Aalto University
School of Business

# The relationship between credit default swap spreads and credit rating announcements: Empirical evidence from the emerging Asian markets

*Bachelor's Thesis*

*Finance*

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**Abstract**

This thesis analyzes the informational content of credit rating announcements made by Moody's and S&P between October 2007 and March 2013, with a focus on the rating announcement reaction in the Asian credit default swap markets. To analyze whether rating events cause abnormal CDS spread changes, I apply traditional event study methods and pooled cross-sectional regression analysis. First, I find evidence that both upgrades and downgrades bring new information to the markets on announcement. Second, I observe the markets to anticipate upgrades 30 days before and downgrades 90 days before the rating change. Third, my results suggest that events occurring around the investment grade threshold generate economically largest abnormal spread changes.



*Keywords*: Credit default swaps; Credit ratings; Market reaction; Event study


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## 1. Introduction

Credit rating agencies are important in providing financial information to the market participants, and their opinions on sovereign and corporate entities' credit quality are thought to alleviate information asymmetries in the markets when it comes to the creditworthiness of an entity. The rating opinions impact stakeholders in the markets in many ways; for example, the rated entity's interest rate on outstanding debt may be conditional on its current credit rating, and should the entity's rating downgrade, its debt capital would become more expensive. On the other hand, institutional investors may have regulation limiting their investment only to entities that are rated above a certain threshold grade (i.e. the lowest investment grade rating Baa3/BBB-) in order to control for the amount of risk they are willing to take. Even though credit ratings are monitored carefully by a number of market participants for implications of a change in credit quality, the informational content of the ratings has been discussed intensely. As the importance of the rating agencies has grown immensely in the past few decades, the accuracy and timeliness of credit ratings has been under scrutiny, especially around significant global events, i.e. 1997 Asian financial crisis and 2008 global financial crisis.

The study on the informational content of credit rating announcements initially focused on reactions to rating changes in the bond markets (e.g. Katz, 1974; Weinstein, 1977), followed by studies analyzing the similar reaction in the stock market (e.g. Holthausen and Letfwich, 1986; Goh and Ederington, 1993). However, a new derivative contract, credit default swap (CDS)[1], was introduced in the mid-1990s and the more recent studies (e.g. Hull et al., 2004; Norden and Weber, 2004) have focused on the rating announcement effect in the CDS markets. CDS spread, also the annual payment of the swap, contains information about the default risk of an entity, and thereby the CDS spreads and credit ratings should reflect the same information.

As all of the three major credit rating agencies originate from the United States, many of the studies on rating announcement effects in general have been conducted using data of US reference entities. Also, past studies focusing on the relationship between CDS spreads and rating announcements have conducted their studies with worldwide data, where European and North American entities represent a great share of the total number of sovereigns and corporates in the sample.[2] Thereby, the

[1] CDS offers protection to its buyer in case the reference entity of the contract goes into default, and in return the seller receives annual CDS spreads from the buyer until default of the reference entity or maturity of the contract.
[2] For example, Hull et al. (2004) and Norden and Weber (2004) use a sample, where only 22% and 9% of the entities, respectively, are not from North America or Europe.

past studies analyzing the CDS market reaction to rating events do not draw conclusions on continent-specific effects. Asia as a continent has increased its economic power significantly since the late-1990s and many of the Asian countries have improved their financial markets through deregulation. Thus, I find it relevant to study the effect of credit rating announcements solely in the Asian CDS markets. To my best knowledge, this thesis is the only study analyzing the announcement effect exclusively in the Asian credit default swap markets, and along with e.g. Li et al. (2006), who study the stock market reaction to rating events in the Japanese markets, one of the few studies analyzing the announcement effect of rating changes in the Asian markets in general.

This thesis provides an in-depth analysis of the Asian CDS markets' reaction to different credit rating announcements made by Moody's or S&P. My final sample covers 373 rating events of 240 reference entities from 15 Asian markets. I study the market reaction to four rating events; upgrades, downgrades, new ratings and rating withdrawals, the first two of which I also divide into two smaller subsamples focusing on rating events within a certain rating grade. My objective is to study the market reaction to the rating events in the event window of 90 days before and after the announcement day zero, to draw conclusions on their informational content in the Asian markets. Thereby, my first research question aims to study whether rating announcements bring new information to the markets and thereby cause significant changes in the CDS spreads immediately on announcement. Secondly, I also aim to find out whether the Asian markets view the informational content of Moody's and S&P ratings differently from one agency to another. My analyses yield results, which suggest that the Asian CDS markets react to both upgrades and downgrades on announcement, but a pending rating change is also anticipated as early as 90 days before a downgrade and 30 days before an upgrade. The observed announcement effect to upgrades is similar across agencies, but the markets do not react significantly to a Moody's downgrade.

This study is structured in the following way; in Section 2, I will introduce the existing literature on market reactions to credit rating announcements, as well as state my hypotheses and discuss the theory base behind them. The third section will detail my CDS spread and credit rating data sets and present my final uncontaminated sample of rating events. The fourth section will specify the methodology I use and present the univariate and multivariate analyses that are extended, and brought to the Asian context, from the methods used by Norden and Weber (2004) and Galil and Soffer (2011). Section 5 summarizes my results from the univariate and multivariate tests and Section 6 extends to discuss the results in greater detail by contrasting them with my hypotheses and with past results. Section 7 concludes my thesis and offers topics to consider for further study.

## 2. Literature review and hypotheses

When assuming the semi-strong-form of market efficiency, if a rating announcement brings new information to the markets, the prices should react to the rating change timely on announcement. However, if credit ratings are only based on publicly available information, a rating announcement would not cause an effect in the markets at all. The question regarding the informational content of credit ratings has puzzled researchers for decades, largely due to the strong presence of credit rating agencies in the global financial markets. Loeffler (2005) studied the timeliness of credit rating announcements and concludes that the rating agencies do not react to public information immediately, since they have policies stating that ratings may only be changed if a reversal of the rating is highly unlikely in the nearby future. This would imply that the market prices have already adjusted to the information before the rating is actually changed.

### 2.1 Bond and stock market response

The studies focusing on analyzing the informational value of credit rating announcements first evolved in the 1970s and tested for market efficiency in the US bond markets (e.g. Katz, 1974; Weinstein, 1977). The early bond market studies do not find evidence of a significant announcement effect and they also note that markets do not anticipate future rating events either. However, a more recent study by Steiner and Heinke (2001) on the German Eurobond market concludes that significant abnormal returns are found on Moody's and S&P downgrade announcements. They also note that markets overreact to the negative announcements and positive abnormal returns follow the initially negative ones three weeks after the announcement. Interestingly, Steiner and Heinke (2001) also present that the bond prices of US-based issuers react stronger to the rating changes than issuers of other nationality.

