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COLLECTIVE ACTION CLAUSES: HOW DO THEY WEIGH ON SOVEREIGNS?

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Abstract

We study the effects of the adoption of collective action clauses (CACs) on government bond yields by exploiting secondary market data on sovereigns quoted in international markets from March 2007 to April 2011. CACs are assessed security by security. Using a panel data approach, we find a U-shaped effect of CACs on yields according to credit rating of the issuer. While the impact is negligible for the highest ratings, there emerges a significant yield discount for mid-ratings, which is smaller for bad ratings and possibly insignificant for the worst ratings. The relationship appears fairly robust across a number of robustness checks. This evidence may reflect the fact that CACs are valuable as they help orderly restructuring unless the perceived probability of default is too small. Nevertheless, at low ratings this relevance can be weakened by an increasing moral hazard risk.

JEL Classification: F34, G15, H63.
Keywords: collective action clauses (CACs), sovereign yields, debt restructuring, default, panel data.

Contents

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1. Introduction

Collective action clauses (CACs) are contractual provisions allowing a qualified majority of bondholders to change the terms of a bond in a way which is legally binding for all holders, in order to facilitate debt restructuring.[1] Since the end of the 1990s, a relatively large number of papers have empirically addressed the relationship between the adoption of CACs on sovereign bonds and their yields.[2] Yet, no consensus has yet emerged on the sign of the link or even about the conditions under which such link could exist. In fact, disagreement among authors arises even on the methodology to follow when conducting the empirical analysis and the nature and structure of the dataset. This is a rather uncomfortable situation, as time is nearing when one of the biggest experiments in the field – the adoption of standardised mandatory CACs on all new euro area sovereign bonds from January 2013[3] – will be implemented.

Against this background, the ambition of this paper is to take stock of a number of lessons gained on both fronts – methodology and dataset – and to offer a wider encompassing approach to the issue of testing the relationship between the adoption of CACs and the bond yields.

We exploit a dataset running from March 2007 to April 2011 with yields on 292 securities listed on major international markets. Thanks to a new feature added by Bloomberg, we are able to determine for each bond whether a CAC is in place or not, overcoming one of the main pitfalls of many earlier studies, which relied on the bond's place of issue – whether New York or London – as a proxy to gauge the adoption of the CAC.

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[1] Examples of collective action clauses include those stating what majority has to be met in order to change the payment terms; those requiring a minimum threshold to initiate litigation or accelerate the bond; those allowing a qualified share of bondholders to prevent acceleration (so-called "deceleration clauses"); those stating if and how bondholders strictly related to the issuer should be excluded from the computation of quorum (so-called "disenfranchisement clauses"); etc.

[2] Though the origin of CACs can be traced back to XIXth century, there was a revival of interest starting in mid-Nineties when CACs were suggested by the official sector and by researchers as a device to reduce resorting to public bailouts.

[3] The main features of the CACs associated to euro area government bonds have been defined consistently with those commonly used under New York and English law. The introduction of standardised CACs aims to promote a uniform legal impact in all euro area countries, despite different legal systems and traditions, in order to preserve a level playing field. The detailed legal arrangements for the inclusion of CACs in euro area government securities have been developed by the EU's Economic and Financial Committee via the Sub-Committee on EU Sovereign Debt Markets. The work of the Sub-Committee was finalised in March 2012, following consultations with market participants and other stakeholders. Euro area CACs will enable creditors to take a qualified majority decision agreeing a legally binding change to the terms of payment when the debtor is unable to pay. Aggregation clauses will be included, allowing several securities issued by the same euro area country to be considered together in negotiations.

The sample is large enough to allow us to focus on sovereigns, enhancing comparability (this means that we choose not to enlarge the dataset with corporates, which could give rise to forms of spurious correlations). Our study encompasses a relatively large number of countries at various stages of development. This stands in sharp contrast with most of previous works, which focussed on emerging market issuers.

As a further add-on, previous studies (e.g., Becker *et al.*, 2003) have stressed the need for using secondary market data rather than primary market ones, as a way to mitigate instances of endogeneity, structural breaks and omitted variable bias. We follow suit compiling our dataset with secondary market yields (average of bid and ask rates) taken monthly, for a total of 50 time periods. The exceptional market patterns occurred throughout the sample should ease the identification of any effect of CACs on yields: in other words, if no such link emerges in the data at a time when the default of sovereigns was more than a marginal tail probability, then doubts should be cast on whether CACs do affect in any way the return asked by the investor when purchasing a security.

The bulk of our empirical analysis is about the estimation of a panel model. This is a relatively novel approach in this strand of literature, as most previous papers focused on a snapshot taken at a given time. The extension of the period under scrutiny (on top of the cross-section) offers two clear advantages: (i) it renders the analysis less dependent on the idiosyncrasies in the data at any specific point in time and (ii) it allows checking if and how the link under examination has evolved with market developments (e.g., the impact of a downgrading of the country issuing the bond).

Naturally, caution must be exerted when judging the aptitude of any empirical analysis to predict the outcome of the euro-area experiment, simply because there is no precedent of a number of rich countries adding standardized and identical CACs on their domestic bonds all at once. Nevertheless, the reliance on a broader range of issuers allows examining in better detail questions like the impact of such clauses on securities rated double or triple A, to name a few.

Anticipating the gist of our results, we find that the inclusion of CACs lowers most yields for bonds whose issuers fall in the middle of the rating scale. For very good ratings, no statistically significant difference in yields emerges due to the use of CACs, while for bad ratings the yield discount is smaller than the one for mid-ratings, to the point of becoming insignificant for the lowest ratings. This relation appears to hold across several robustness checks.

These results hint that collective action clauses are *ex post* useful for an ordered restructuring, should a default occur. So, a first condition is that the probability of default has to be non-negligible. The lack of this requirement helps explaining why, for best-rated countries, CACs do not seem to have any effect. Second, by making a default easier, CACs might also make it more likely, if the issuer behaves opportunistically (the so-called *moral hazard* effect). The second requirement is thus that the greater *ex ante* moral hazard does not offset the *ex post* benefit. For worst rated countries, the moral hazard-enhancing effects of CACs is arguably larger than for better rated countries.

The rest of the paper is organised as follows. In Section 2 we review the empirical literature on the effect of collective action clauses on bond yields; in Section 3 we present the dataset and show some descriptive statistics. In Section 4 we report the econometric analysis on the panel data and comment on the main results; while in Section 5 we address several issues related to sensitivity analysis. Finally, Section 6 concludes.

2. Literature review

Previous studies on the effect of collective action clauses on bond yields depict a number of different approaches, with respect to either the methodology or the dataset used. Subsequent research has often moved from criticizing the pitfalls of previous works, either from a methodological point of view, on the ground of sample-construction or both.

The first systematic study on the yield effect of CACs is acknowledged to Tsatsaronis (1999), who considered primary market data on a variety of international sovereign bonds issued after 1990. Since information on CACs was not available at bond level, the governing law of issuance was used as a proxy, *i.e.*, all bonds issued under UK's governing law were assumed to be endowed with collective auction clauses, while those issued under the US' were not, in accordance with the common practice in those countries. The author finds some evidence that CACs measured that way imply greater yields, but the difference is not statistically significant. Eichengreen and Mody (2000) assess the impact of CACs on borrowing costs, recognising the importance of controlling for endogeneity in the choice of governing law (used, as usual in the first strand of the literature, as a proxy for the very presence of CACs). Using the same proxy-dummy for CACs and primary market data on a wide set of bonds including corporates, the authors find that CACs reduce the interest burden for more credit-worthy issuers, arguing that well rated borrowers may benefit from issuing bonds subject to renegotiation-friendly governing law. In a later article Eichengreen and Mody (2004) focus on different sub-samples according to the rating group of the issuer: they find that CACs reduce yields for well-rated issuers but rise them for bad-rated ones, suggesting that for the latter the moral hazard risk implied by CACs is heavier than any benefit.

Becker *et al.* (2003) point out a number of pitfalls stemming from the use of primary market data, arguing that secondary market data should be preferred: the latter arguably rise less endogeneity issues (whereas the former would require modelling the supply side too), need less control variables (since there is no need to account for general market conditions changing over time), and, presumably, benefit of more accurate data. However, selecting two dates, one in 1998 one in 2000, they get contrasting results: in the first one, CACs lower well rated bond yields and have no effect for bad-rated countries, while in the second CACs rise well-rated countries' yields and lower bad-rated borrowers' (which is basically the opposite of what Eichengreen and Mody found in 2004). Pooling data together, no significance emerges. The methodology of Becker *et al.* (2003) is followed by Gugiatti and Richards (2003), who consider for the first time post-2003 data, after the extensive debate on the inclusion of CACs in a big Mexican issuance in US dollars. The authors detect a negative effect of CACs, which nevertheless disappears when the interaction with rating is introduced.

Eichengreen *et al.* (2003) add four additional points in time, selected in correspondence with very high or very low levels of market credit risk premium. To reconcile previous contrasting results, they consider a triple interaction term between the risk premium mirrored by bond spreads, CACs and ratings, to capture the idea that the rating value at which the effect of CACs on yields turns from negative to positive depends on market sentiment. However, assessing the effects of CACs using a triple interaction term is not straightforward: the suggested interpretation of the random effect estimation is that, in good times, CACs are beneficial for all but the worse issuers, while in bad times, CACs penalize all but the best issuers. Nevertheless, Häseler (2009) points out that this is not a true reconciliation of previous findings, nor the results seem to be robust since, for instance, the coefficient on the triple interaction is barely significant. Moreover, only four time snapshots are considered, thus shedding doubts on the legitimacy to fully extend the relation with market sentiment; moreover, the validity of random effects estimation should be tested.

A more substantial criticism is that posed by Gugiatti and Richards (2004), who question the goodness of the governing law dummy as a proxy for the inclusion of CACs. Through a detailed inspection of a sample of bonds documentation they find that, already well before 2003, several bonds under the US governing law were endowed with CACs, while several securities under UK governing law were not.[4] The fallacy of the governing law proxy casts serious uncertainty on the extent results from previous literature were affected by that spurious identification.

This scepticism somehow slowed the pace of research on the topic. Only after some years the exploration of the effect of collective action clauses on yields was resumed by Bradley *et al*. (2008), who restricted the analysis to US law-issued bonds and investigated contractual terms to detect the actual presence of CACs. However, the sample remained quite heterogeneous, as it was made up of primary market data on bonds issued since 1986, thus encompassing very different frameworks and

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4 Gelpern and Gulati (2008) also document a number of bonds endowed with CACs under the New York law which were issued before 2003.

situations and being exposed to the afore mentioned criticism on primary market data. The authors found some weak evidence of a negative effect of CACs on yields, but they warned that that might have hidden a more general structural break between the pre- and post-2003 periods (when post-2003 time dummies are added, the coefficient on the CACs dummy loses significance). Moreover, their study did not account for the interaction between CACs and rating. Finally, Häseler (2010) took advantage of the newly available CACs field in Bloomberg inquiries to collect a sample of secondary market data not limited to emerging markets, as customary in previous literature.[5] However, the effects of CACs are not the main focus of this study, so that the CACs dummy is included in one specification only (finding a negative but non-significant coefficient) and its interaction with issuer rating remained unexplored.

3. The dataset and some group-wise distribution of CACs

Securities in the dataset are selected from Bloomberg according to the following criteria: issuance after January 1, 2003, denominated in US dollars, issued in Global, Euro MTN and Eurodollar markets;[6] issuer being either a national or a regional government; maturity type being either *bullet*, *putable* or *callable;*[7] bonds being either zero coupon or having a fixed or floating coupon. Besides the features described above, for each security the following fields are downloaded:[8] inclusion of CACs ; maturity date; amount issued; registration at SEC; issuer's country. The dataset is described in more detail in Table 1.

Yields are collected on a monthly basis from March 2007 to April 2011, totalling 50 observations.[9] In Figure A.1 of the Appendix, yields over time are shown through their mean, median, first and third quartile.

The bonds duration and the credit rating of the issuer country assigned by the three main rating agencies: Standard & Poor's, Fitch and Moody's are recorded over the same timespan. Ratings are mapped into a numerical variable starting from 1 for the best rating and increasing by 1 for each notch towards the worst rating. The composite rating variable used in the econometric

5 Analysing the Bloomberg inquiry across countries and time, Häseler (2010) confirms that previous studies misclassified more than 20%, perhaps up to one third of bonds, by assuming a straight correspondence between the use of CACs and the governing law before 2003.

6 The Global Bond market refers to bonds issued and traded outside the country and outside the regulations of a single country. In Euro MTN market bonds are traded that require fixed dollar payments issued and traded outside the US and Canada; securities in this market are issued with maturities of less than five years and are generally part of a program. Eurodollar bonds are US-dollar denominated bonds held in a foreign institution outside both the US and the issuer's home nation.

7 With a bullet maturity the face value is paid on the maturity date of the bond. When maturity is callable the issuer has the faculty of redeeming the bond prior its natural maturity date; on the contrary, with a putable maturity the investor may mature the bond at an earlier date than the natural expiration date.

8 Other bond features are downloaded but not used in the analysis, as they displayed no variation with respect to the inclusion of CACs (dummies for private placement, subordinated debt, zero coupon, collateralized debt).

9 Yields are taken for each month at closing quote of the Friday before mid-month to limit irregularities related to calendar issues potentially occurring at the end/beginning of month or at the beginning of the week.

model is obtained averaging across the three agencies. Other time series variables collected at the same dates are the VIX index as a proxy for market volatility, the spread between triple A and triple B US corporate bonds as a measure for credit risk market premium, and the US benchmark yield at 10 years to account for patterns in the general level of yields over time.

Table 1

Dataset variables description

Variable	Unit	Description	Series features
Y	%	Mid-yield closing quote	Bond specific, time variant
CAC	Dummy	1 if bonds include CACs	Bond specific, time invariant
Size	Millions	Issued amount	Bond specific, time invariant
Duration	Years	Bond duration at date	Bond specific, time variant
Rating	Index. Scaled from 1 (AAA) to 18 (CCC)	Average of rating given by S&P, Fitch and Moody's	Country specific, time variant
Region	Dummy	1 if issued by regional government	Bond specific, time invariant
Call	Dummy	1 if bond is callable	Bond specific, time invariant
Put	Dummy	1 if bond is putable	Bond specific, time invariant
MTN /Eurodollar	Dummies	1 if traded in MTN / Eurodollar market (Global as reference category)	Bond specific, time invariant
SEC	Dummy	1 if registered at SEC	Bond specific, time invariant
VIX	Index	Chicago Board Option Exchange Market Volatility Index	Bond invariant, time variant
BBB-AAA spread	b.p.	Spread between AAA and BBB US corporate bonds 5 years tenure (bp)	Bond invariant, time variant
Bmk 10	%	US benchmark yields at 10 years tenure	Bond invariant, time variant

Source: Bloomberg.
Number of bonds: 292.
Number of time periods: 50.
Note: Time range: March 2007 to April 2011, monthly observations. All time-dependent variables are taken at closing value of closest Friday before mid-month.

