8] and the parameters employed are the long difference estimators of the dynamic panel equation [9].

Source: Own elaboration.

4. Conclusions

This paper employs the long difference estimator of Hahn *et al*. (2007) to estimate the speed of adjustment of Spanish public firms toward target debt ratios. The long difference estimator is known to have less bias than the more popular system GMM estimator. We find that Spanish firms adjust slowly toward their target leverage. The speed of adjustment using the long difference estimator is 17.5 percent, which implies that only one-fifth of the gap between the actual and target debt ratios is closed each year. As expected, the estimator of the speed of adjustment under the system GMM suggests a faster speed toward target leverage. It is estimated to be 24.9 percent, which is closer to the latest evidence available for U.S. firms.

We also analyze the importance of macroeconomic conditions on the speed of adjustment. The results suggest that Spanish public firms move faster toward their target market debt ratios during contractions than during expansions. This is especially the case when we introduce macroeconomic conditions in the definition of target leverage. However, there is also some evidence of a countercyclical speed of adjustment when we allow the speed of adjustment to be time varying. This result contrasts with previous international empirical evidence and with the theoretical model of Hackbarth *et al*. (2006). Given the weak economic fluctuations characterizing the Spanish economy during our sample period, we measure the distance between target market debt ratios and actual leverage to check whether the previous result is a direct consequence of Spanish firms being farther away from target during relatively bad economic times. This turns out to be the case. We attribute the countercyclical nature of the speed of adjustment during our sample period to the distance between actual and target leverage rather than to macroeconomic conditions.

Appendix. Some econometric issues in dynamic panel data models

Consider a dynamic panel data that can be understood as a simplified version of equation [4]

$$MDR_{jt} = \beta MDR_{jt-1} + \left(\mu_j + \varepsilon_{jt}\right), \quad \text{[A.1]}$$

where the residual term is composed of the unobserved time-invariant, firm-specific effect captured by μ_j (fixed effect) and the usual residual given by ε_{jt}.The problem is that the residual component of MDR_{jt} is correlated with the unobserved effect in the error term. Hence, a (pooled) ordinary least squares (OLS) estimated coefficient on MDR_{jt-1} without fixed effects will be upwardly biased (or the speed of adjustment downwardly biased). These simple regressions can also be estimated by the two-step procedure of Fama-MacBeth (FM hereafter, 1973), as suggested by Fama and French (2002). The authors recommend FM estimators to mitigate the underestimation of coefficient standard errors. Flannery and Rangan (2006) show that both the traditional (pooled) OLS and FM regressions yield similar coefficient estimates but the *t*-statistics are much smaller when the FM estimates are used in the regression. The problem is that the FM estimates also fail to recognize the data's panel characteristics. As pointed out above, if cross-sectional differences in target market debt ratios are driven by time-invariant, firm-specific components, either the FM procedure or the (pooled) OLS estimates will be upwardly biased and the speed of adjustment downwardly biased.

The FM estimation procedure runs, for each observation *T,* an OLS cross-sectional regression with *N* observations (firms):

$$MDR_{jt} = \beta_{t-1} MDR_{jt-1} + u_{jt} \quad ; j = 1,\ldots,N . \quad \text{[A.2]}$$

We keep the estimates of each beta, $\hat{\beta}_{t-1}$, for each cross-sectional regression from $t = 2$ to $t = T$. The final estimates are

$$\hat{\hat{\beta}} = \frac{1}{T-1}\sum_{t=2}^{T}\hat{\beta}_{t-1}$$

$$\hat{\sigma}\left(\hat{\hat{\beta}}\right) = \sqrt{\frac{\sum_{t=2}^{T}\left(\hat{\beta}_{t-1} - \hat{\hat{\beta}}\right)^2}{(T-1)(T-2)}} \qquad \text{[A.3]}$$

This discussion implies that a panel regression with unobserved (fixed) effects is more appropriate if firms have relatively stable and unobserved variables affecting their leverage targets. There are two simple possibilities to adjust for this bias. A common way to estimate panel data models with fixed effects is to perform a "within" transformation of (A.1) and then estimate using OLS. This transformation expresses all variables as deviations from their firm-specific time series means. This eliminates μ_j from the regression, since it is time invariant and thus provides consistent estimates. However, it should also be noted that the within transformations of the lagged dependent variables and error terms are

$$MDR_{jt} - \frac{1}{T}\sum_{t=1}^{T}MDR_{jt} \quad \text{and} \quad u_{jt} - \frac{1}{T}\sum_{t=1}^{T}u_{jt}\,. \qquad \text{[A.4]}$$