Several studies have extended the analysis on the informational content of credit ratings to the stock market as well. Holthausen and Leftwich (1986) study the effect of positive and negative rating announcements in the US stock market, and conclude that, although upgrades do not generate abnormal returns during the announcement day, downgrades by both agencies and S&P watch list announcements result in abnormal returns on announcement. A later study on stock market reaction (Goh and Ederington, 1993) divides downgrades into two samples. The division is based on whether the downgrade is a result of new negative information or the rated entity's increased leverage. They find that only downgrades that contain non-public information, e.g worsening prospects of the company's financials, cause significant abnormal returns during on announcement in the US stock market and no abnormalities can be observed around downgrades reflecting already public information, e.g changes in the entity's capital structure. Contrary to majority of the literature, Elayan et al. (2003) conducted their study focusing solely on the market efficiency in the New Zealand stock market. They conclude that US based credit rating agencies provide highly valuable information to the global markets about a small country's entities and as a result, both positive and negative rating actions cause significant market reactions on announcement.

### 2.2 CDS market response

After the introduction of credit defaults swaps in mid-1990s, the scope of the studies on credit rating announcement effects has widened to analyze the CDS market as well. Hull et al. (2004) and Norden and Weber (2004) were the first to study the relationship between CDS spreads and credit rating announcements, both using a global sample consisting of mainly North American and European companies. Hull et al. (2004) consider only Moody's ratings and conclude that the CDS markets do not react to Moody's downgrades on announcement, but show significant anticipation of all negative rating events starting as early as 90 days before the rating announcement. Norden and Weber (2004) extend their study to all three major rating agencies and note that all negative events are anticipated by the CDS market and that the results are similar across all the rating agencies. Interestingly, they also conclude that the CDS markets react to rating changes earlier than the stock markets do. Neither of the initial papers on CDS market reaction finds evidence that upgrades would cause abnormal spread changes during any event window.

The first CDS market studies provided evidence of the CDS market reaction before and on a downgrade announcement, which a number of following studies also support (e.g. Daniels and Jensen, 2005; Micu et al., 2006). However, when analyzing the American CDS market, Daniels and Jensen (2005) find evidence of a lagging market reaction, since they observed that the CDS spreads continue to adjust to the new information until 15 days past the announcement. Micu et al. (2006) take a different approach in controlling for sample contamination and use a sample with over 8000 rating events.[3] They are the first to note that positive rating events, especially upgrades to Ba/BB, also right below investment grade, result in abnormal CDS spread changes. Micu et al. (2006) also conclude that clustering of similar events does not weaken the significance of a single rating event.

[3] An uncontaminated sample is defined differently across studies. Hull et al. (2004) and Norden and Weber (2004) clear their event sample of all events that are preceded by another event in the previous 90 days. Micu et al. (2006) only remove events if a rating is preceded by another one in the previous 10 days or if two or more ratings occur on a given day.

Galil and Soffer (2011) add to the study of clustered events and analyze the CDS market reaction to credit rating announcements using three samples controlling for different degrees of contamination by other rating events. They conclude that as credit ratings tend to cluster, the use of uncontaminated samples underestimates the overall CDS market reaction. Although Galil and Soffer (2011) find that both upgrades and downgrades have an effect in the CDS market on announcement, they also find clustering of events to weaken the significance of downgrades following any other negative rating action. As the most recent study in the topic, Finnerty et al. (2013) conclude that all positive rating announcements generate abnormal spread changes on announcement, but are not as anticipated as negative rating changes are, suggesting that CDS market participants monitor worsening prospects of entities more than their improvement.

## 2.3 Market response in Asia

A number of studies have tested the informational content of credit ratings in the Asian markets to analyze how much impact the US based rating agencies have in the Asian markets and whether regional agencies are better in assessing the creditworthiness of Asian entities. Li et al. (2006) study the Japanese stock market for abnormal returns around rating announcements of both American and Japanese rating agencies. They find evidence that Japanese markets react stronger to Moody's and S&P downgrades compared those of the Japanese agencies JCR and R&I. Additionally, they note that the market reaction around upgrades of any rating agency remain insignificant and that the reaction to downgrades is similar, regardless of whether it is announced by Moody's or S&P.

Yamori et al. (2006) also compare the foreign and local rating agencies in the Japanese bond market and they find evidence that international agencies tend to assign entities with lower ratings than their Japanese equivalents. Poon and Chan (2008) and Ferri et al. (2013) extend the study to the Chinese and Korean markets, respectfully. Poon and Chan (2008) conclude that, although downgrade announcements by Chinese rating agencies result in abnormal returns in the Chinese stock market, the rating announcements by international agencies have more informational content in the Chinese market. On the contrary, Ferri et al. (2013) find that in the Korean bond market, ratings by local agencies generate higher abnormal returns on rating announcement days, when compared to the announcements by large international rating agencies. Their study suggests that the Korean market participants view local agencies more reputable, which then corresponds to higher informational value of national agencies' ratings.

### 2.4 Hypotheses

Assuming the Asian CDS markets are efficient in the semi-strong-form, the CDS spreads should reflect all publicly available information at any point of time. Thereby, if credit ratings contain both public and non-public information regarding the creditworthiness of an entity, the announcement of a credit rating change would provide new relevant information to the market participants. This emerging information would result in an immediate CDS market reaction towards the expected direction, as the CDS spreads adjust to the new level of default risk. In the case of a rating upgrade, decreasing credit risk should correspond to a decrease in the CDS spread as well, whereas in the case of a downgrade the increasing credit risk of a reference entity should correspond to an increase in the CDS spread. My first hypothesis revolves around the initial market reaction and I expect to observe the following:

*H1 (Informational value hypothesis)*: All credit rating announcements contain new information to which the CDS markets react directly on announcement.

The large international credit rating agencies are conscious about their market reputation and are thereby reluctant to assign ratings unless the need to reverse it in the future is highly unlikely. Conflicting rating actions within a short time period would decrease the agency's reliability and, as a result, the importance of its announcements would lessen and the CDS markets would not react significantly to its announcements. Given the large global presence of Moody's and S&P, I do not find evidence why the markets would view the ratings of either agency less reliable and thereby I expect to observe:

*H2 (Realiability hypothesis)*: CDS market reaction to rating announcements does not depend on the announcing rating agency.

## 3. Description of the data set

### 3.1. The CDS data set

My credit default swap spread data consists of daily CDS spreads collected from Thomson Reuters Datastream. The set includes CDS spreads of all Asian reference entities that are available on the portal and covers the time period from October 1, 2003 to March 18, 2014. The number of reference entities in the raw sample totals to 816 and they range from corporates and financials to sovereigns and quasi-sovereigns. The data set contains a total of over 1,040,000 individual spread quotes and it includes entities from 21 Asian countries.