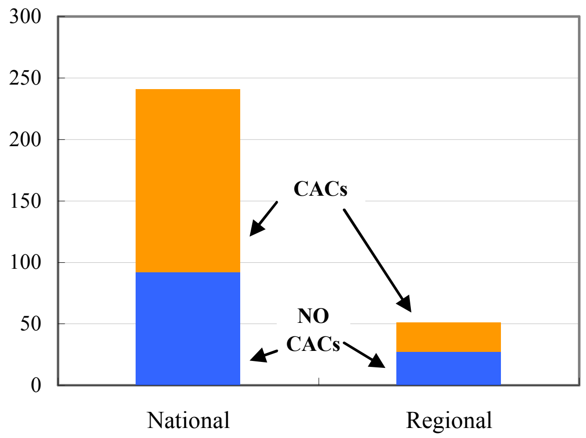
From a preliminary tabulation of the distribution of CACs across the dataset variables, a number of descriptive findings can be identified (see Figure 1). No clear-cut divide emerges for the use CACs depending on **issuer's type** or **issuance year**. Both national and regional governments have issued bonds with and without clauses, although states tend to do so relatively more frequently (62% of the sovereign securities in our dataset feature a CAC, vs 47% of regional or other instances of local government). The proportion of CACs in new issuances followed a decreasing trend until 2008, reverting to positive thereafter. Within the dataset, we have **countries** which never rely on CACs (Austria, Germany, Hungary), and countries which always or almost always adopt them (Brazil, Columbia, Peru, Uruguay). There exists also a mixed breed, as some countries have followed either one of the two approaches (e.g., Italy, Canada, Venezuela, Turkey, Indonesia, etc.). There does not seem to be a sort of turning point when the sovereign switches from one approach to the other, as they seem to continue to intertwine over time.

Figure 1

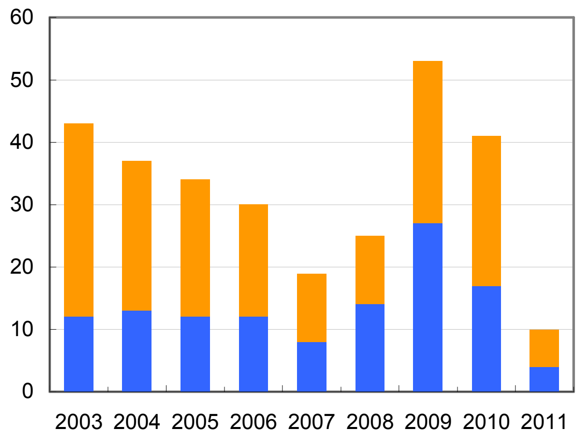
Distribution of Collective Action Clauses

(number of bonds in the sample)

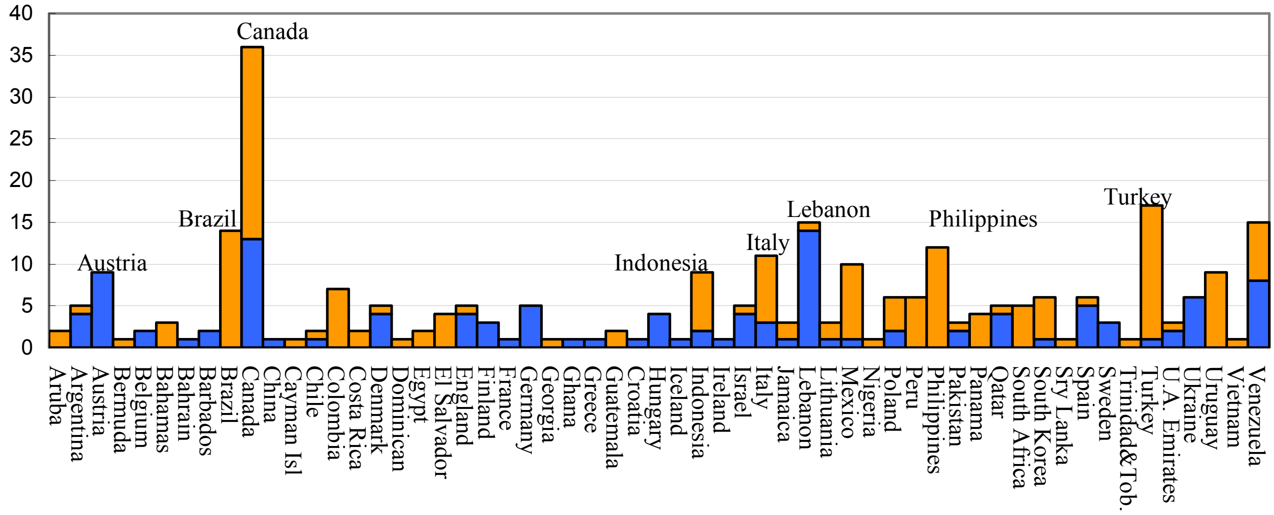
(a) by issuer



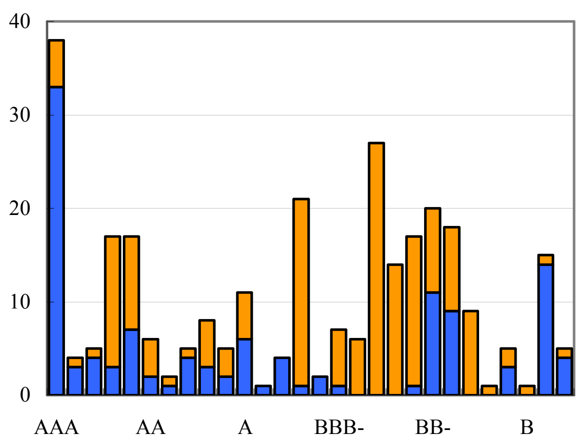
(b) by issuance year



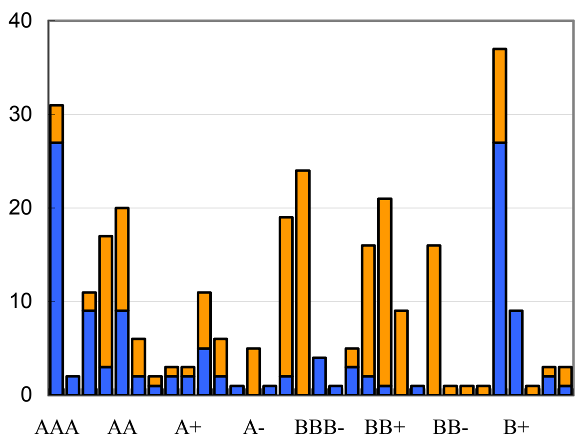
(c) by country



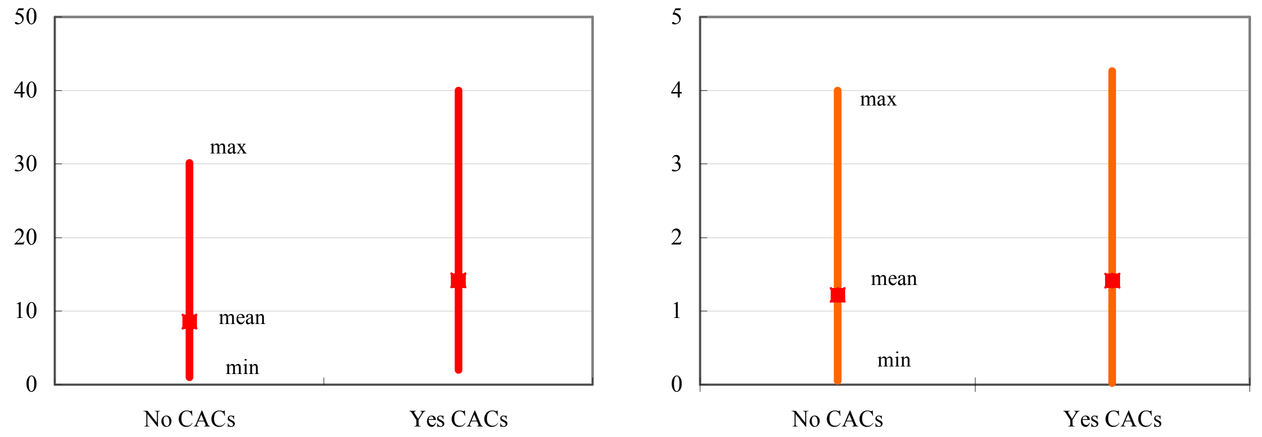
(d) by initial rating distribution (Mar 2007)



(e) by final rating distribution (Apr 2011)



(f) by maturity at issuance *(years)* (g) by amount at issuance *(billions)*



CACs are relatively more frequent for issuers in the middle of the **rating scale**, while very well rated and very badly rated issuers make less use of CACs. This will be a recurring theme in our analysis, as the advantages in bringing in such clauses may be diverse in a non-linear way depending on the issuer's rating, while those with a middle ranking could gain most. CACs are more likely to be present in bonds of longer **maturity**: the average tenure at issuance of securities endowed with collective action clauses is about 14 years and 2 months versus 8 years and 7 months for those without CACs. This could be related, *ceteris paribus,* to greater uncertainty and hence higher credit risk for bonds with longer maturity. There is no appreciable difference, instead, as far as the issued **amount** is concerned.[10]

4. The econometric analysis

The econometric analysis is carried out taking a panel data approach over three time sub-sets, which can be related to different economic situations. The first interval ranges from March 2007 to September 2008, in coincidence with the Lehman Brothers collapse. The second interval starts from October 2008 and ends in April 2010, including the aftermath of the financial crisis, the partial recovery from the banking crisis when market concerns about sovereigns initially were still in the background, to emerge thereafter. The third period starts in May 2010, when sovereigns came clearly under the spotlight.[11] This time division is also useful to take advantage from panel data

[10] Finally, as shown in Figure A.2 in the Appendix, the sample is made up mainly by bonds traded in the Global market (78% of which include CACs; while such a share is more balanced in the Eurodollar market: 50%). More than 90% of bonds in the sample have a standard bullet maturity, and little more than a half of them are endowed with CACs. Among securities not registered with the SEC (which make up almost 70% of sample), CACs are relatively more frequent than among registered bonds.

[11] While the first break-point is somehow uncontroversial, given the symbolic identification of the financial crisis with the Lehman collapse, the second one is, admittedly, more arbitrary. Though there is no clear-cut divide, we picked

techniques without increasing too much the time dimension and imposing constant coefficients over a long period. In Section 5 we also consider repeated cross-section regressions for each period as a robustness checks for the time pattern shown by estimated coefficients.

We perform the estimation allowing for non-linearities. In so doing, we do not split the sample in sub-groups of ratings, like done by others (Eichengreen and Mody, 2004, Becker *et al.*, 2003). Keeping the whole sample allows to gain in efficiency and helps detecting the genuine non-linear effect (if any) of collective action clauses on yields, as other controls variables are forced to have the same effect across ratings, and we do not have to make arbitrary sample restrictions or divides.

4.1. The benchmark econometric model

In order to take into account possible non-linearities in the impact of CACs according to the rating merit of the issuer, we consider two econometric specifications which permit such effects: in the first one we include the interaction of the CAC dummy with a linear distance from the middle of the rating scale for issuers either above or below it; in the second specification we consider a quadratic form.

The basic panel model is specified as follows:

$$y_{i,t} = \alpha + \boldsymbol{\beta' CACS\_VARIABLES_{i,t}} + \boldsymbol{\gamma' \, CONTROL\_VARIABLES_{i,t}} + \eta_i + \varepsilon_{i,t} \qquad (1)$$

where the dependent variable is the log of the mid-yield in the secondary market of security i at month t, $\eta_{i,t}$ represents a unit-specific time invariant effect, while $\varepsilon_{i,t}$ is an idiosyncratic error which varies both across space and time. The *CACS_VARIABLES* includes the set of the main variable of interest and depends on the specification (see below), while the *CONTROL_VARIABLES* vector includes:

- $RATING_{i,t}$, a measure of the rating of the security i at time t along a scale which takes value 1 when rating is AAA and increases at each notch (say, 2 for AA+, 3 for AA, etc.);
- $SIZE_i$, the log of the amount issued of the security;
- $REGION_i$, that tells whether the issuer is a local government (in which case the variable takes value 1) or not (value 0);
- $DURATION_{i,t}$ and $DURATION\_SQ_{i,t}$, compiled with bond's duration and its square;

the one between April and May 2010 as it is characterized by a number of considerable events and decisions: tensions on peripherals EU sovereign debts sharpened showing evidence of contagion, financial support to Greece was agreed (May 2), the Security Market Program was introduced (May 10), the euro-dollar swap line was reactivated (May 10), etc. Moreover, May 2010 marked the beginning of the widening of the spread with Eonia swap rates in opposite direction for core countries and peripherals (see Di Cesare *et al.*, 2012).

- PUT_i and $CALL_i$, dummy variables which take value 1 if the related options are embedded in the security and 0 otherwise;
- $EMTN_i$ and $EURODOLLAR_i$, that specify the market where the security is negotiated with the Global Market being the reference point;
- SEC_i , highlighting whether the bond is subject to registration with the SEC;
- VIX_t , the VIX index at time t to capture market volatility;
- $BBBAAA_t$, an index of the spread between triple A- and triple B-rated corporate bonds which may be regarded as a proxy for general market premium for credit risk;
- $BM10Y_t$, the value of benchmark US yields at time t to account for general movement in yield levels.