Then, MDR_{j2} is correlated with u_{j2}, MDR_{j3} is correlated with u_{j3}, and so on. This implies that the coefficient of the lagged dependent variable, β, is downwardly biased (the speed of adjustment is upwardly biased) by a factor of (approximately) $1/T$. In panel data sets with large T, this bias becomes insignificant, but in panel data with capital structure observations, large N, and a relatively small T, the bias can be substantial. A similar alternative procedure consists of running FM cross-sectional regressions for each t once we express all variables as deviations from their firm-specific time series mean. This straightforward alternative can also be downwardly biased (upwardly biased for the speed of adjustment).

It is possible to obtain unbiased estimates of the levels regression (A.1) if an instrument is found that is correlated with the lagged dependent variable but not with the error term. We can use the lagged book debt ratio as the instrument. The new estimate should be between the OLS and the within estimates. This is the procedure employed by Flannery and Rangan (2006) in their Tables 2 (columns 5 to 7) and A.1 (column 3). In particular, the authors use a two-stage least squares in which they substitute a fitted value for the lagged dependent variable, using the lagged book value of leverage and the X's, from equation [3], as instruments. They first regress the lagged market debt ratio on the lagged book debt ratio and the X's, and then the fitted value from that regression is employed in the second stage as the independent variable. We can then regress the market debt ratio on the previous fitted value of the lagged market debt ratio and the control variables. A serious drawback is that the previous correction relies on the book debt ratio being a reasonable instrument for the market debt ratio. Of course, finding reliable instruments can be difficult and a number of alternative econometric possibilities are available in literature.

These techniques generally first-difference the model (A.1) to eliminate fixed effects and use lagged dependent variables to instrument for the lagged first difference:

$$MDR_{jt} - MDR_{jt-1} = \beta\left(MDR_{jt-1} - MDR_{jt-2}\right) + \left(\varepsilon_{jt} - \varepsilon_{jt-1}\right). \qquad [A.5]$$

It is also well known that these first-difference methodologies rely on two key assumptions to produce unbiased and consistent estimates. First, the error term in (A.1), ε_{jt}, should be serially uncorrelated, because the first-order serial correlation would make the lagged dependent variable correlated with the (differenced) regression residual. In other words, lags of the dependent variable fail the exogeneity assumption if the residual is serially correlated. Moreover, the dependent variable should not have (near) unit root properties. If the dependent variable has high persistence (as is usually the case in these studies), then the first difference will be close to zero and the instruments used will be weak. Unfortunately, both the first-difference approach of the Anderson-Hsiao (1981) instrumental method and the GMM procedure of Arellano and Bond (1991) are unlikely to yield consistent results.

To avoid these additional difficulties, Arellano and Bover (1995) and Blundell and Bond (1998) propose a system GMM (extended GMM) that imposes additional moment conditions. It has been argued that it performs better with persistent data series than first-differencing estimators. In fact, it has been used by Antoniou *et al*. (2008), Faulkender *et al*. (2008), and Lemmon *et al*. (2008). We first take the first difference of equation (A.1) to obtain (A.5), and both equations (A.1) and (A.5) are then simultaneously estimated as a "system." The estimator uses the lagged differences ($MDR_{jt-2} - MDR_{jt-3}$, ..., $MDR_{jt-1} - MDR_{jt-0}$) as instruments for equation (A.1), and the lagged levels ($MDR_{jt-2} - MDR_{jt-0}$) as instruments for equation (A.5). The practical problem is that the set of moment conditions for the system GMM estimator tends to explode as the time series increases. Moreover, the estimates may also be sensitive to the choice of instruments. Hahn *et al*. (2007) show that the use of the full set of moment conditions does not provide proper guidance in the dynamic panel data model when the autoregressive parameter is close to one.