The quotes are the daily mid spreads of senior CDS contracts with a 5-year maturity. I focus my study on 5-year contracts for two reasons; first, the 5-year CDS is considered the most liquid of credit default swaps and thereby it is thought to be the benchmark maturity on the market (Bomfim, 2005). Second, past studies of Hull et al. (2004) and Norden and Weber (2004) both focus on 5-year quotes and using contracts of the same maturity allows for better comparability between the results. Even though the 5-year CDS is considered the most liquid CDS, some illiquidity still exist in the markets and as a result, some entities are missing CDS quotes on a small number of days. To complete the time series, I use a similar approach to Micu et al. (2006) and hold the most recent price constant until new information arrives in the markets and results in a new quoted price. [4]

The availability of credit rating data made me eventually limit the number of reference entities to 240 and limit the timespan to cover the period from October 1, 2007 to March 18, 2014. All entities without an available credit rating at some point during the period were removed from the data set, resulting in a final CDS spread sample consisting of corporates and financials and 399,819 daily CDS quotes.[5] After leaving all sovereigns, quasi-sovereigns and number of corporates out of the final sample, it covers 15 Asian markets with 55% of the entities originating from Japan. The complete list of reference entities by country of origin can be found in Appendix.

### 3.2. The credit rating data set

My unconditional credit rating data set includes all issuer ratings of the sampled entities covering the time period between October 1, 2007 and March 18, 2014. [6] The data is collected from Thomson Reuters's Thomson ONE and it comprises of ratings announced by Moody's and S&P, also of two of the three leading credit rating agencies. My rating event data consists of the following rating actions: upgrades, downgrades, new ratings and rating withdrawals. The initial unconditional rating event sample includes 508 different rating announcements, of which Moody's rating actions account for 29% and S&P's for 71%.

[4] Hull et al. (2004) and Norden and Weber (2004) do not assume constancy of CDS spreads and instead linearly interpolate the missing quotes.

[5] Previous studies use a final sample of 233,620 (Hull et al., 2004), 60,827 (Norden and Weber, 2004) 1,176,640 (Galil and Soffer, 2011) individual CDS quotes.

[6] Moody's ratings in the sample are long term issuer ratings and the S&P ratings are long-term foreign issuer ratings. Issuer ratings are used to properly account for entity-specific default risk instead of issue-specific risk.

To control for the contamination of the results by extreme outlying quotes, similarly to Galil and Soffer (2011), I disregard all ratings below Moody's rating B3 (S&P rating B-) from my sample. Hereby, I allow ratings events within non-investment grade in my final sample as well, whereas Hull et al. (2004) and Norden and Weber (2004) only study the rating changes within investment grade. However, to allow for comparability of the results, I also analyze subsamples of upgrades and downgrades taking place only within investment grade.

As rating events usually tend to cluster, I need to control for contamination of the results by overlapping events windows. Contamination of an event may result from all other rating events, both within and across agencies, in the same time interval. As my aim is to study the rating effect in the [-90, 90] –time window, I have constructed my uncontaminated rating event sample, presented in Table 1, by excluding all rating changes that are preceded by another rating event in the previous 90 days of the rating announcement day.[7] Hence, the maximum overlap between two rating events in the final sample is 90 days. After the controlling measures, my final uncontaminated rating sample comprises of 373 events, which includes 98 upgrades and 152 downgrades. Moody's rating events amount to 31% and S&P's events to 69% of the total events. Most of the upgrades and downgrades occur within investment grade, which I also study using separate subsamples (UIG and DIG). I also construct two separate samples for the exploratory events (new ratings and rating withdrawals) and two subsamples for the ratings around the threshold grade (UTIG and DTNIG).

**Table 1. Number of rating events by rating type and credit rating agency – uncontaminated sample**

| Moody's or S&P | Events | Moody's | Events | % of total | S&P | Events | % of total |
|---|---|---|---|---|---|---|---|
| Upgrades | 98 | Upgrades | 19 | 19 % | Upgrades | 79 | 81 % |
| Downgrades | 152 | Downgrades | 61 | 40 % | Downgrades | 91 | 60 % |
| New ratings | 30 | New ratings | 20 | 67 % | New ratings | 10 | 33 % |
| Withdrawn ratings | 93 | Withdrawn ratings | 17 | 18 % | Withdrawn ratings | 76 | 82 % |
| Total | 373 | Total Moody's | 117 | 31 % | Total S&P | 256 | 69 % |
| UIG | 68 | UIG | 17 | 25 % | UIG | 51 | 75 % |
| DIG | 121 | DIG | 55 | 45 % | DIG | 66 | 55 % |
| UTIG | 10 | UTIG | 2 | 20 % | UTIG | 8 | 80 % |
| DTNIG | 13 | DTNIG | 6 | 46 % | DTNIG | 7 | 54 % |
| Total | 212 | Total Moody's | 80 | 38 % | Total S&P | 132 | 62 % |

This table breaks downs the total number of rating events in the final uncontaminated sample by event type and credit rating agency. Sample includes Moody's (S&P) rating events above B3 (B-) from October 1, 2007 to March 18, 2014 with CDS spread data available in the full event window of [-90, 90], also 90 days before and after the announcement day zero. The sample consists of the rating events of 240 reference entities from 15 Asian CDS markets. All rating events: upgrades, downgrades, new ratings and rating withdrawals are included in the sample, as well as smaller subsamples of the first two event types. The subsamples are; UIG: upgrades within investment grade; DIG: downgrades within investment grade; UTIG: upgrades to investment grade and DTNIG: downgrades to non-investment grade.

[7] This approach is similar to Norden and Weber (2004), whereas Micu et al. (2006) only require that there are no other rating events in the past 10 days or on the same day of the announcement.

### 3.3. Limitations of data

Limited availability of credit rating data from the Asian markets made me alter my intended data set. I was unable to consider Moody's credit reviews (S&P's watch listings) and credit outlook announcements in my final rating event sample. Previous studies (Hull et al., 2004; Norden and Weber, 2004; Micu et al., 2006) have had evidence that, along with actual rating changes, other rating events also have a significant effect in the CDS markets, and omitting them from my sample may lead to me observing weaker or insignificant market reactions. However, Norden and Weber (2004) also conducted their study without having credit outlooks in their event sample.

Initially I also planned to include the ratings of Fitch in my credit rating sample, but as their ratings cover only less than 10% of the entities in my final CDS sample, I decided to leave Fitch outside of my study.[8] Since over 50% of the entities in my CDS data are traded in the Japanese market, I also considered adding the ratings of the two major Japanese credit rating agencies, JCR and R&I, in my rating event sample. However, the rating data of JCR and R&I is not importable from Thomson ONE and thereby, I am unable to study whether rating announcements made by local and foreign agencies cause different effects in the CDS markets.