In the linear specification, the *CACS_VARIABLES* vector includes:

- CAC_i , a dummy variable which takes value 1 if the clause is adopted and 0 otherwise;
- $CACDISTGOOD_{i,t}$ and $CACDISTBAD_{i,t}$, the interaction between CAC_i and the linear distance of issuer rating ($RATING_{i,t}$) from the middle rating either for better ($CACDISTGOOD_{i,t}$) or worse ($CACDISTBAD_{i,t}$) rated issuers, as in Table 2:

Table 2

Linear specification: CACDISTGOOD and CACDISTBAD

(interaction of CAC with linear distance from the median value of the rating scale)

Rating notches	*RATING*	Linear distance from median value	*CACDISTGOOD*		*CACDISTBAD*	
			CAC=Y	*CAC*=N	*CAC*=Y	*CAC*=N
AAA	1	8	8	0	0	0
AA+	2	7	7	0	0	0
AA	3	6	6	0	0	0
AA-	4	5	5	0	0	0
A+	5	4	4	0	0	0
A	6	3	3	0	0	0
A-	7	2	2	0	0	0
BBB+	8	1	1	0	0	0
BBB	9	0	0	0	0	0
BBB-	10	1	0	0	1	0
BB+	11	2	0	0	2	0
BB	12	3	0	0	3	0
BB-	13	4	0	0	4	0
B+	14	5	0	0	5	0
B	15	6	0	0	6	0
B-	16	7	0	0	7	0
CCC	17	8	0	0	8	0

Significant coefficients of the same sign on both *CACDISTGOOD* and *CACDISTBAD* mirror a non-monotonic effect of CACs depending on ratings: it would support the occurrence of a U-shaped (inverse U-shaped) if the coefficient is positive (negative).

In the quadratic specification, *CACS_VARIABLES* are:

- CAC_i , as above;

- $CACRTG_{i,t}$ and $CACRTG\_SQ_{i,t}$: the interaction between CAC_i and the issuer's rating at time t and the interaction between CAC_i and the square of issuer's rating.

A positive (negative) significant coefficient on $CACRTG\_SQ_{i,t}$ would suggest a convex (concave) relations between CAC and rating, possibly non-monotonic if $CACRTG_{i,t}$ has the opposite sign.

We view these two specifications as complementary as one can mitigate the drawbacks of the other: on the one hand, the linear specification produces more intuitive estimates as the sign on the coefficients of interest is already informative on the shape of the net effect, and it does not overload the weight at the extreme of the rating scale; on the other hand, the linear form is centred by construction on the mid-rating, thus being more rigid, and it features a kink, while the quadratic form is more flexible and smoother. None of the specifications constraints the effects of CAC to be symmetric for well- and badly-rated issuers. In the linear case, the symmetry between the two effects can be easily tested through a standard post-estimation parametric test.

The use of a panel data model allows us to exploit techniques precluded to cross-section analysis. In particular, the presence of unit-specific unobserved characteristics, which may cause wrong inference or inconsistent estimation in a cross-section framework, can be tested and controlled for.[12]

More in detail, we perform a fixed-effect (FE) estimation for each and every time sub-set, in order to detect evidence of unit-specific effects and obtain robust estimates against correlations between these effects and the regressors. The FE approach transforms the original data by taking the so-called "within variation", *i.e.*, the deviation from the individual mean taken over time. By doing so, the FE estimator involves a huge loss of efficiency, since as many unit factors have to be estimated as there are units in the sample. Even more importantly, the FE technique prevents any inference on time-invariant regressors (thus including CAC_i), because all the "between variation" is cancelled out. A straight way to estimate these regressors would be adopting a random effect (RE) technique, which is also more efficient than the FE one. In the RE estimator, the unit-specific effects are dealt with as if they were drawn from a random distribution. However, the RE estimator is consistent only as long as the unit-specific factors and the regressors are exogenous; this condition can be tested in each sub-set through a Hausman-type test (Hausman, 1978).[13] When the Hausman statistics does not bring evidence against the RE estimator, the Breusch-Pagan test

[12] See Baltagi (2008) and Wooldridge (2001) for a comprehensive presentation of panel data techniques.

[13] The test compares the time-variant coefficients estimated through random-effects with those obtained from the fixed-effects estimator. Under the null hypothesis both estimators are consistent, the RE one being preferred, if the null is not rejected, for being more efficient, otherwise, besides the FE, other types of estimators have been developed to identify coefficients on time-invariant variables, such as Hausman-Taylor (1981) or error component two stage least squares (see Baltagi, 2008). In the few previous works based on a panel data framework, the RE estimator has been taken without reporting the test to support its consistency.

(Breusch Pagan, 1979) allows to discriminate between the RE and the pooled OLS estimators: since the former basically belongs to the GLS-type, a failure to reject the null can be interpreted in favour of the latter, being it the best linear unbiased estimator.

Once the estimator is selected, we proceed to inference and to test the significance of the net effect of collective action clauses at each rating.

4.2. The regression results

Results from the panel regressions in each sub-set are reported in Table 3, for the linear and the quadratic specifications. The quality of fit appears adequate with a R^2 slightly above 0.70 in each sub-sets. In each sub-set and specification, the random effect estimator proves to be preferable, given the failure to reject the null hypothesis in the Hausman tests, combined with the strong rejection of the null hypothesis in the Breusch-Pagan tests. The control variables generally take the expected signs, but not all of them add substantial explanatory power to the model. *RATING* has a positive and significant coefficient, implying that issuers with worse rating (which means a larger value of the variable, according to our scale) have to concede larger yields. Coefficients on the two duration variables suggest that bonds with greater duration are more exposed to interest rate risk and hence have to pay higher yields, however this occurs at decreasing rates because of a convexity effect. The *VIX* index tends to increase yields in market situations characterized by high volatility, while it is important to control also for *BM10y* as it captures overall movements in yields. There is also a positive effect of general credit risk premia (*BBBAAA*), but it is not significant, possibly because already captured by bond-specific credit risk measures such as rating. The yield tends also to rise when a call option is embedded, since the investor wants to hedge the risk that the security is called back before maturity, typically when their price is low, but the significance is not high.[14] The expected sign but no statistical significance is found for *REGION* and *SIZE*.[15]

Let us now focus on the variables related to the use of collective auction clauses. In the linear specification, the coefficient on CAC is negative, strongly significant and with a fairly similar magnitude across all sub-sets. The interaction variables *CACDISTGOOD* and *CACDISTBAD* both show a positive sign in all the sub-set and they are generally at least 95% significant, except for *CACDISTBAD* in the second sub-sample. The statistical test for symmetry between these two

[14] On the other hand, the evidence for the coefficient *PUT* is less regular, but this could reflect the scarcity of this option in the sample bonds (only three instances are present).

[15] The sign is expected because regional governments might have to grant a yield premium when compared to states. Similarly, bonds with lower outstanding might have to concede a liquidity premium. With respect to liquidity, there are other indicators which could be considered in place of the amount issued. A natural candidate in this respect would be the bid-ask spread; however its inclusion among explanatory variables would highly increase endogeneity issues. Since the main objective of the present work is not to identify the effects of liquidity on yields, we keep the amount issued.

variables reveals that the null of equal coefficients cannot be rejected in sub-set I and III, while it can be in sub-set II, but only at 90% significance level.
The net effect on yields due to the use of CACs, *ceteris paribus*, is given by:

$$\beta_1\, CAC + \beta_2\, CACDISTGOOD + \beta_3\, CACDISTBAD$$

Hence, it depends on the position in the rating scale. In the middle of the rating spectrum, β_1 prevails and our estimates predict that bonds endowed with CACs enjoy a discount with respect to bonds which are not. For issuers placed at both ends of the rating scale, this effect is contrasted by the one that moves in opposite direction, related respectively to β_2 and β_3. The net effect along the rating scale and its 95% confidence interval is charted in Figure 2. Since β_2 is larger than β_3, the net effect tends to vanish more quickly toward the well-rated issuers, however only in sub-set II the difference acquires some significance, since worse-rated issuers continue to enjoy some discount from featuring CACs. In sub-set I and III the net impact of collective action clauses at the extremes is not statistically different from zero. In all samples, there is a significant yield discount for bonds of issuers in the middle of the rating scale (loosely speaking, from single A to single B). In the quadratic specification, the coefficients on the interactions between CAC and rating (and its square) are always positive and significant. Their implication in terms of net effects are less easy to read from the point estimate than in the linear case, but their sign suggests that the net impact might be U-shaped. This is confirmed in Figure 2, where the net effect:

$$(\beta_1\, CAC + \beta_2\, CACRTG + \beta_3\, CACRTG\_SQ)$$

is plotted against the rating spectrum for each sub-set.

The net effect of CACs implied by the quadratic specification is similar to the one under the linear specification: in all sub-sets, issuers in the middle of the rating scale enjoy a discount for bonds endowed with CACs; for well-rated issuers the effect is largely non-significant in all sub-samples; bad-rated issuers may enjoy some discount but to a lesser extent than mid-rated issuers, moreover this discount does not appear always clearly significant and in fact it is often not. Both specifications seem to suggest that in the second sub-set there is more asymmetry, with the net effect for the bad-rated being in between the well-rated and the mid-rated issuer's ones.

Finally, both the linear and the quadratic specifications show wider confidence intervals for bad-rated issuers, so that the net impact of CACs is harder to predict with accuracy.[16]

[16] This may partly be due to a greater heterogeneity, suggesting investigation of the determinants of this variability to further research.

Table 3

Benchmark panel model regressions

(dep. var.: y_{it}, log mid-yield)

	Linear specification			Quadratic specification		
	Sub-set I	Sub-set II	Sub-set III	Sub-set I	Sub-set II	Sub-set III
CAC	–0.427 ***	–0.416 ***	–0.395 ***	0.156	0.170 *	0.173 **
CACDISTGOOD	0.059 ***	0.061 ***	0.058 ***			
CACDISTBAD	0.038 **	0.019	0.038 **			
CACRTG				–0.106 ***	–0.098 ***	–0.100 ***
CACRTG_SQ				0.005 ***	0.004 ***	0.005 ***
RATING	0.092 ***	0.098 ***	0.088 ***	0.092 ***	0.098 ***	0.089 ***
EMTN	0.031	0.014	–0.051	0.032	0.016	–0.049
EURODOLLAR	0.098	0.118	0.107	0.102	0.122	0.113
SEC	0.011	–0.013	0.050	0.007	–0.016	0.046
CALL	0.299 **	0.340 **	0.267 **	0.287 **	0.330 **	0.253 **
PUT	–0.119	0.065	–0.065	–0.124	0.061	–0.075
REGION	0.052	0.082	0.008	0.046	0.081	0.006
SIZE	–0.030 *	–0.026	–0.012	–0.031 *	–0.027	–0.013
DURATION	0.399 ***	0.394 ***	0.302 ***	0.402 ***	0.397 ***	0.303 ***
DURATION_SQ	–0.021 ***	–0.021 ***	–0.015 ***	–0.021 ***	–0.022 ***	–0.016 ***
VIX	0.298 ***	0.356 ***	0.385 ***	0.301 ***	0.358 ***	0.388 ***
BBBAAA	0.038	0.012	0.009	0.038	0.013	0.010
BM10y	0.520 ***	0.573 ***	0.606 ***	0.527 ***	0.579 ***	0.617 ***
R^2	0.72	0.73	0.73	0.72	0.73	0.73
Estimator	RE	RE	RE	RE	RE	RE
Hausman (*p*-val)	0.83	0.44 [c]	0.10	0.18	0.98 [c]	0.43
Breusch Pagan (*p*-val)	0.00	0.00	0.00	0.00	0.00	0.00
Jointly null (*p*-val) ([a])	0.00	0.00	0.00	0.00	0.00	0.00
Symmetry (*p*-val) ([b])	0.36	0.07	0.28			
Groups	285	281	282	285	281	282
Time obs	19	19	12	19	19	12
Avg Obs per group	10.0	10.3	6.3	10.0	10.3	6.3

Stdandard errors robust and clustered by bonds. Legend: *** 99% significant, ** 95% significant, * 90% significant. RE: Random Effects Estimator.

Sub-periods: I: March 2007-September 2008; II: October 2008-April 2010; III: May 2010-April 2011

Constant coefficients not shown. ([a]) In the linear specification: test for coefficients on *CACDISTGOOD* and on *CACDISTBAD* jointly equal to zero; in the quadratic specification: test for *CAC, CACRTG* and *CACRTG_SQ* jointly equal to zero. ([b]) In the linear specification: Test for coefficients on *CACDISTGOOD* equal the one on *CACDISTBAD.* ([c]) The Hausman test could not be computed as the empirical variance covariance matrix is not positive definite. The reported statistics refer to an analogous regression excluding BBBAAA (highly insignificant in the original model), for which the Hausman statistics was computable. The two regressions yield almost identical coefficients and levels of significance. Results are available from the authors.

4.3. Discussion

The intuitions behind this evidence are perhaps not surprising. Collective action clauses are meant to assist ordered debt restructuring when a default event occurs. Hence, they are valuable *ex post*, so the market may well be keen to acknowledge a value to CACs as they help limiting disordered default, holdout risks, prisoners' dilemma outcomes, and delays detrimental both to the debtor and to the majority of creditors. The benefits of majority restructuring through collective action clauses have been modelled, amongst others, by Eichengreen *et al.* (2003), Weinschelbaum and Wynne (2005), Ghosal and Thampanishvong (2007) and Fernandez and Fernandez (2007). Kletzer (2003) shows that CACs improves efficiency in lending and repayment, thus improving welfare relative to unanimous consent clauses.

However, the effectiveness of collective action clauses is subordinated to default being a non-negligible chance. For very well-rated issuers, the probability of default is so small that the market does not really care whether CACs are included, as they are less likely to be helpful; hence, there is basically no reason to acknowledge a discount for them.

Besides that, another condition for collective action clauses being valuable to creditors is that the probability of default is, loosely speaking, exogenous to the inclusion of these clauses. Nevertheless, by making debt-restructuring easier, CACs might make it also more likely, or at least might make this suspicion arise. This is the so-called moral-hazard enhancing (*ex ante*) effect which spurs a demand for a yield premium, in the opposite direction with respect to the ordered-restructuring (*ex post*) effect. Several studies have formalized the moral hazard channel in theoretical models: see, e.g., Weinschelbaum and Wynne (2005) and Ghosal and Thampanishvong (2007), while the issue has been empirically investigated by Esho *et al.* (2004).[19]

Hence, these two contrasting effects can be detected: *ex ante,* the effect of increasing the (perceived) default probability and, *ex post,* the effect of reducing the loss given default. Our empirical investigation shows that the issuer's rating matters for the market in assessing their relative importance. One of the greatest costs of default for debtors willing to maintain access to markets is in terms of reputation and risk of being precluded that access, see, e.g., Eaton and Gersovitz (1981) and Sturzenegger and Zettelmeyer (2006).