Recently, Hahn *et al*. (2007) and Huang and Ritter (2009) have argued that the so-called long differencing estimator is much less biased than the GMM estimator. This estimator alleviates the problem of weak instruments and relies on a less than full set of moment conditions. To understand their estimation procedure, let us assume that the market debt ratio at the end of year $t - k$ is given by

$$MDR_{jt-k} = \beta MDR_{jt-k-1} + \left(\mu_j + \varepsilon_{jt-k}\right). \qquad [A.6]$$

Subtracting equation (A.6) from equation (A.1), we obtain

$$MDR_{jt} - MDR_{jt-k} = \beta\left(MDR_{jt-1} - MDR_{jt-k-1}\right) + \left(\varepsilon_{jt} - \varepsilon_{jt-k}\right). \qquad [A.7]$$

The GMM estimation employs MDR_{jt-k-1} as a valid instrument. Using this instrument, we can first estimate equation (A.7) to obtain the initial $\hat{\beta}$. Then we obtain the residuals and use ($MDR_{jt-1} - \hat{\beta}MDR_{jt-2}$), ..., ($MDR_{jt-k} - \hat{\beta}MDR_{jt-k-1}$) and those residuals as instruments. This is the first iteration. Both Hahn *et al*. (2007) and Huang and Ritter (2009) use three iterations to obtain the final $\hat{\beta}$. Additionally, using data from 1972 to 2001, Huang and Ritter (2009) try four alternative lag values for k and show that the coefficient of the speed of adjustment varies from 22.3 percent when $k = 4$ to 17.6 for $k = 28$.

In particular, for the long difference estimator without macroeconomic variables, the dynamic equation at the end of year t with fixed effects is given by

$$MDR_{jt} = (\lambda\beta)' X_{jt-1} + (1-\lambda) MDR_{jt-1} + (\mu_j + u_{jt}). \qquad [A.8]$$

The dynamic equation at the end of year t-k with fixed effects is given by

$$MDR_{jt-k} = (\lambda\beta)' X_{jt-k-1} + (1-\lambda) MDR_{jt-k-1} + (\mu_j + u_{jt-k}). \qquad [A.9]$$

Then, subtracting (A.9) from (A.8), we get the equation to be estimated under the long difference estimator:

$$MDR_{jt} - MDR_{jt-k} = (\lambda\beta)' \left(X_{jt-1} - X_{jt-k-1}\right) + (1-\lambda)\left(MDR_{jt-1} - MDR_{jt-k-1}\right) + \\ + \left(u_{jt} - u_{jt-k}\right) \qquad [A.10]$$

The instruments would be the residuals as in (A.7), MDR_{jt-k-1} and X_{jt-k-1}.

Finally, the equation in which we recognize the potential effects of macroeconomic conditions would be

$$MDR_{jt} - MDR_{jt-k} = (\lambda_0\beta)' \left(X_{jt-1} - X_{jt-k-1}\right) + (1-\lambda_0)\left(MDR_{jt-1} - MDR_{jt-k-1}\right) \\ - \lambda'\left(Y_{jt-1} - Y_{jt-k-1}\right) + \eta'\left(Z_{jt-1} - Z_{jt-k-1}\right) + \left(u_{jt} - u_{jt-k}\right) \qquad [A.11]$$

The instruments would be the residuals as in (A.7), MDR_{jt-k-1}, Y_{jt-k-1}, and X_{jt-k-1}.

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RESUMEN

Nuestra evidencia sugiere que las empresas cotizadas españolas realizan un lento ajuste hacia su estructura de capital objetivo. La empresa típica reduce aproximadamente, cada año, un quinto de la distancia existente entre su ratio de endeudamiento y su nivel óptimo de deuda. Este resultado parece contrario a la evidencia previa existente para empresas españolas. Sin embargo, nuestros resultados están basados en procedimientos de estimación especialmente diseñados para series de alta persistencia, como es el caso de los ratios de endeudamiento empresarial. Asimismo, nuestra evidencia no parece apoyar que las condiciones macroeconómicas, al menos las experimentadas por la economía española durante el periodo muestral disponible, afectan a la velocidad de ajuste. Parece más probable que estemos observando una mayor rapidez hacia el endeudamiento objetivo de las empresas en aquellos estados en los que los porcentajes de endeudamiento están más alejados del óptimo.

Palabras clave: ratio de endeudamiento, intercambio dinámico, endeudamiento objetivo, velocidad de ajuste, macroeconomía, distancia desde el endeudamiento objetivo.

Clasificación JEL: G32, C33, E30.