## 4. Methodology

### 4.1. Univariate event study

I conduct my event study analysis in a similar way as Norden and Weber (2004) and Galil and Soffer (2011) did. I begin by defining the rating announcement day as day zero and, since I use an event window of 90 business days before and after the rating announcement day 0, I match the CDS spread data around a rating event with the event time variable taking values from the event time range [-90, 90]. I have divided the event window into nine shorter intervals, which I use to better analyze the CDS market reaction before, on and after the rating event. The shorter event windows, which are also used by Galil and Soffer (2011), include the intervals before the rating event [-90, -61], [-60, -31], [-30, -11] and [-10, -2], the announcement day interval [-1, 1] and the intervals after the rating event [2, 10], [11, 30], [31, 60] and [61, 90].

[8] Ratings by Fitch were also omitted by Hull et al. (2004) and Galil and Soffer (2011). Norden and Weber (2004) and Micu et al. (2006) included rating events of all three agencies, however, the former study received evidence that Fitch ratings and reviews do not have a significant effect on CDS markets. Additionally, Micu et al. (2006) do not analyze their results on an agency level.

Although CDS spreads act as a fair representation of company specific default risk, they are also affected by significant region-wide and global effects and thereby my results may be contaminated by changes in the overall market conditions. To control for this, I construct ten CDS spread indices, where an index corresponds to a mean of daily CDS spreads within a whole letter rating class.[9] For Moody's rating events, I use the whole letter rating indices of Aaa/Aa, A, Baa and Ba/B and for S&P they are AAA/AA, A, BBB and BB/B. To analyze new ratings and rating withdrawals in a similar way, I have also constructed a "Not rated" –index for both Moody's and S&P. Due to the fact that the entities without a rating may theoretically belong to any available rating class, the "Not rated" –index may not be as efficient as the whole letter rating indices are, but it does provide a controlling measure for the changes in the overall market conditions to some extent.

As my CDS spread sample consists of entities from 15 Asian markets, the whole letter rating indices have their shortcomings as well, since they assume significant events to cover all global markets, or at least the regional markets in question, and thereby the indices do not account for possible country-specific events. However, constructing the indices across countries would not be relevant in this study for two reasons; first, a great number country-specific indices would have less than 10 observations in my case. Second, the nationality-based indices would omit the magnitude with which the entities of different rating classes react to the significant events in the CDS markets.

To calculate the abnormal spread change, for each day and entity, I deduct the daily CDS index changes from the raw CDS spread changes:

(1) $$ASC_{i,t} = \begin{cases} (CDS_{i,t} - CDS_{i,t-1}) - (I_{o,t} - I_{o,t-1}), \text{when } t < 0 \\ (CDS_{i,t} - CDS_{i,t-1}) - (I_{n,t} - I_{n,t-1}), \text{when t} \geq 0 \end{cases}$$

where $ASC_{i,t}$ is the abnormal spread change for entity *i* on day *t*

$CDS_{i,t}$ is the raw CDS spread for entity *i* on day *t*

$I_{o,t}$ is the CDS spread index for old rating *o* on day *t*

$I_{n,t}$ is the CDS spread index for new rating *n* on day *t*

To better capture the rating class-specific variance resulting from significant market events, I follow the method used by Norden and Weber (2004) and adjust the raw CDS spread changes with the

[9] I construct the indices based on whole letter rating classes due to an insufficient number of intermediate rating observations, e.g. Moody's A3 or S&P BB+. Previous studies (Norden and Weber, 2004; Micu et al., 2006) also base their indices on whole letter ratings.

index of the old rating class *o* before the rating event, also [-90, -1], and with the index of the new rating class *n* on the announcement day zero and after the rating event, also [0, 90].[10]

In my event study analysis, I use the cumulative abnormal spread changes ($CASC_{i,\tau 1,\tau 2}$) to study the reaction in the CDS markets during the nine event time intervals. I calculate $CASC_{i,\tau 1,\tau 2}$ for each entity and for each of the nine time intervals $[\tau_1, \tau_2]$, where $\tau_1$ is the beginning day and $\tau_2$ is the ending day of the individual event windows, as follows:

(2) $$CASC_{i,\tau 1,\tau 2} = \sum_{t=\tau_1}^{\tau_2} ASC_{i,t}$$

For each event time interval, I also calculate the mean cumulative abnormal spread change, also $\overline{CASC}_{\tau 1,\tau 2}$, on the ending days of each event windows across entities *i*. In my event study, I use the parametric Student's *t*-test and the non-parametric Wilcoxon sign-test. Similarly to Galil and Soffer (2011), to obtain the test statistics to test the significance $\overline{CASC}_{\tau 1,\tau 2}$ during different event windows, I assume cross-sectional independence of $ASC_{i,t}$, meaning that abnormal spread changes do not correlate across entities. I calculate the test statistic for the Student's *t*-test in the following way:

(3) $$t = \overline{CASC}_{\tau 1,\tau 2} / S(ASC_0)$$

where $S(ASC_0)$ is the estimate of the standard deviation of the abnormal spread changes $\sigma(ASC_0)$ during an event window $[\tau_1, \tau_2]$. Assuming cross-sectional independence, the following equation holds:

(4) $$\sigma(ASC_0) = \sqrt{\left(\frac{1}{N^2}\right) * \sum_{i=1}^{N} \sigma^2(ASC_{i,0})}$$

where $\sigma^2(ASC_{i,0})$ is the variance of abnormal spread changes during an event interval $[\tau_1, \tau_2]$, which I estimate by calculating the sample variances of $ASC_{i,t}$ within all the time intervals $[\tau_1, \tau_2]$.

With the event window-specific *t*-statistics calculated as above, I test the mean cumulative abnormal spread change around upgrades and downgrades for significant differences towards their anticipated direction. I study the one-tailed Student's *t*-test against the following null hypotheses:

[10] Hull et al. (2004) and Micu et al. (2006) use a different approach, where they do not change the CDS index on rating event day 0 and instead use the index of the old rating class for the whole event time window. However, the difference in methods only affects upgrades or downgrades that occur from a whole letter rating to another, for example from Baa1 to A3.