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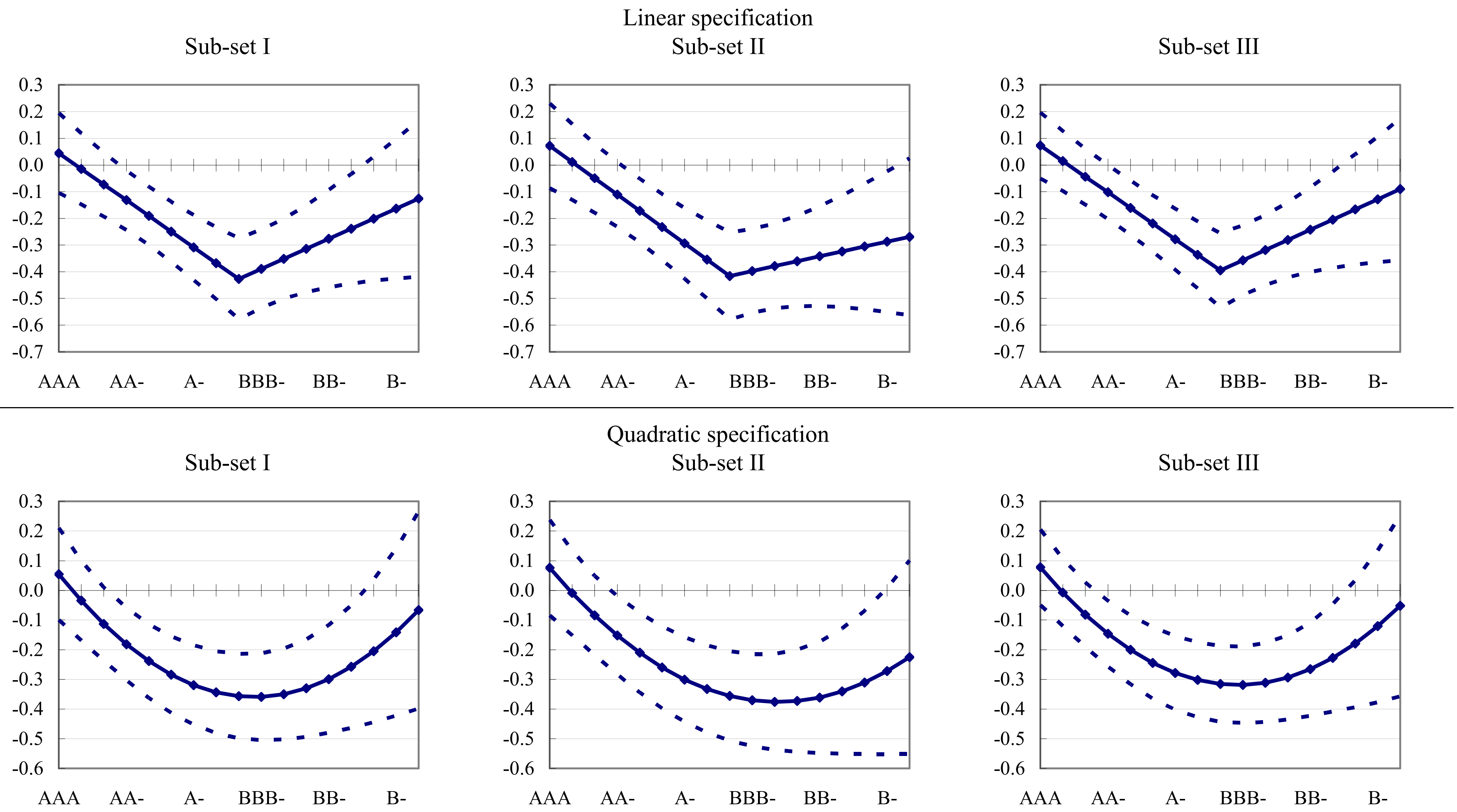
19 The risk that CACs may encourage opportunistic behaviour is claimed also by Dooley (2000) and Cline (2001).

Figure 2

Net impact of CACs by rating and sub-sets

(solid lines: point estimates; dashed lines: 95% confidence intervals)



Sub-sets: I: March 2007-September 2008; II: October 2008-April 2010; III: May 2010-April 2011.

These constraints are much weaker for badly rated debtors, who have a low reputation anyway and are typically less reliant on international bond markets for funding. In this respect, Reinhart and Rogoff (2009) remark that poorly rated countries have usually less access to international bond markets, being their funding sources mainly made up by subsidies and loans from the official sector, and have higher propensity to debt repudiation. Hence it is not surprising to find that market might show higher moral-hazard fears about collective action clauses for bad-rated countries rather than for well rated ones.

The empirical evidence suggests that for well-rated issuers the inclusion of collective action clauses has no substantial impact on ratings because the event of default is so unlikely that any benefit from an ordered restructuring would not matter much. For bad-rated issuers, even if the event of a true default is concrete, the benefit of an *ex post* ordered default is partially counterbalanced by the risk that the default might be declared opportunistically. For sufficiently bad ratings, at least in sub-set I and III, this risk premium appears high enough to almost wipe out the other effect. At medium levels of ratings, the net yield discount is the highest because the chance of a true default is not negligible, while the risk of an opportunistic default is still low given that the higher reputations costs and the fear of foiling its access to the market work as incentives for the issuer to meet its obligations.

In sub-set II, a net yield discount seems to be acknowledged on a wider range of ratings moving toward the end of the rating scale. A possible explanation could be found in the flight away from stock market to bond market in the aftermath of the crisis, when sovereign debt was not yet perceived as a serious issue and government bonds were reputed a safer option. In this situation, the search-for-yield outside of the stock market might have led to invest in more risky countries, while acknowledging at the same time a value to the greater safety promised by CACs.

5. Sensitivity analysis

In this section we perform a number of robustness checks in order to assess the validity of the benchmark model results. In particular, we consider: (i) a different mapping of ratings by grouping them into four classes (Section 5.1); (ii) a dynamic specification for the panel model, allowing for a direct effect of the lagged dependent variable (Section 5.2); (iii) a cross-section analysis repeated by the period (Section 5.3); (iv) an extension to disentangle a possible non-linear effect of ratings not interacted with CACs (Section 5.4); and (v) several other checks in order to address problems of different sample sizes, outliers, etc. (Section 5.5).

5.1. Classes of rating

In the basic specifications, rating values are mapped into a scale starting from 1 for AAA issuers and increasing by 1 for each notch (see Section 4.1). This implies a constant partial impact of rating moving from a notch to the next one. However, there are rating levels which may be seen as more "critical" than others, e.g., the passage from investment grade (at least triple B) to the speculative grade. In order to take that into account, we construct a different measure of credit standing by sub-grouping ratings notches into four classes.[20] However, in so doing we lose information in terms of regressors' variance, because changes in ratings within a certain class are not considered; on the other hand, differences in ratings that might be expected to be more significant for the market are highlighted.

The rating classes (*RTGCLS*) are created as follows: class 1: ratings from AAA to AA-; class 2: ratings from A+ to BBB-; class 3: ratings from BB+ to B+; class 4: ratings lower than B+. In this way, incidentally, investment grade ratings belong to the first two classes, while speculative grade ratings belong to the last two.

The following model is considered:[21]

$$y_{i,t} = \alpha_i + \boldsymbol{\beta}'(CAC_i \otimes RTGCLS_{i,t}) + \boldsymbol{\gamma}'\,\boldsymbol{CONTROL\_VARIABLES_{i,t}} + \eta_i + \varepsilon_{i,t} \qquad (2)$$

Results from random effects estimates are shown in Table 4. As expected, the rating classes coefficients are significantly positive and show a monotonic pattern in all sub-sets. Control variables maintain the same sign and – generally – significance as in the benchmark model. As for the CACs' effect, the first two rating-classes feature coefficients with a negative sign but never significant. The third class of ratings shows the largest coefficient in absolute values; it is negative and highly significant in all sub-sets. The fourth class shows a negative coefficient as well, but it is smaller in absolute terms and less statistically significant: actually, it is significant at 95% only in sub-set III and not even at 90% in sub-set I.

With respect to model (1), model (2) displays a higher degree of collinearity among regressors, a fact which tends to produce less efficient estimates, thus lowering significance levels. Despite that, it seems that the main implication from the benchmark specifications can be retained (see Figure 3):[22] for rather good ratings, the yield effect due to collective action clauses seems to be negligible; there is instead a discount for worse ratings, which, however, does not occur monotonically, as it is smaller (and possibly not highly significant) for worst-rated issuers.

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[20] We thank an anonymous referee for this suggestion.

[21] The *CONTROL_VARIABLES* vector is the same as in the benchmark model, except, of course, for the exclusion of *RATING*.

[22] The net impact of collective action clauses and its significance is given by the straight test on the *CAC* coefficient for rating class 1 and by the test for the null *CAC* + *CACRTGCLS2* = 0 for class 2, 3 and 4.

Table 4

Panel regressions: rating classes specification

(dep. variable: y_{it} log of mid-yield)

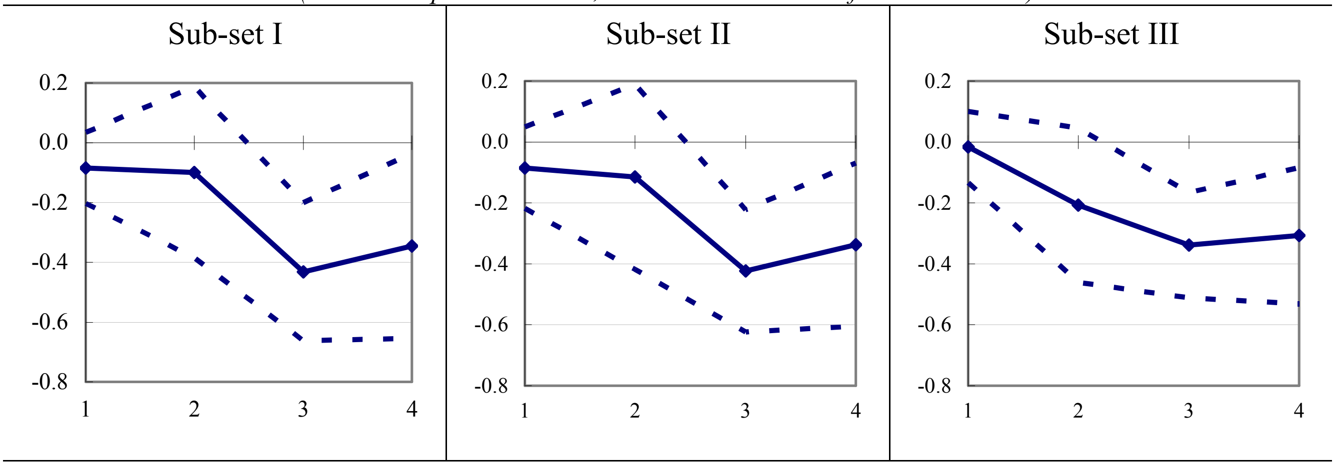
	Sub-set **I**		Sub-set **II**		Sub-set **III**	
CAC	–0.084		–0.084		–0.016	
CAC*RTGCLS2	–0.015		–0.030		–0.191	
CAC*RTGCLS3	–0.348	***	–0.339	***	–0.322	***
CAC*RTGCLS4	–0.261		–0.253	*	–0.291	**
RTGCLS2	0.343	**	0.383	**	0.540	***
RTGCLS3	0.904	***	0.905	***	0.896	***
RTGCLS4	1.162	***	1.192	***	1.125	***
EMTN	–0.050		–0.075		–0.074	
EURODOLLAR	0.116		0.133		0.145	
SEC	0.047		0.046		0.090	
CALL	0.258	*	0.299	**	0.196	**
PUT	–0.135		0.038		–0.095	
REGION	0.002		0.034		0.007	
SIZE	–0.033	*	–0.029		–0.013	
DURATION	0.415	***	0.411	***	0.314	***
SQUARE_DURATION	–0.022	***	–0.023	***	–0.016	***
VIX	0.320	***	0.372	***	0.398	***
BBBAAA	0.018		–0.007		–0.002	
BM10y	0.536	***	0.571	***	0.631	***
R^2	0.68		0.69		0.68	
Estimator	RE		RE		RE	
Hausman (p-val)	n.c.	(c)	0.072		n.c.	(c)
Breusch Pagan (p-val)	0.000		0.000		0.000	
Jointly null (p-val) (a)	0.002		0.001		0.002	
Joint equality (p-val) (b)	0.287		0.212		0.163	
Groups	286		282		283	
Time obs	19		19		12	
Avg Obs per group	10.0		10.2		6.3	

Legend: *** 99% significant, ** 95% significant, * 90% significant. Standard errors robust and clustered by bonds.
Sub-periods: I: March 2007-September 2008; II: October 2008-April 2010; III: May 2010-April 2011.
Constant coefficients not shown. (a) Test for coefficients on *CAC*, *CACRTGCL2, CACRTGCL3 and CACRTGCL4* jointly equal to zero. (b) Test for coefficients on *CAC*, *CACRTGCL2, CACRTGCL3 and CACRTGCL4* jointly equal.
(c) Not computable. The Hausman test could not be computed as the empirical matrix is not positive semi definite.

Figure 3

Net impact of CACs by rating classes and sub-sets

(solid lines: point estimates; dashed lines: 95% confidence intervals)



Rating classes: 1 ∈ [AAA, AA–]; 2 ∈]A+, BBB–]; 3 ∈]BB+, B+]; 4 ∈]B+ ,]
Sub-sets: I: March 2007-September 2008; II: October 2008-April 2010; III: May 2010-April 2011

5.2. Dynamic model

The benchmark model in (1) represents a static panel model. In order to check whether it suffers from an omitted variable problem due to possible significant effects of lagged values of the dependent variables, we turn to a dynamic panel model where $y_{i,t\text{-}1}$ is allowed to have an impact on $y_{i,t}$ as in:

$$y_{i,t} = \alpha_i + \rho\, y_{i,t\text{-}1} + \boldsymbol{\beta' \underline{CACS\_VARIABLES_{i,t}}} + \boldsymbol{\gamma'\, CONTROL\_VARIABLES_{i,t}} + \eta_i + \varepsilon_{i,t} \qquad (3)$$

As a preliminary step, we perform a panel data unit root test *à la* Fisher (1932) for each of the three sub-sets, as described in Maddala and Wu (1999).[23] In all sub-sets the null of unit root is strongly rejected, hence the series do not appear to be integrated and they can be handled as stationary, thus supporting a specification like (3).

In a dynamic panel model, the OLS and the fixed effect estimator are inconsistent and biased respectively upwards and downwards (see Bond, 2002). The inconsistency relies in $E(y_{i,t\text{-}1}, \eta_i) \neq 0$; the within transformation is not able to eliminate this problem, though it can attenuate it for a high T. A well-known approach to solve this issue is the so-called Diff-GMM or Arellano-Bond estimator,[24] which is based on a preliminary transformation of the data to eliminate the unit-specific effect (typically by first-differencing). It then exploits a number of orthogonal conditions between the transformed equation and lagged values of endogenous and/or predetermined regressors.[25] If there are exogenous regressors, they can be used as instruments too. However, the Diff-GMM proves to be scarcely useful in our framework, given that one of the main interesting regressors (*CAC*) is time-invariant; similarly, the fixed-effect estimator too – albeit its bias would be lowered by the high time dimension – cannot provide an estimate for this variable, as it would eliminate it through the within transformation. In situations such as this, the so-called SYS-GMM estimator developed by Blundell and Bond (1998), can be used. Provided some assumptions on initial conditions are met,[26] this technique combines moment conditions on the first difference with the level equations further moments conditions, relating the (lagged) first difference of endogenous or predetermined regressors to the equation in level (possibly in addition to other exogenous

[23] Since the panel is unbalanced because bonds enter and exit the sample according to their issuance and maturity dates, several panel unit root tests are not implementable. The Fisher panel unit root test is instead suitable for unbalanced panel as well, as it based on a combination of p-values from Phillips-Perron unit root tests of cross-section units. See Maddala and Wu (1999).

[24] See Arellano and Bond (1991).

[25] For pre-determined regressors, lags can be used since step one as it orthogonal to the first differenced disturbance: $E(x_{i,t\text{-}1}\, \Delta\varepsilon_{i,t} = 0)$; instead for endogenous regressors (such as the lagged dependent variable), the first lag would be correlated so that only lags since step two can be used as instruments: $E(x_{i,t\text{-}1}\, \Delta\varepsilon_{i,t}) \neq 0$; $E(y_{i,t\text{-}1}\, \Delta\varepsilon_{i,t}) \neq 0$; $E(x_{i,t\text{-}2}\, \Delta\varepsilon_{i,t}) = 0$; $E(y_{i,t\text{-}2}\, \Delta\varepsilon_{i,t}) = 0$.

[26] In particular it is necessary to assume that $E(\Delta y_{i,2}\ \eta_i) = 0$, which amounts to assume a restriction on the behaviour of $y_{i,1}$. Intuitively, it has to be assumed that the same model has generated the y_{it} series for "long enough" prior to the sample period; i.e., for a sufficiently long time to ensure that the true start-up process has become negligibly small. Clearly, this in turns depend on the persistence of the series. See Blundell and Bond (1998).

regressors).[27] With the Sys-GMM estimator, coefficients for time invariant regressors can be estimated as well.

Though the efficiency of the SYS-GMM estimator asymptotically improves with more instruments, in small samples it may perform poorly (Ziliak, 1997). In particular, the exogeneity test on instruments as a whole and on their sub-sets may lose power. For this reason, the use of instruments should be parsimonious and standard errors should be computed through the Windmeijer (2005)'s correction for small samples.[28] We take that into account in the selection of the maximum number of lags for gmm-type instruments and in the computations of standard errors.

Results and econometric tests are shown in Table 5 for the linear specification and in Table 6 for the quadratic one. The Arellano-Bond statistics on the autocorrelation of residuals yields the expected results for a proper panel GMM model, exhibiting correlation at the first order and no correlation at the second order.[29] The robust Hansen statistics for overidentification of restrictions is not rejected both for the whole group and for different sub-sets of instruments.[30] Also, the difference in Hansen statistic for sub-sets of instruments, aimed at detecting whether their specific use as instruments is correct, is never rejected for a standard 95% significance level.

Control variables generally maintain the sign and significance levels; in particular, controlling for rating, duration and its square, market volatility, and the general level of yields is important, as expected. The lagged dependent variable is significant in none of the three sub-sets whatever the specification,[31] thus providing a robustness check for the validity of the static model. Though this might look surprising at first glance, let us recall that, *ceteris paribus*, yields tend to co-move, and we already controlled for the general level of yields, as well as for credit risk and volatility risk factors.

Going into details for CACs-related variables, the following can be observed. In the linear specification (Table 5), the highly significant and negative coefficient for CACs is confirmed, while the interaction variables *CACDISTGOOD* and *CACDISTBAD* are still positive, but only the former

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[27] If a regressor at time t is correlated only with the idiosyncratic disturbance at time t, but not with the unobserved unit-specific heterogeneity η_i, then its lagged value can be used as an instrument in the level equation: $E(x_{i,t-1}(\eta_i + \varepsilon_{i,t})) = 0$. Even if a regressor at time t is correlated with both the idiosyncratic disturbance at time t and with the observed unit-specific heterogeneity η_i, it might be that its lagged first difference it is not, and hence it can be used as an instrument in the level equation: $E(\Delta x_{i,t-1}(\eta_i + \varepsilon_{i,t})) = 0$.

[28] All estimations are performed through the routine xtabond2 implemented for STATA as shown in Roodman (2006).

[29] If the model dynamics is correctly specified, $\varepsilon_{i,t}$ is uncorrelated with $\varepsilon_{i,t-1}$. Therefore, by construction: $E(\Delta\varepsilon_{i,t}\, \Delta\varepsilon_{i,t-1}) = -\varepsilon^2_{i,t-1} \neq 0$.

[30] The Hansen test may be weakened by too many instruments, yielding p-values inflated towards 1. For this reason the number of lags is chosen parsimoniously. p-values on Hansen statistics in Table 5 and Table 6 do not suggest a lack of power problem.

[31] The lagged dependent variable's coefficient tends to rely between the OLS one (upward biased) and the Fixed Effect one (downward biased), considerably more shifted toward the latter as the fixed-effect bias is reduced by the large T dimension.

is significant in all sub-sets. On the other hand, the test for symmetry cannot be rejected in sub-sets I and III, while in sub-set II it can be (but not strongly). These findings are quite similar to the outcome of the benchmark linear (static) specification.

Also, from the quadratic specification (Table 6) we find that it is important to consider interactions between collective action clauses and ratings (coefficients on interacted variables are generally significant), though inferring the net effect of CACs from them is less straightforward. For this purpose, the resulting net effect of collective action clauses across the rating scale is plotted in Figure 4.

In the bottom panel, the U-shaped form of the net impact implied by the quadratic specification resembles the one found in the benchmark (static) version. As in the static models, the second sub-set displays a lower degree of symmetry in both the linear and the quadratic specifications, implying that the discount acknowledged to collective action clauses affected worst-rated issuers more than in the other sub-sets. Hence, this finding too appears rather robust to a dynamic extension of the benchmark model.

Table 5

Dynamic panel regressions: linear distance specification

(dep. var.: y_{it}, log of mid-yield)

	Sub-set **I**	Sub-set **II**	Sub-set **III**
$\boldsymbol{y_{t-1}}$	–0.013	–0.021	–0.013
CAC	–0.379 ***	–0.539 ***	–0.313 **
CACDISTGOOD	0.063 ***	0.103 ***	0.063 **
CACDISTBAD	0.044 *	0.042	0.033
RATING	0.068 ***	0.095 ***	0.083 ***
EMTN	–0.080	–0.216 *	–0.040
EURODOLLAR	0.151	0.052	0.163
SEC	0.087	0.098	0.026
CALL	0.169	0.075	0.182 **
PUT	–0.039	0.041	0.015
REGION	–0.042	–0.026	0.049
SIZE	0.039	0.000	0.019
DURATION	0.304 ***	0.221 ***	0.204 ***
DURATION_SQ	–0.017 ***	–0.011 ***	–0.011 ***
VIX	0.384 ***	0.424 ***	0.479 ***
BBBAAA	–0.087	–0.084	–0.198 **
BM10y	0.549 ***	0.620 ***	0.431 ***
Estimator	SYS-GMM	SYS-GMM	SYS-GMM
Ar-Bond AR(1) (p-val)	0.00	0.00	0.00
Ar-Bond AR(2) (p-val)	0.46	0.31	0.71
Hansen test	0.35	0.52	0.70
Sub-set of instruments			
GMM instruments for levels			
Hansen (p-val)	0.58	0.69	0.28
Difference (p-val)	0.20	0.31	0.94
Gmm (L.ly)			
lags	1-3	1-2	1-3
Hansen (p-val)	0.56	0.49	0.38
Difference (p-val)	0.21	0.52	0.88
Gmm (cacrtg cacrtg2)			
lags	1	1	1
Hansen (p-val)	0.28	0.62	0.64
Difference (p-val)	0.47	0.07	0.81
Gmm (rating4 l10y lvix)			
lags	1	1	1
Hansen (p-val)	0.34	0.98	0.74
Difference (p-val)	0.48	0.13	0.54
IV (cac mtn eurdol sec call put regio) level eq.			
Hansen	0.30	0.49	0.59
Difference	0.76	0.53	0.87
Tests on CACDISTGOOD, CACDISTBAD			
Joint null (p-val)	0.01	0.00	0.11
Symmetry (p-val)	0.51	0.05	0.35
Sample			
Groups	243	241	214
time	19	19	12
avg per group	8.0	8.19	5.7

Two-step standard errors clustered for bonds and robust to small sample dimension are used.

Table 6

Dynamic panel regressions: quadratic specification

(dep. var.: y_{it}, log of mid-yield)

	Sub-set **I**		Sub-set **II**		Sub-set **III**	
$\boldsymbol{y_{t-1}}$	–0.018		–0.011		–0.013	
CAC	0.254	**	0.400	***	0.354	
CACRTG	–0.116	***	–0.158	***	–0.134	**
CACRTG_SQ	0.006	***	0.008	***	0.007	*
RATING	0.085	***	0.089	***	0.082	***
EMTN	–0.077		–0.233	*	–0.063	
EURODOLLAR	0.127		0.044		0.169	
SEC	0.061		0.124		0.016	
CALL	0.129		0.125		0.183	**
PUT	–0.010		0.051		0.029	
REGION	0.076		–0.025		0.010	
SIZE	0.059		0.004		0.019	
DURATION	0.225	***	0.186	***	0.193	***
DURATION_SQ	–0.012	***	–0.009	**	–0.010	***
VIX	0.401	***	0.475	***	0.504	***
BBBAAA	–0.071		–0.130	*	–0.223	**
BM10y	0.660	***	0.648	***	0.432	***
Estimator	SYS-GMM		SYS-GMM		SYS-GMM	
Ar-Bond AR(1) (*p*-val)	0.00		0.00		0.00	
Ar-Bond AR(2) (*p*-val)	0.44		0.35		0.71	
Hansen test	0.32		0.35		0.67	
Sub-set of instruments						
GMM instruments for levels						
Hansen (*p*-val)	0.45		0.78		0.27	
Difference (*p*-val)	0.28		0.11		0.93	
Gmm (L.ly)						
lags	1–2		1–3		1–3	
Hansen (*p*-val)	0.32		0.39		0.46	
Difference (*p*-val)	0.37		0.20		0.78	
Gmm (cacrtg cacrtg2)						
lags	1		1		1	
Hansen (*p*-val)	0.47		0.67		0.66	
Difference (*p*-val)	0.28		0.13		0.45	
Gmm (rating4 l10y lvix)						
lags	1		1		1	
Hansen (*p*-val)	0.81		0.44		0.72	
Difference (*p*-val)	0.05		0.34		0.52	
IV (cac mtn eurdol sec call put regio) level eq.						
Hansen	0.27		0.40		0.63	
Difference	0.74		0.22		0.62	
Tests on CAC, CACRTG, CACRTG_SQ						
Joint null (*p*-val)	0.00		0.00		0.01	
Joint null interactions (*p*-val)	0.00		0.00		0.07	
Sample						
Groups	243		241		214	
time	19		19		12	
avg per group	8.0		8.19		5.7	

Two-step standard errors clustered for bonds and robust to small sample dimension are used.