Upgrades: $H_0: \overline{CASC}_{\tau1,\tau2} \geq 0$

$H_1: \overline{CASC}_{\tau1,\tau2} < 0$

Downgrades: $H_0: \overline{CASC}_{\tau1,\tau2} \leq 0$

$H_1: \overline{CASC}_{\tau1,\tau2} > 0$

When testing the $\overline{CASC}_{\tau1,\tau2}$ around new ratings or rating withdrawals, I prefer to use the two-tailed Student's *t*-test to test whether these rating announcements cause abnormal changes of any kind in the event window. Contrary to upgrades and downgrades, the anticipated direction to which the abnormal spread changes deviate around new ratings and rating withdrawals not as evident. Thereby, I test $\overline{CASC}_{\tau1,\tau2}$against the following null hypothesis:

New ratings & $H_0: \overline{CASC}_{\tau1,\tau2} = 0$

rating withdrawals $H_1: \overline{CASC}_{\tau1,\tau2} \neq 0$

As Student's *t*-test assumes the population of $\overline{CASC}_{\tau1,\tau2}$to be normally distributed, I also conduct the non-parametric Wilcoxon sign-test to control for abnormal distribution of $ASC_{i,t}$.[11] I calculate the percentage of positive $\overline{CASC}_{\tau1,\tau2}$ across all event windows and rating event types and compare the sign against the evenly the distributed probability of 0.5 between signs. My null hypotheses under the sign-test are as follows:

Upgrades: $H_0: p \geq 0.5$

$H_1: p < 0.5$

Downgrades: $H_0: p \leq 0.5$

$H_1: p > 0$

New ratings & $H_0: p = 0.5$

rating withdrawals $H_1: p \neq 0.5$

### 4.2. Multivariate analysis

To test whether upgrades and downgrades generate a significant reaction in the CDS market, especially close to the actual rating event, I run eight separate multivariate regressions, first testing the combined impact of the two rating agencies, followed by an analysis of the agencies individually. Similarly to Norden and Weber (2004), I analyze the relationship between raw CDS

[11] Test-statistics for the sign-test are obtained through a standard binomial test. Similarly to the Student's *t*-test, I use a one-tail test for upgrades and downgrades and conduct a two-tailed test for new ratings and rating withdrawals.

spread changes and CDS index changes in the pooled cross-sectional regressions. I use the same sampled events as in the univariate analysis, however, due to the equivocal nature of the "Not rated" –index, I decide to leave new ratings and rating withdrawals outside of the multivariate analysis, and thereby the number of events totals to 250. Norden and Weber (2004) considered rating reviews and watch listing in their multivariate analysis sample as well, but due to limitations regarding Asian data, I focus my analysis on upgrades and downgrades only.

First, I study the market reaction around the announcement window [-1, 1] with events of either of the two agencies. As a great proportion of my sampled entities originate from Japan, I also add a control variable to detect possible region-specific (Japanese compared to non-Japanese) differences. Using the Huber/White/sandwich estimator of variance, I estimate a separate model for upgrades (5) and for downgrades (6) in the following way:

(5) $$\Delta CDS_{i,t} = \alpha + \beta_1 \Delta I_t + \beta_2 UM_{i,t} + \beta_3 USP_{i,t} + \beta_4 JAP_i + a_i + u_{i,t}$$

(6) $$\Delta CDS_{i,t} = \alpha + \beta_1 \Delta I_t + \beta_2 DM_{i,t} + \beta_3 DSP_{i,t} + \beta_4 JAP_i + a_i + u_{i,t}$$

where $\Delta CDS_{i,t}$ is the raw spread change for entity *i* on day *t*

$\Delta I_t$ is the CDS spread index change on day *t*

$UM_{i,t}/DM_{i,t}$ is a dummy variable taking the value 1 around Moody's events

$USP_{i,t}/DSP_{i,t}$ is a dummy variable taking the value 1 around S&P events

$JAP_i$ is a dummy variable taking the value 1 if the entity is Japanese

$a_i$ denotes entity-specific fixed effects and $u_{i,t}$ denotes the *i* and *t* variant error

I also want to extend my analysis to study the time intervals close to the announcement day, and therefore I add three event time dummy variables, which take values during the following event windows: [-7, -2], [-1, 1], and [2, 7]. Again, using the Huber/White/sandwich estimator of variance, I estimate separate models for upgrades (7) and for downgrades (8) as follows:

(7) $$\Delta CDS_{i,t} = \alpha + \beta_1 \Delta I_t + \beta_2 U72_t + \beta_3 U101_t + \beta_4 U27_t + \beta_5 JAP_i + a_i + u_{i,t}$$

(8) $$\Delta CDS_{i,t} = \alpha + \beta_1 \Delta I_t + \beta_2 D72_t + \beta_3 D101_t + \beta_4 D27_t + \beta_5 JAP_i + a_i + u_{i,t}$$

where $\Delta CDS_{i,t}$, $\Delta I_t$, $JAP_{i,t}$, $a_i$ and $u_{i,t}$ are as above

$U72_t/D72_t$ is a dummy variable taking the value 1 during interval [-7, -2]

$U101_t/D101_t$ is a dummy variable taking the value 1 during interval [-1, 1]

$U27_t/D27_t$ is a dummy variable taking the value 1 during interval [2, 7]

## 5. Results

### 5.1. Univariate event study results

This section reports the results of my univariate event study. I analyze the CDS market reaction to upgrades and downgrades within and across agencies using both the uncontaminated event sample and the subsamples of upgrades and downgrades within investment grade. I also report the results of two exploratory subsamples: upgrades to investment grade and downgrades to non-investment grade, as well as of two exploratory rating events: new ratings and rating withdrawals. In my analysis, I consider event windows, where both the parametric- and the non-parametric test reject the null hypothesis to show the most powerful evidence.[12]

*5.1.1 Uncontaminated sample: Upgrades and downgrades*

First, I consider the CDS market reaction to upgrades and downgrades using the entire uncontaminated sample, presented in Table 2. Both events show a significant announcement effect (at a 10%-level) in the expected direction during the event window [-1, 1], and thereby provide evidence for *H1*. Although $\overline{CASC}_{\tau1,\tau2}$is positive around upgrades, the sign of all CASCs suggests that the spreads decrease significantly on announcement. Downgrades, on the other hand, increase on average by 3.922 basis points during the announcement window. I also find partial evidence for *H2*, as the markets react significantly to both agencies' upgrade announcement but only S&P downgrades generate economically and statistically significant abnormal changes in the markets.

The CDS market starts to anticipate downgrades of both agencies as early as 90 day prior to the rating announcement, whereas Moody's upgrades are anticipated earlier (starting 60 days vs. 30 days earlier[13]). Interestingly, the CDS spreads also show a significant post-announcement effect at a 1%-level starting 61 days after an upgrade (31 days after a downgrade). The CDS spreads also continue to decrease on average by 1.762 basis points after a rating upgrade (at a 5%-level) in the event window [2, 10], which largely results from large proportion of similar S&P observations. Changes remain insignificant immediately after the rating event for Moody's upgrades and both agencies' downgrades.

[12] Similarly to Galil and Soffer (2011), the possibility of abnormally distributed spread changes suggests me to prefer the non-parametric test results over the parametric test results.

[13] S&P shows a statistically significant (at a 10%-level) decrease in price already during [-60, -31], but the mean CASC remains close to zero.