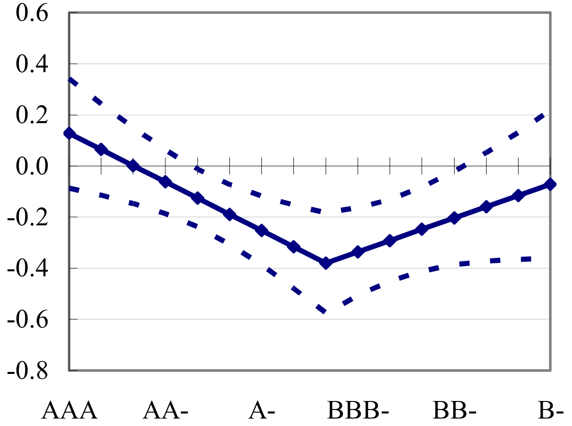
Figure 4

Net impact of CACs by rating and sub-sets

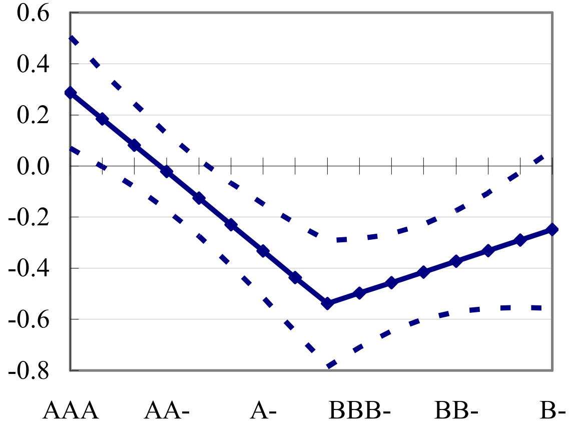
(solid lines: point estimates; dashed lines: 95% confidence intervals)

Linear specification

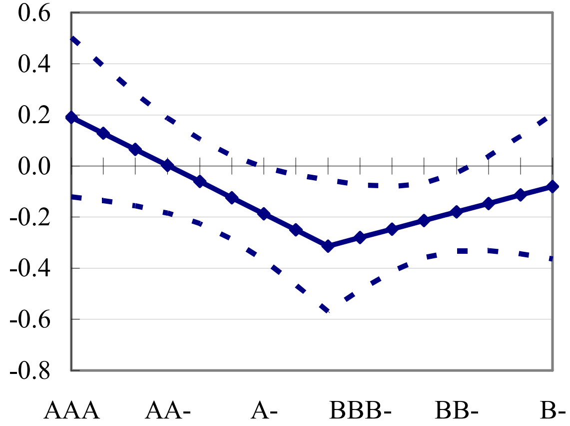
Sub-set I



Sub-set II

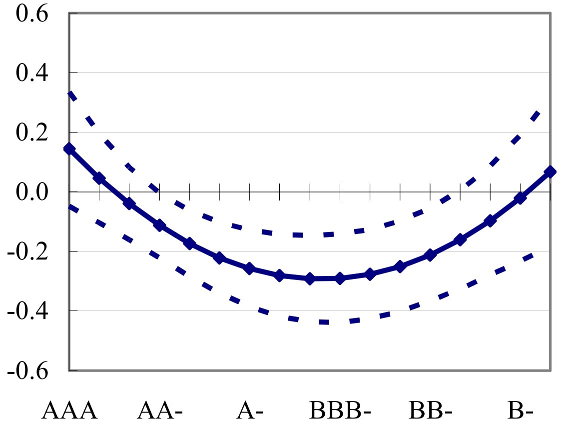


Sub-set III

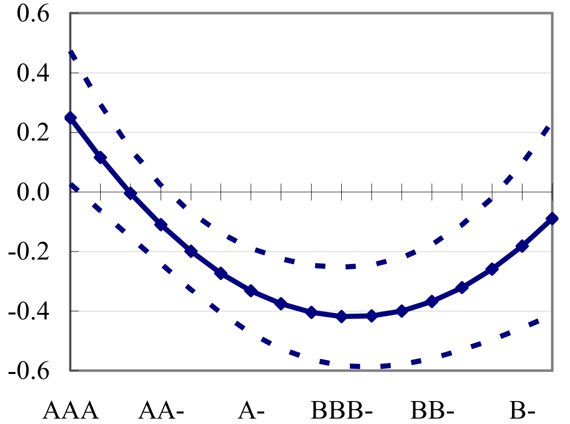


Quadratic specification

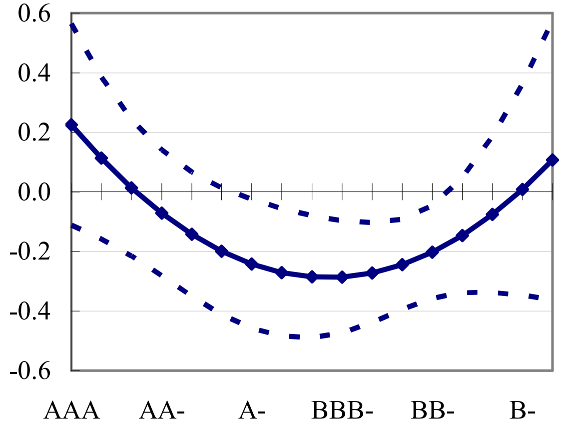
Sub-set I



Sub-set II



Sub-set III



Sub-sets: I: March 2007-September 2008; II: October 2008-April 2010; III: May 2010-April 2011

5.3. Repeated cross-sections

The panel estimation produces synthetic results, encompassing several observations through time. However, data-pooling may hide a different pattern within each sub-set. In order to address this issue, we perform cross-section regressions for each sample period. Note that this approach itself represents a contribution to research as cross-section based regressions on secondary market data up to now have considered only a few dates, thus being subject to the risk of drawing general conclusions out of very particular market situations.

The equation estimated at each time t by OLS with robust standard errors is the same as in (1) but for unit-invariant time-variant regressors (*BM10Y*, *BBBAAA*, and *VIX*), which are excluded because of perfect collinearity.[32] The R^2 coefficients is above 0.70 in every period and we checked that control variables maintain the expected sign when significant. The point estimates and their significance for the variables related to collective action clauses are reported in Table 7 for either specifications. Being generally significant, they confirm the results of the panel analysis. For instance, in the linear specification we can observe the negative significant coefficient of CAC and the positive significant effect on *CACDISTGOOD* and *CACDISTBAD* in almost all of the cross-section regressions (see also Figure 5), which support the hump-shaped effect of collective action clauses (Figure 6). Not surprisingly, the months with lower significance for *CACDISTBAD* and more frequent rejections of the symmetrical effects hypothesis are concentrated in sub-set II, for which the panel analysis also revealed a similar outline. The quadratic specification as well supports in every period the variation in a non-linear way of the effect of collective action clauses with rating (see the significance on *CACRTG* and *CACRTG_SQ*).

When we consider the net effect produced by collective action clauses on yields through time, our results approach the ones from the panel analysis: the effect is negative and higher in absolute terms for issuers approximately in the middle of the rating scale, and non-significant for those at the extremes (see Figure 6).

The graphs equivalent of Figure 5 and Figure 6 for the quadratic specification are shown in the Appendix. The cross-section analysis essentially confirms the results of the panel approach, which remains the preferred one as it produces more parsimonious results, exploits information better, allows for a more proficient control over unobserved heterogeneity, and makes it possible to account for the effects of time series variables (common across units), describing the general market situation, in order to disentangle the genuine CACs' effect.

[32] For the sake of space, the results of each regression are not reported, but they are available from the authors upon request.

Table 7

Repeated cross-sections: coefficients estimates and significance

	Linear specification				Quadratic specification				No.
	CAC	*CACDISTGOOD*	*CACDISTBAD*	Eq.test	*CAC*	*CACRTG*	*CACRTG_SQ*	Jointly	
Mar 2007	–0.090 *	0.016 **	0.015		0.074 ***	–0.035 **	0.002 *	**	85
Apr 2007	–0.065 *	0.014 **	0.003	*	0.071 **	–0.023 *	0.001	*	84
May 2007	–0.072 *	0.016 ***	0.004	*	0.084 ***	–0.027 *	0.001	*	84
Jun 2007	–0.135 ***	0.023 ***	0.018		0.092 ***	–0.043 ***	0.002 **	***	86
Jul 2007	–0.150 ***	0.025 ***	0.021 *		0.099 ***	–0.047 ***	0.003 ***	***	88
Aug 2007	–0.144 ***	0.024 ***	0.014		0.092 ***	–0.043 ***	0.002 **	***	85
Sep 2007	–0.165 ***	0.026 ***	0.023 *		0.102 ***	–0.053 ***	0.003 ***	***	85
Oct 2007	–0.215 ***	0.034 ***	0.033 ***		0.122 ***	–0.060 ***	0.003 ***	***	84
Nov 2007	–0.178 ***	0.030 ***	0.028 **		0.123 ***	–0.056 ***	0.003 ***	***	87
Dec 2007	–0.186 ***	0.031 ***	0.026 **		0.115 ***	–0.053 ***	0.003 ***	***	90
Jan 2008	–0.207 ***	0.033 ***	0.035 ***		0.124 ***	–0.063 ***	0.004 ***	***	92
Feb 2008	–0.236 ***	0.043 ***	0.039 **		0.199 ***	–0.083 ***	0.005 ***	***	93
Mar 2008	–0.299 ***	0.054 ***	0.052 ***		0.233 ***	–0.096 ***	0.005 ***	***	94
Apr 2008	–0.266 ***	0.047 ***	0.045 ***		0.206 ***	–0.087 ***	0.005 ***	***	97
May 2008	–0.345 ***	0.052 ***	0.054 ***		0.179 ***	–0.099 ***	0.006 ***	***	100
Jun 2008	–0.337 ***	0.045 ***	0.051 ***		0.129 ***	–0.096 ***	0.006 ***	***	100
Jul 2008	–0.296 ***	0.040 ***	0.048 ***		0.113 ***	–0.083 ***	0.005 ***	***	99
Aug 2008	–0.369 ***	0.053 ***	0.058 ***		0.167 ***	–0.101 ***	0.006 ***	***	108
Sep 2008	–0.412 ***	0.056 ***	0.065 **		0.160 ***	–0.113 ***	0.007 ***	***	108
Oct 2008	–0.452 ***	0.056 ***	0.079 **		0.127 *	–0.121 ***	0.008 **	***	113
Nov 2008	–0.506 ***	0.060 ***	0.079 *		0.081	–0.105 **	0.006 *	**	104
Dec 2008	–0.390 ***	0.045 **	0.049 **		0.056	–0.083 ***	0.005 ***	***	106
Jan 2009	–0.628 ***	0.087 ***	0.090 ***		0.242 **	–0.160 ***	0.009 ***	***	114
Feb 2009	–0.769 ***	0.095 ***	0.123 ***		0.173	–0.178 ***	0.011 ***	***	119
Mar 2009	–0.680 ***	0.083 ***	0.107 ***		0.151	–0.156 ***	0.010 ***	***	124
Apr 2009	–0.696 ***	0.082 ***	0.075 **		0.076	–0.121 **	0.006 **	**	106
May 2009	–0.715 ***	0.096 ***	0.100 ***		0.216 *	–0.162 ***	0.009 ***	***	132
Jun 2009	–0.786 ***	0.109 ***	0.111 ***		0.279 **	–0.193 ***	0.011 ***	***	134
Jul 2009	–0.587 ***	0.089 ***	0.092 **		0.286 *	–0.158 ***	0.009 **	***	133
Aug 2009	–0.379 ***	0.075 ***	0.037	*	0.356 ***	–0.125 ***	0.006 **	***	184
Sep 2009	–0.410 ***	0.084 ***	0.032	**	0.389 ***	–0.128 ***	0.006 ***	***	188
Oct 2009	–0.480 ***	0.093 ***	0.043	**	0.416 ***	–0.147 ***	0.007 ***	***	187
Nov 2009	–0.417 ***	0.084 ***	0.026	**	0.407 ***	–0.137 ***	0.006 ***	***	203
Dec 2009	–0.492 ***	0.096 ***	0.042	**	0.459 ***	–0.164 ***	0.008 ***	***	207
Jan 2010	–0.379 ***	0.076 ***	0.036 *	**	0.387 ***	–0.135 ***	0.007 ***	***	213
Feb 2010	–0.338 ***	0.071 ***	0.034	*	0.364 ***	–0.121 ***	0.006 ***	***	218
Mar 2010	–0.307 ***	0.063 ***	0.036 *		0.303 ***	–0.104 ***	0.005 ***	***	227
Apr 2010	–0.311 ***	0.062 ***	0.038 *		0.289 ***	–0.102 ***	0.005 ***	***	226
May 2010	–0.370 ***	0.072 ***	0.033	*	0.324 ***	–0.115 ***	0.005 ***	***	224
Jun 2010	–0.437 ***	0.077 ***	0.071 **		0.341 ***	–0.146 ***	0.008 ***	***	221
Jul 2010	–0.487 ***	0.078 ***	0.055 **		0.269 **	–0.131 ***	0.007 ***	***	230
Aug 2010	–0.551 ***	0.087 ***	0.058 **		0.295 **	–0.145 ***	0.007 ***	***	236
Sep 2010	–0.562 ***	0.091 ***	0.055 *		0.326 ***	–0.154 ***	0.008 ***	***	237
Oct 2010	–0.585 ***	0.098 ***	0.065 **		0.370 **	–0.166 ***	0.008 ***	***	241
Nov 2010	–0.590 ***	0.097 ***	0.065 **		0.368 ***	–0.170 ***	0.009 ***	***	241
Dec 2010	–0.520 ***	0.093 ***	0.050 *		0.388 ***	–0.156 ***	0.008 ***	***	245
Jan 2011	–0.480 ***	0.088 ***	0.058 **		0.380 **	–0.150 ***	0.008 ***	***	246
Feb 2011	–0.432 ***	0.088 ***	0.044 *	*	0.419 ***	–0.143 ***	0.007 ***	***	246
Mar 2011	–0.471 ***	0.090 ***	0.049 *		0.390 ***	–0.143 ***	0.007 ***	***	247
Apr 2011	–0.394 ***	0.076 ***	0.037		0.325 **	–0.118 ***	0.005 **	***	249

Legend: * 90% significance, ** 95% significance; *** 99% significance.

Equality test is between *CACDISTGOOD* and *CACDISTBAD*; Joint test is for the joint significance of *CACRTG* and *CACRTG_SQ*.

Figure 5

Coefficients and confidence interval for *CAC* and *CACDISTGOOD/BAD*

(solid lines: coefficient OLS estimates; dashed lines: robust 95% confidence interval)
(left-hand scale: CACDISTGOOD/BAD; right-hand scale: CAC)

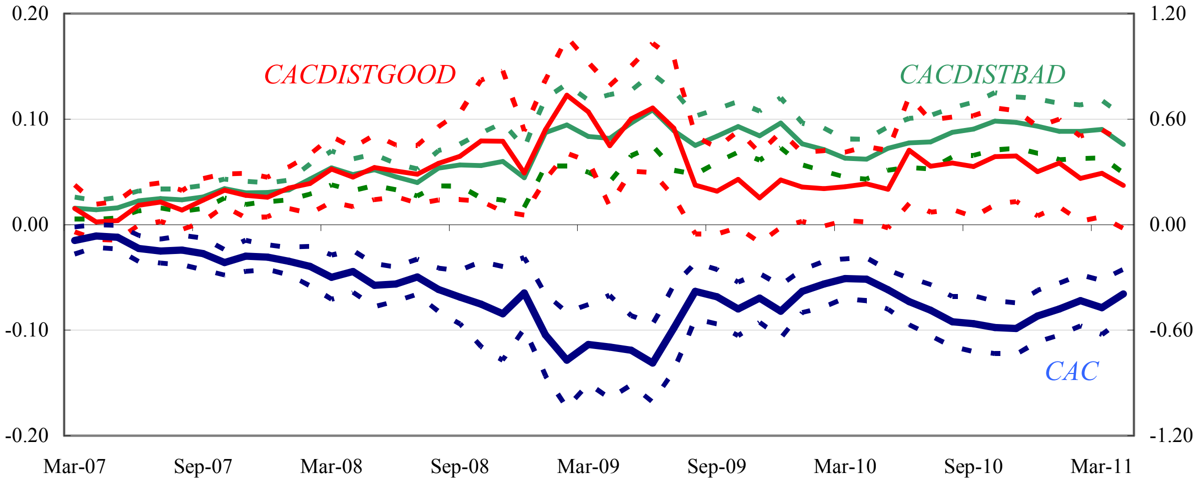
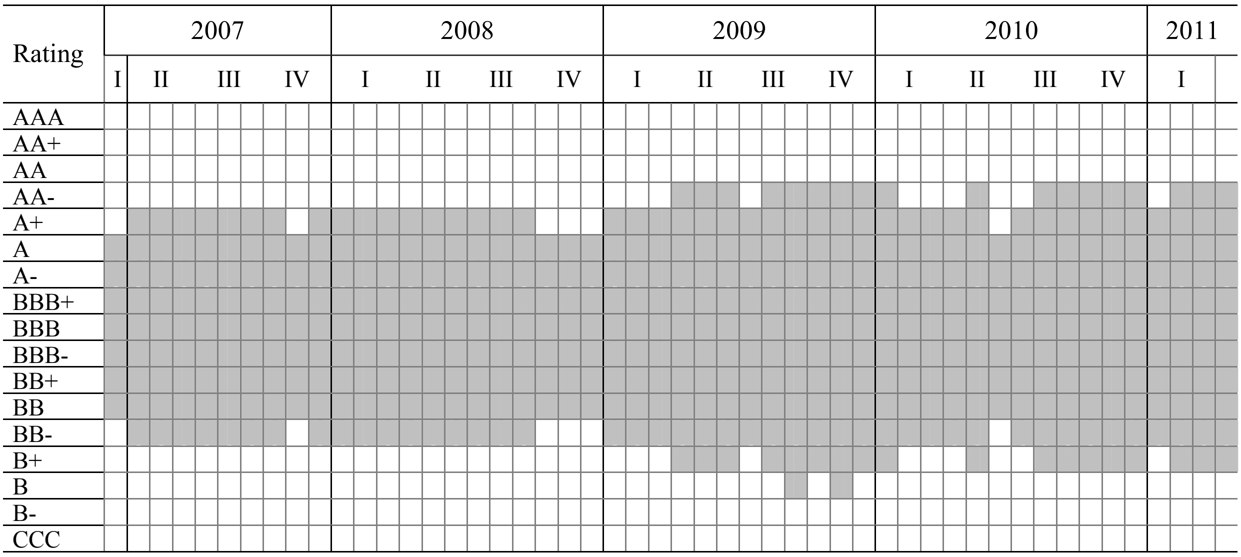


Figure 6

Repeated cross-sections with linear specification
Significance of net effects of CACs through rating and time

(shadowed cells denote 95% significance)



Results from two tails t-tests for the null hypothesis: H_0: β_1 CAC + β_2 $CACDISTGOOD_r$ for r= 0, …,8 for ratings from BBB to AAA; H_0: β_1 CAC + β_3 $CACDISTBAD_r$ = 0 for r= 1, …,8 for ratings from BBB to CCC.

5.4. Controlling for non-linear effects of ratings

In model (1) we have considered the interactions of the CAC dummy with rating measures, *i.e.*, the distance from the centre of the rating spectrum in the linear specification and the rating itself (and its square) in the quadratic specification. However, among the control variables we have considered only the level of ratings and not these other measures (*i.e.*, the rating distances in the linear specifications and the rating squared in the quadratic one) as further control variables. In fact, the very effect of rating on yields might be non-linear and U-shaped, irrespectively from the presence of collective action clauses.

Controlling for that would require adding the following variables: *DISTGOOD* or *DISTBAD* (measuring the distance from the centre of the rating spectrum respectively for better and worse ratings) in the linear specification;[33] *RATING_SQ* (measuring the square of rating) in the quadratic specification. Unfortunately, multicollinearity of regressors would become a serious issue after the inclusion of these regressors, as they are highly dependent by (in fact, constructed from) some of the other explanatory variables. With multicollinearity, standard errors are amplified, thus reducing the significance level and making inferences less clear.

Taking that into account, we nonetheless perform an augmented version of the quadratic specification (including also *RATING_SQ*) as a robustness check to assess whether the new variable shows up to be significant and whether the U-shaped effect of collective action clauses is preserved.[34] As shown in Table 8,[35] the new variable *RATING_SQ* is non-significant in either of the three sub-sets. On the other hand, the significance levels of CAC-related variables are lower too because of multicollinearity, though their sign and point estimate are preserved. However, it is interesting to observe that, in spite of the enlarged standard errors, the test for the interaction variables being jointly equal to zero are rejected in all sub-sets at 99%. Therefore, after this robustness check, the effect of collective action clauses appears to still depend on the rating as in the benchmark model. The net effects displayed in Figure 7 show indeed a picture substantially similar to that in Figure 2, though with larger and more irregular confidence intervals because of higher regressors' collinearity.

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[33] Actually, only one of these variables can be used as regressor because of multicollinearity, unless *RATING* itself is dropped.

[34] We present the results for the quadratic specification because its multicollinearity problems are less prominent than in the linear specification.

[35] See Table A.1 in the Appendix for complete results.

Table 8

Panel regressions: quadratic specification adding *RATING_SQ*

(dep. var.: y_{it}, log of mid-yield)

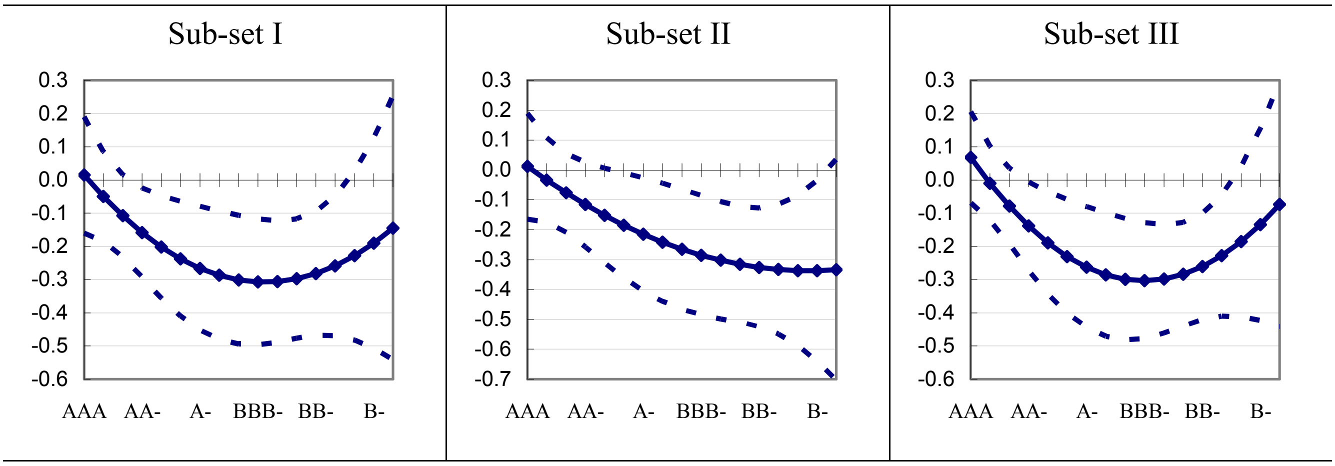
	Sub-set I	Sub-set II	Sub-set III
CAC	0.088	0.062	0.156
CACRTG	0.056	0.042	0.079 **
CACRTG_SQ	–0.076 *	–0.051	–0.092 **
RATING	0.004	0.002	0.005 *
RATING_SQ	0.002	0.003	0.001
Jointly null (p-val) ([a])	0.000	0.011	0.004
Interaction (p-val) ([b])	0.000	0.084	0.007

Legend: *** 99% significant, ** 95% significant, * 90% significant.
Std. Err. adjusted for clusters in bonds.
Sub-periods: I: March 2007-September 2008; II: October 2008-April 2010; III: May 2010-April 2011.
Constant coefficients not shown. For further details see Table A.4 in the Appendix.
([a]) Test for coefficients on *CAC* , *CACRTG, CACRTG_SQ* jointly equal to zero.
([b]) Test for coefficients on *CACRTG, CACRTG_SQ* jointly equal to zero.

Figure 7

Net impact of CACs by rating and sub-sets. Quadratic Specification adding *RATING_SQ*

(solid lines: point estimates; dashed lines: 95% confidence intervals)



Sub-sets: I: March 2007-September 2008; II: October 2008-April 2010; III: May 2010-April 2011

5.5. Different sample sizes, outliers-driven results, exclusion of worst ratings, country effects

The sample size is not constant over time as new bonds are issued and others expire or go out of the sample because their residual life falls below one year (see the last column of Table 7). The changes in the sample size can raise the issue of whether results are due to the presence of CACs of different vintages and recomposition effects so that the aggregate evidence hides a different pattern for CACs of the same vintage. This is also relevant as we have a dummy-type information on collective action clauses, but we do not know whether and how they differ from each other. We address this point by restricting the sample only to bonds available since the first period (March 2007) and then checking whether these subsamples share similar implications with the benchmark model. Cross-section regressions on the restricted sample actually confirm the negative sign for *CAC* and the positive ones for *CACDISTGOOD* and *CACDISTBAD,* though the latter is not always significant. The implied net effect of CACs on yields is negative and significant in the middle of the rating scale, while it tends to vanish at the extremes.[36] As in the benchmark model cross-sections, the equality test on *CACDISTGOOD* and *CACDISTBAD* coefficients can be rejected only in very few cases (5 out of 50).

Another potential issue is whether results may be driven by some outliers. Some bonds appear indeed to be outliers once the effects of explanatory variables are controlled for. Several of them belong to few countries, in particular Venezuela.[37] We tackle this potential problem by adding a Venezuela dummy to the basic model, but the benchmark results are substantially all confirmed, in particular those related to the effects of CACs (see Figure A.5 in the Appendix).[38] Note that adding a country dummy for each and every country is not very useful in our main sample: as shown in Figure 1, plenty of countries with just few issuances make up the sample; so that including a dummy per country would increase collinearity and reduce the degree of freedom, recommending restraint. Finally, we exclude from the sample the worst-rating categories, suspected to be misleading. Again, the estimation on the restricted sample gives results similar to the benchmark's.[39]

6. Conclusions

Collective Action Clauses are contractual provisions included in the issues of sovereign bonds to ensure orderly debt restructuring. The European Council of 24-25 March 2011 decided that standardized and identical CACs would be included in all new euro area government bonds from 2013 onwards. The conclusions of the March meeting state that "the inclusion of CACs in a bond will not imply a higher probability of default or of debt restructuring relating to that bond". On the other side, De Grauwe (2011) provides some evidence of the bond spread increases following the first proposal to introduce CACs on the euro area stage.

The impact of CACs on borrowing costs is an empirical issue over which research has not yet found a sound consensus. In this work we update previous studies encompassing the financial and

36 Results, not reported here for space limitation, are available from the authors upon request. With respect to the benchmark results, we observe a slightly larger region of significant negative impacts on yields in the middle of the time interval, and a slightly smaller one in the last part of the time interval.

37 Venezuela is the leading country for number of defaults in the modern era: 10 since its independence in 1830 (Reinhart and Rogoff, 2009).

38 For the sake of space only graphs for the net effects are reported. Estimation are available from the authors upon request. The inclusion of Venezuela dummy increases the R^2 coefficient by about 0.04.

39 For the sake of space estimation results are not displayed but are available from the authors upon request.

the sovereign debt crisis. Moreover, we contribute to the literature by a more homogenous sample and a more accurate bond-specific tracking for the inclusion of CACS.

The empirical analysis suggests that the effect of CACs on yields may vary in a non-linear way according to the issuer's rating. In particular, a U-shaped impact of collective action clauses on yields seems to emerge, with a discount acknowledged to issuers in the middle of the rating scale and no effect for those at the extremes. Our interpretation is that for mid-rated issuers' creditors the advantage from CACs is greater because the probability of default is not negligible (so CACS are actually valuable) and at the same time the debtor is less suspected of opportunistic behaviour. Instead, for very well rated issuers the chance of default is low, thus reducing the value of ordered restructuring, while very badly-rated issuers face lower reputational costs and are more suspected of moral hazard if they choose to include CACs that favour debt restructuring. Noticeably, across all specification and robustness checks, irrespectively of rating, there emerges no evidence that the inclusion of collective action clauses implies an increase, *ceteris paribus,* in sovereign bond yields.