**Table 2. CDS market reaction around upgrades and downgrades – uncontaminated sample**

| Days | | | [-90, -61] | [-60, -31] | [-30, -11] | [-10, -2] | [-1, 1] | [2, 10] | [11, 30] | [31, 60] | [61, 90] |
|---|---|---|---|---|---|---|---|---|---|---|---|
| *Panel A: Mean CASCs around upgrades* | | | | | | | | | | | |
| Moody's | All | N | 98 | 98 | 98 | 98 | 98 | 98 | 98 | 98 | 98 |
| or S&P | | CASC-mean (bps) | 3.583 | -1.588 | -4.373 | -1.874 | 1.729 | -1.762 | 2.275 | 3.717 | -12.997 |
| | | *t*-test *p*-value | 0.999 | 0.152 | 0.001 *** | 0.020 ** | 0.771 | 0.036 ** | 0.995 | 0.999 | <0.001 *** |
| | | % of CASC>0 | 51.02 | 42.86 | 46.94 | 46.94 | 42.86 | 46.94 | 46.94 | 45.92 | 38.38 |
| | | sign test *p*-value | 0.540 | 0.065 * | 0.240 | 0.240 | 0.065 * | 0.240 | 0.240 | 0.182 | 0.008 *** |
| Moody's | All | N | 19 | 19 | 19 | 19 | 19 | 19 | 19 | 19 | 19 |
| | | CASC-mean (bps) | 6.867 | -7.416 | -5.922 | -3.019 | -1.710 | 0.999 | 7.543 | -14.061 | -5.324 |
| | | *t*-test *p*-value | 0.993 | 0.016 ** | 0.103 | 0.089 * | 0.064 * | 0.830 | 0.999 | <0.001 *** | 0.000 *** |
| | | % of CASC>0 | 68.42 | 42.11 | 47.37 | 52.63 | 42.11 | 52.63 | 52.63 | 31.58 | 36.84 |
| | | sign test *p*-value | 0.917 | 0.180 | 0.324 | 0.500 | 0.180 | 0.500 | 0.500 | 0.032 ** | 0.084 * |
| S&P | All | N | 79 | 79 | 79 | 79 | 79 | 79 | 79 | 79 | 79 |
| | | CASC-mean (bps) | 2.793 | -0.187 | -4.001 | 1.599 | 2.555 | -2.426 | 1.008 | -1.229 | -14.842 |
| | | *t*-test *p*-value | 0.993 | 0.457 | 0.001 *** | 0.959 | 0.806 | 0.023 ** | 0.852 | 0.119 | <0.001 *** |
| | | % of CASC>0 | 46.84 | 43.04 | 46.84 | 45.57 | 43.04 | 45.57 | 45.57 | 49.37 | 39.24 |
| | | sign test *p*-value | 0.250 | 0.088 * | 0.250 | 0.184 | 0.088 * | 0.184 | 0.184 | 0.411 | 0.021 ** |
| *Panel B: Mean CASCs around downgrades* | | | | | | | | | | | |
| Moody's | All | N | 152 | 152 | 152 | 152 | 152 | 152 | 152 | 152 | 152 |
| or S&P | | CASC-mean (bps) | 6.658 | -4.653 | 5.790 | 7.380 | 3.922 | -1.695 | -4.409 | 7.441 | 1.507 |
| | | *t*-test *p*-value | <0.001 *** | 0.999 | 0.001 *** | <0.001 *** | 0.067 * | 0.812 | 0.999 | <0.001 *** | 0.143 |
| | | % of CASC>0 | 54.61 | 53.29 | 53.95 | 61.18 | 53.29 | 48.03 | 48.68 | 51.32 | 49.34 |
| | | sign test *p*-value | 0.146 | 0.233 | 0.186 | 0.004 *** | 0.233 | 0.715 | 0.657 | 0.404 | 0.596 |
| Moody's | All | N | 61 | 61 | 61 | 61 | 61 | 61 | 61 | 61 | 61 |
| | | CASC-mean (bsp) | 7.247 | -4.704 | 2.981 | 10.719 | 1.174 | -3.132 | 10.975 | 2.118 | 2.295 |
| | | *t*-test *p*-value | <0.001 *** | 0.999 | 0.103 | 0.001 *** | 0.231 | 0.971 | <0.001 *** | 0.228 | 0.181 |
| | | % of CASC>0 | 50.82 | 52.46 | 47.54 | 57.38 | 55.74 | 45.90 | 49.18 | 54.10 | 54.10 |
| | | sign test *p*-value | 0.500 | 0.399 | 0.696 | 0.153 | 0.221 | 0.779 | 0.601 | 0.305 | 0.305 |
| S&P | All | N | 91 | 91 | 91 | 91 | 91 | 91 | 91 | 91 | 91 |
| | | CASC-mean (bsp) | 7.237 | -2.156 | 7.507 | 3.648 | 5.338 | -0.647 | -13.588 | 11.471 | 6.573 |
| | | *t*-test *p*-value | <0.001 *** | 0.932 | 0.001 *** | 0.027 ** | 0.084 * | 0.586 | 0.999 | <0.001 *** | <0.001 *** |
| | | % of CASC>0 | 58.24 | 53.85 | 57.14 | 62.64 | 50.55 | 50.55 | 48.35 | 50.55 | 46.15 |
| | | sign test *p*-value | 0.071 * | 0.265 | 0.104 | 0.010 *** | 0.500 | 0.500 | 0.662 | 0.500 | 0.799 |

This table presents the means of cumulative abnormal spread changes around upgrades and downgrades during time intervals $[\tau_1, \tau_2]$. Sample includes Moody's (S&P) rating events above B3 (B-) from October 1, 2007 to March 18, 2014 with CDS spread data available in the event window [-90, 90]. Null hypothesis for upgrades (downgrades) is under Student's *t*-test: CASC-mean $\geq 0$ (CASC-mean $\leq 0$) and under Wilcoxon sign test: % of positive CASC $\geq 0.5$ (% of positive CASC $\leq 0.5$). * indicates $p<0.10$, ** indicates $p<0.05$, *** indicates $p<0.01$.

To allow for better comparability between my results, and the past studies using investment grade events in their analyses (e.g. Hull et al., 2004; Norden and Weber, 2004), I also conduct the same tests using two smaller subsamples consisting of rating changes occurring only within investment grade (Table 3). The announcement effect of upgrades remains significant at a 10%-level with an average decrease of 1.176 basis points on announcement. CDS spreads decrease significantly (at the 5%-level) by 2.349 basis points around Moody's upgrades, whereas S&P upgrades do not generate abnormal changes, providing evidence against *H2*. Downgrades do not cause a significant change around [-1, 1], aside from the spreads increasing on average by 9.896 basis points on S&P downgrade announcement (significant at a 5%-level), suggesting that the findings do not support either *H1* or *H2*.