A. Appendix

A.1 Further descriptive analysis

Figure A.1

Sovereign bond yields
(percent)

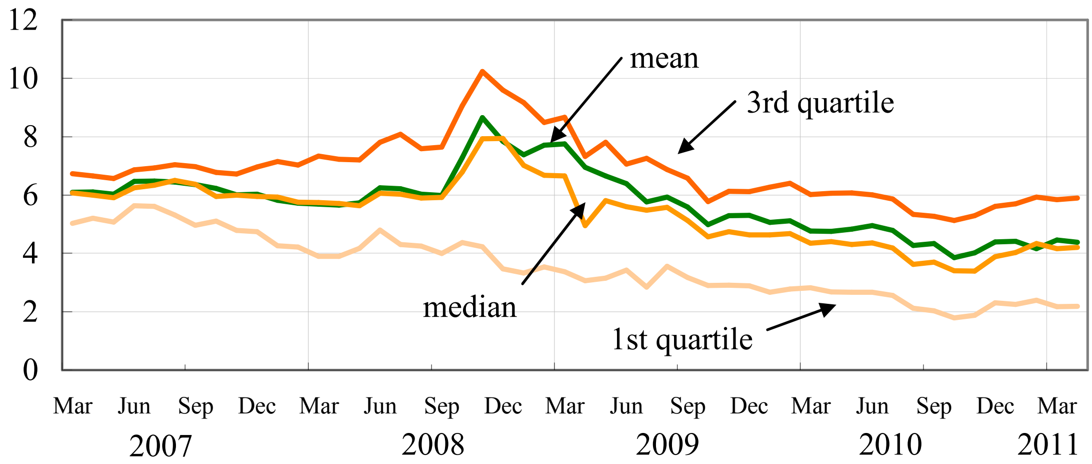
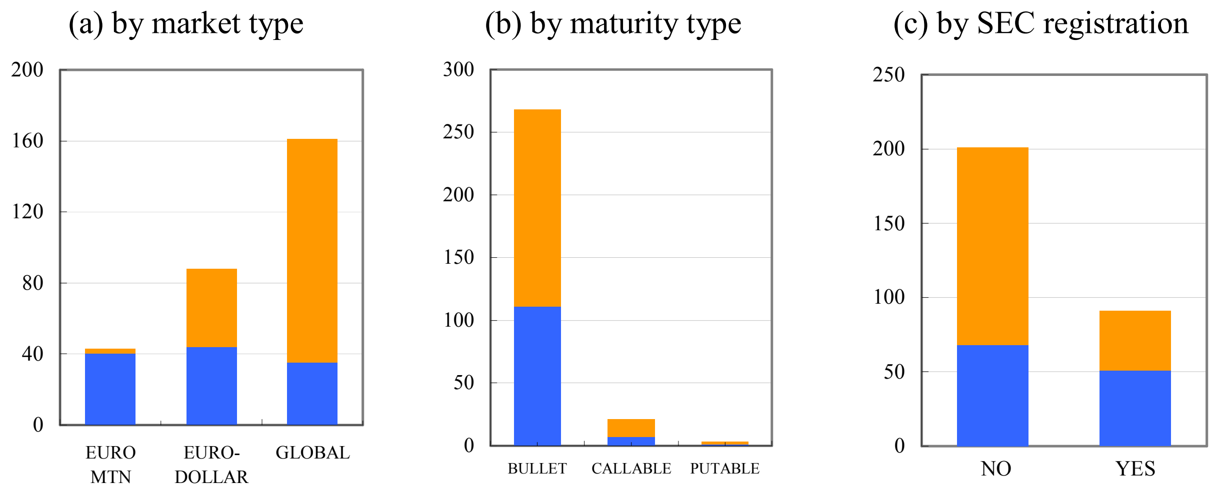


Figure A.2

Distribution of Collective Action Clauses
(number of bonds in the sample)



A.2 Cross section regressions

Figure A.3

Coefficients and confidence interval for *CAC, CACRTG, CACRTG_SQ*

(solid lines: coefficient OLS estimates; dashed lines: robust 95% confidence interval)

(Left-hand scale: CAC, CACRTG; right-hand scale: CACRTG_SQ)

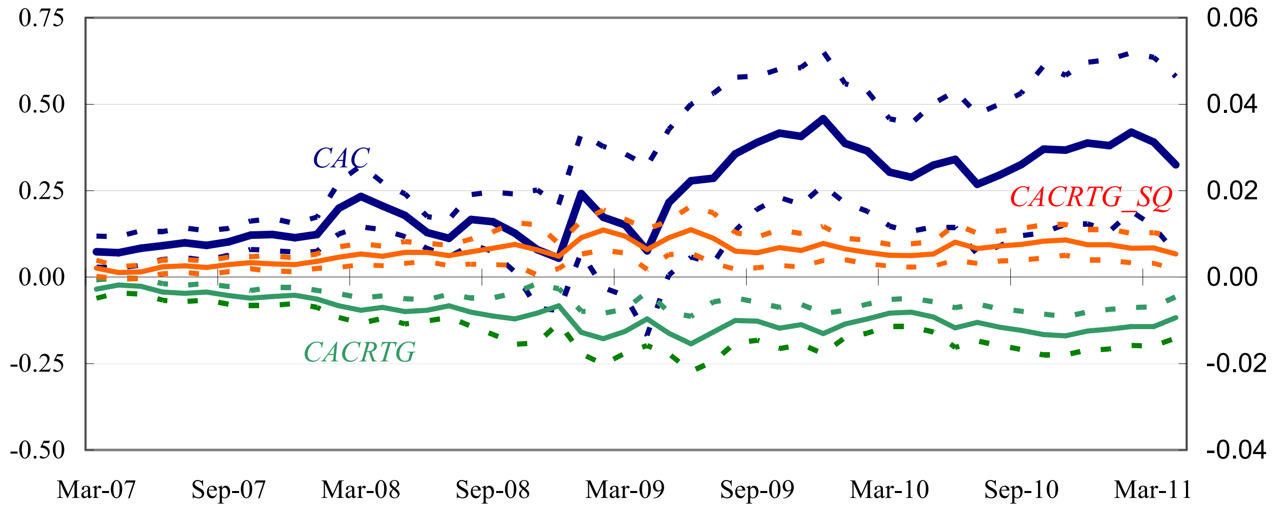
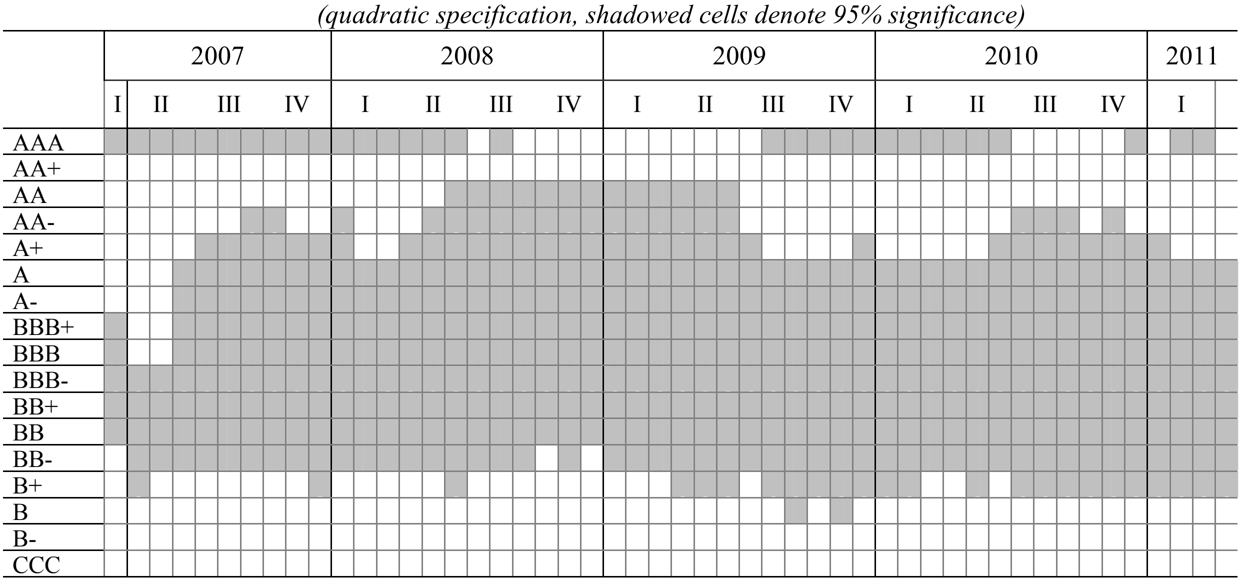


Figure A.4

Significance of net effects of CACs through rating and time

(quadratic specification, shadowed cells denote 95% significance)



Results from two tails *t*-tests for the null hypothesis: $H_0 : \beta_1 + \beta_2 CACDIST_r = 0$ for r= 0, …,8 (the steps of rating distance are 8).

A.3 Further robustness checks on the panel model

Table A.1

Panel regressions: quadratic specification adding *RATING_SQ*

(dep. var.: y_{it}, log of mid-yield)

	Sub-set I	Sub-set III	Sub-set III
CAC	0.088	0.062	0.156
CACRTG	0.056	0.042	0.079 **
CACRTG_SQ	–0.076 *	–0.051	–0.092 **
RATING	0.004	0.002	0.005 *
RATING_SQ	0.002	0.003	0.001
EMTN	–0.014	–0.055	–0.062
EURODOLLAR	0.085	0.096	0.108
SEC	0.023	0.012	0.050
CALL	0.297 **	0.347 **	0.256 **
PUT	–0.108	0.073	–0.072
REGION	0.013	0.031	–0.002
SIZE	–0.033 *	–0.030	–0.013
DURATION	0.403 ***	0.399 ***	0.303 ***
SQUARE_DURATION	–0.021 ***	–0.022 ***	–0.016 ***
VIX	0.299 ***	0.355 ***	0.388 ***
BBBAAA	0.038	0.011	0.010
BM10y	0.523 ***	0.569 ***	0.616 ***
R^2	0.72	0.72	0.72
Estimator	Random effects	Random effects	Random effects
Hausman (*p*-val)	0.18	0.00	0.00
Breusch Pagan (*p*-val)	0.00	0.00	0.00
Jointly null (*p*-val) ([a])	0.00	0.01	0.04
Interaction (*p*-val) ([b])	0.00	0.08	0.07
Groups	285	281	282
Time obs	19	19	12
Avg Obs per group	10.0	10.3	6.3

Legend: *** 99% significant, ** 95% significant, * 90% significant.
Std. Err. adjusted for clusters in bonds.
Sub-periods: I: March 2007-September 2008; II: October 2008-April 2010; III: May 2010-April 2011.
Constant coefficients not shown.
([a]) Test for coefficients on *CAC* , *CACRTG, CACRTG_SQ* jointly equal to zero.
([b]) Test for coefficients on *CACRTG, CACRTG_SQ* jointly equal to zero.

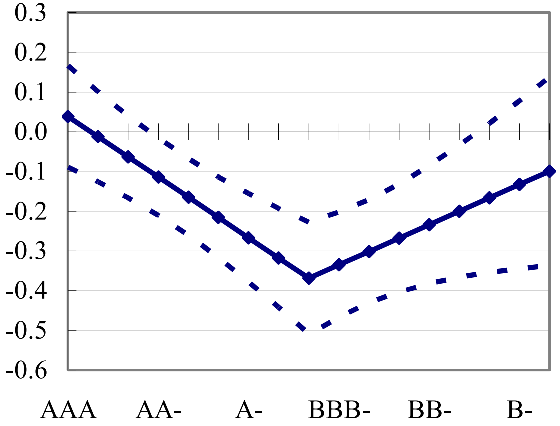
Figure A.5

Net impact of CACs by rating and sub-sets when Venezuela dummy is added

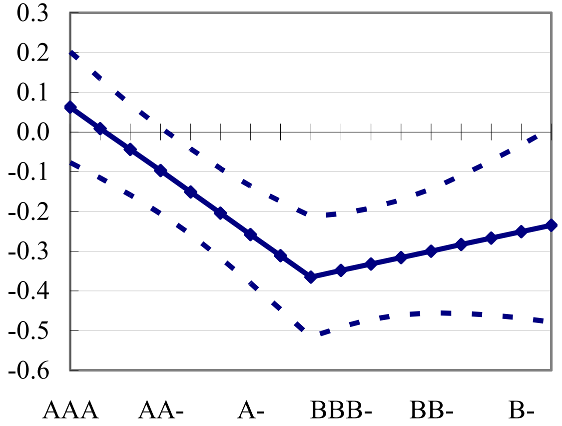
(solid lines: point estimates; dashed lines: 95% confidence intervals)

Linear specification

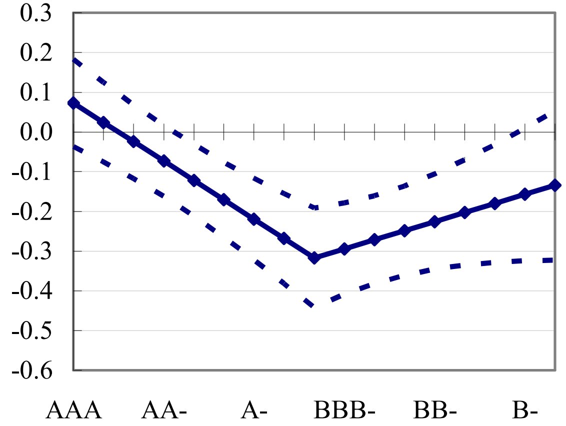
Sub-set I



Sub-set II

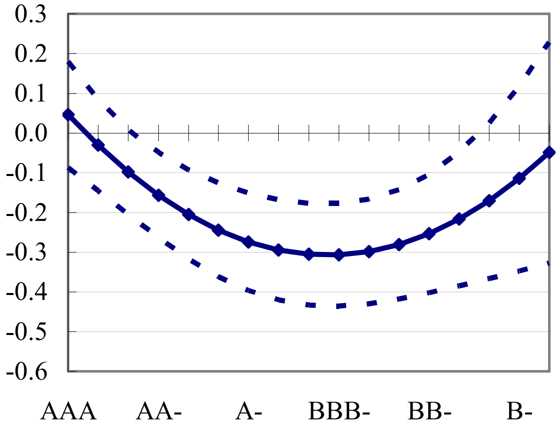


Sub-set III

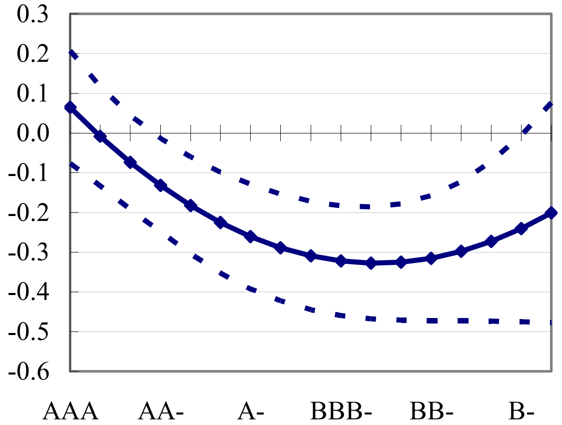


Quadratic specification

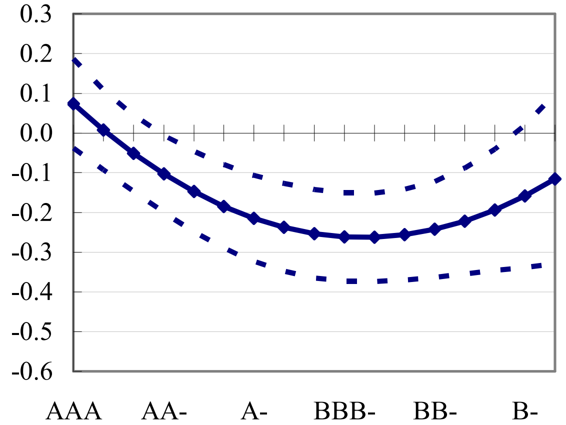
Sub-set I



Sub-set II



Sub-set III



Sub-sets: I: March 2007-September 2008; II: October 2008-April 2010; III: May 2010-April 2011.

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