**Table 3. CDS market reaction around upgrades and downgrades within investment grade**

| Days | | | [-90, -61] | [-60, -31] | [-30, -11] | [-10, -2] | [-1, 1] | [2, 10] | [11, 30] | [31, 60] | [61, 90] |
|---|---|---|---|---|---|---|---|---|---|---|---|
| *Panel A: Mean CASCs around upgrades within investment grade* | | | | | | | | | | | |
| Moody's | UIG | N | 68 | 68 | 68 | 68 | 68 | 68 | 68 | 68 | 68 |
| or S&P | | CASC-mean (bps) | 3.533 | 2.322 | -0.063 | -3.321 | -1.176 | -2.430 | 3.068 | 0.891 | -15.018 |
| | | *t*-test *p*-value | 0.999 | 0.991 | 0.475 | <0.001 *** | 0.093 * | 0.001 *** | 0.999 | 0.897 | <0.001 *** |
| | | % of CASC>0 | 50.00 | 44.12 | 44.12 | 41.18 | 45.59 | 44.12 | 47.06 | 44.12 | 38.24 |
| | | sign test *p*-value | 0.548 | 0.138 | 0.138 | 0.057 * | 0.198 | 0.138 | 0.272 | 0.138 | 0.019 ** |
| Moody's | UIG | N | 17 | 17 | 17 | 17 | 17 | 17 | 17 | 17 | 17 |
| | | CASC-mean (bps) | 3.999 | 1.068 | -9.254 | -3.468 | -2.349 | 1.548 | 11.825 | -12.411 | -1.408 |
| | | *t*-test *p*-value | 0.998 | 0.756 | <0.001 *** | 0.002 *** | 0.044 ** | 0.902 | 0.999 | <0.001 *** | 0.150 |
| | | % of CASC>0 | 64.71 | 47.06 | 47.06 | 52.94 | 35.29 | 52.94 | 58.82 | 35.29 | 41.18 |
| | | sign test *p*-value | 0.834 | 0.315 | 0.315 | 0.500 | 0.072 * | 0.500 | 0.686 | 0.072 * | 0.166 |
| S&P | UIG | N | 51 | 51 | 51 | 51 | 51 | 51 | 51 | 51 | 51 |
| | | CASC-mean (bps) | 3.378 | 2.486 | 3.001 | -3.271 | -0.785 | -3.756 | 0.149 | 5.326 | -19.554 |
| | | *t*-test *p*-value | 0.999 | 0.983 | 0.986 | 0.002 *** | 0.202 | <0.001 *** | 0.578 | 0.999 | <0.001 *** |
| | | % of CASC>0 | 45.10 | 43.14 | 43.14 | 37.25 | 49.02 | 41.18 | 43.14 | 47.06 | 37.25 |
| | | sign test *p*-value | 0.201 | 0.131 | 0.131 | 0.024 ** | 0.390 | 0.080 * | 0.131 | 0.288 | 0.024 ** |
| *Panel B: Mean CASCs around downgrades within investment grade* | | | | | | | | | | | |
| Moody's | DIG | N | 121 | 121 | 121 | 121 | 121 | 121 | 121 | 121 | 121 |
| or S&P | | CASC-mean (bps) | 19.252 | 24.216 | 22.783 | 17.700 | 10.734 | -2.444 | -7.557 | 0.764 | -16.410 |
| | | *t*-test *p*-value | <0.001 *** | <0.001 *** | <0.001 *** | <0.001 *** | 0.194 | 0.655 | 0.928 | 0.431 | 0.999 |
| | | % of CASC>0 | 57.85 | 54.55 | 53.72 | 61.16 | 54.55 | 51.24 | 49.59 | 52.07 | 46.28 |
| | | sign test *p*-value | 0.051 * | 0.182 | 0.234 | 0.009 *** | 0.182 | 0.428 | 0.572 | 0.358 | 0.818 |
| Moody's | DIG | N | 55 | 55 | 55 | 55 | 55 | 55 | 55 | 55 | 55 |
| | | CASC-mean (bsp) | 0.651 | 46.549 | 19.808 | 22.647 | 11.554 | -7.040 | -32.026 | -21.813 | 10.614 |
| | | *t*-test *p*-value | 0.287 | <0.001 *** | 0.001 *** | <0.001 *** | 0.320 | 0.756 | 0.997 | 0.999 | 0.001 *** |
| | | % of CASC>0 | 49.09 | 54.55 | 45.45 | 56.36 | 54.55 | 47.27 | 49.09 | 54.55 | 58.18 |
| | | sign test *p*-value | 0.606 | 0.295 | 0.791 | 0.209 | 0.295 | 0.705 | 0.606 | 0.295 | 0.140 |
| S&P | DIG | N | 66 | 66 | 66 | 66 | 66 | 66 | 66 | 66 | 66 |
| | | CASC-mean (bsp) | 43.400 | 10.279 | 24.729 | 12.778 | 9.896 | 1.622 | 48.866 | 58.946 | -59.354 |
| | | *t*-test *p*-value | <0.001 *** | <0.001 *** | <0.001 *** | <0.001 *** | 0.050 ** | 0.414 | <0.001 *** | <0.001 *** | 0.999 |
| | | % of CASC>0 | 66.67 | 56.06 | 60.61 | 65.15 | 56.06 | 56.06 | 51.52 | 51.52 | 36.36 |
| | | sign test *p*-value | 0.005 *** | 0.195 | 0.054 * | 0.009 *** | 0.195 | 0.195 | 0.451 | 0.451 | 0.991 |

This table presents the means of cumulative abnormal spread changes around upgrades (UIG) and downgrades (DIG) within investment grade during time intervals $[\tau_1, \tau_2]$. Sample includes Moody's (S&P) rating events above Baa3 (BBB-) from October 1, 2007 to March 18, 2014 with CDS spread data available in the event window [-90, 90]. Null hypothesis for upgrades (downgrades) is under Student's *t*-test: CASC-mean $\geq 0$ (CASC-mean $\leq 0$) and under Wilcoxon sign test: % of positive CASC $\geq 0.5$ (% of positive CASC $\leq 0.5$). * indicates $p<0.10$, ** indicates $p<0.05$, *** indicates $p<0.01$.

Upgrades (downgrades) result in a pre-announcement decrease (increase) in the average CDS spread starting 10 days (90 days) before the rating event, which is significant at a 1%-level. As with the uncontaminated sample, the CDS markets anticipates Moody's upgrades earlier (starting 30 days vs. 10 days earlier) and S&P downgrades are anticipated before Moody's downgrades (starting 90 days vs. 60 days earlier). The CDS markets show an average abnormal decrease of 2.430 basis points, significant at a 1%-level, after an upgrade in the event window [2, 10]. Aside from the period-end similarities with the previous results, no other post announcement effects are detected.

### *5.1.2 Exploratory samples: Changes around investment threshold, new ratings and withdrawals*

Micu et al. (2006) find the strongest CDS market reaction during ratings events that occur close to the investment threshold between investment grade and non-investment grade. To analyze whether the market reaction is similar in the Asian CDS markets, I test two exploratory subsamples; upgrades to investment grade (UTIG) and downgrades to non-investment grade (DTNIG). Additionally, as previous literature on CDS market reaction (e.g. Hull et al., 2004; Norden and Weber, 2004) has mostly studied the announcement effect around positive and negative rating events, I extend the analysis to two implicitly neutral announcements, new ratings and rating withdrawals, to analyze whether these two omitted events also cause abnormal spread changes in the Asian CDS markets.

Table 4 presents the results for rating changes between investment- and non-investment grades. On announcement, UTIG show a significant market reaction at a 10%-level, whereas a DTNIG announcement does not result in any abnormal changes, thus providing only partial evidence for *H1*. Neither event results in significant changes in CDS spreads shortly before the announcement, but UTIG (DTNIG) are anticipated already during the period [-60, -31] ([-90, -61]) at a 1%-level. CDS spreads increase significantly (at a 1%-level) after a DTNIG announcement throughout all the periods after second post-announcement day. Abnormal changes in the CDS market can also be detected after an UTIG announcement at the 1%-level, but the post-announcement reaction starts later as with downgrades.

**Table 4. CDS market reaction around upgrades and downgrades between investment- and non-investment grade**

| Days | | | [-90, -61] | [-60, -31] | [-30, -11] | [-10, -2] | [-1, 1] | [2, 10] | [11, 30] | [31, 60] | [61, 90] |
|---|---|---|---|---|---|---|---|---|---|---|---|
| *Panel A: Mean CASCs around upgrades to investment grade* | | | | | | | | | | | |
| Moody's | UTIG | N | 10 | 10 | 10 | 10 | 10 | 10 | 10 | 10 | 10 |
| or S&P | | CASC-mean (bps) | 18.307 | -80.622 | -12.804 | -1.160 | -3.549 | -2.488 | -9.359 | 10.414 | -13.784 |
| | | *t*-test *p*-value | 0.999 | <0.001 *** | 0.081 * | 0.398 | 0.169 | 0.131 | <0.001 *** | 0.999 | <0.001 *** |
| | | % of CASC>0 | 50.00 | 10.00 | 30.00 | 60.00 | 30.00 | 40.00 | 40.00 | 40.00 | 30.00 |
| | | sign test *p*-value | 0.623 | 0.001 *** | 0.055 * | 0.623 | 0.055 * | 0.172 | 0.172 | 0.172 | 0.055 * |
| *Panel B: Mean CASCs around downgrades to non-investment grade* | | | | | | | | | | | |
| Moody's | DTNIG | N | 13 | 13 | 13 | 13 | 13 | 13 | 13 | 13 | 13 |
| or S&P | | CASC-mean (bps) | 173.777 | -212.019 | -65.942 | 6.739 | 11.704 | 23.663 | 141.722 | 69.941 | 531.851 |
| | | *t*-test *p*-value | <0.001 *** | 0.999 | 0.998 | 0.337 | 0.127 | 0.010 *** | <0.001 *** | 0.006 *** | <0.001 *** |
| | | % of CASC>0 | 69.23 | 53.85 | 69.23 | 61.54 | 69.23 | 30.77 | 69.23 | 53.85 | 46.15 |
| | | sign test *p*-value | 0.133 | 0.500 | 0.133 | 0.291 | 0.133 | 0.954 | 0.133 | 0.500 | 0.709 |

This table presents the means of cumulative abnormal spread changes around upgrades (UTIG) and downgrades (DTNIG) between investment grade and non-investment grade during time intervals $[\tau_1, \tau_2]$. Sample includes Moody's (S&P) rating events above B3 (B-) from October 1, 2007 to March 18, 2014 with CDS spread data available in the event window [-90, 90]. Null hypothesis for upgrades (downgrades) is under Student's *t*-test: CASC-mean $\geq 0$ (CASC-mean $\leq 0$) and under Wilcoxon sign test: % of positive CASC $\geq 0.5$ (% of positive CASC $\leq 0.5$). * indicates $p<0.10$, ** indicates $p<0.05$, *** indicates $p<0.01$.

The exploratory analysis on new ratings and rating withdrawals, presented in Table 5, is conducted using two-tail tests, and thereby I may only note whether abnormal spread changes of any direction occur during the event windows. New ratings do not result in a significant announcement effect, whereas a significant (at a 5%-level) spread change is observed during the announcement window

[-1, 1] around a rating withdrawal. In the case of new ratings, significant spread changes (at a 1%-level) can be detected around all pre-announcement event windows and all but one post-announcement intervals [30, 60]. Also, CDS markets react to withdrawals with significant abnormal spread changes at a 1%-level during all pre-announcement time windows. Contrary to the new ratings, significant changes are only observed during one period, [31, 60], starting 31 days after the rating has been withdrawn.

**Table 5. CDS market reaction around new ratings and rating withdrawals**

| Days | | | [-90, -61] | [-60, -31] | [-30, -11] | [-10, -2] | [-1, 1] | [2, 10] | [11, 30] | [31, 60] | [61, 90] |
|---|---|---|---|---|---|---|---|---|---|---|---|
| *Panel A: Mean CASCs around new ratings* | | | | | | | | | | | |
| Moody's | All | N | 30 | 30 | 30 | 30 | 30 | 30 | 30 | 30 | 30 |
| or S&P | | CASC-mean (bps) | -4.652 | -19.324 | -6.307 | 4.897 | -0.441 | 4.218 | -7.575 | -3.590 | 9.115 |
| | | *t*-test *p*-value | 0.002 *** | <0.001 *** | <0.001 *** | 0.022 *** | 0.558 | <0.001 *** | <0.001 *** | 0.132 | <0.001 *** |
| | | % of CASC>0 | 40.00 | 46.67 | 40.00 | 46.67 | 56.67 | 63.33 | 40.00 | 50.00 | 56.67 |
| | | sign test *p*-value | 0.200 | 0.585 | 0.200 | 0.585 | 0.585 | 0.200 | 0.200 | 0.856 | 0.585 |
| *Panel B: Mean CASCs around rating withdrawals* | | | | | | | | | | | |
| Moody's | All | N | 93 | 93 | 93 | 93 | 93 | 93 | 93 | 93 | 93 |
| or S&P | | CASC-mean (bps) | 81.673 | -20.333 | 17.842 | 11.322 | 6.313 | 2.752 | -0.127 | -13.813 | 0.441 |
| | | *t*-test *p*-value | <0.001 *** | <0.001 *** | <0.001 *** | <0.001 *** | 0.110 | 0.193 | 0.943 | <0.001 *** | 0.842 |
| | | % of CASC>0 | 56.99 | 35.48 | 65.59 | 72.04 | 62.37 | 51.61 | 54.84 | 60.22 | 72.04 |
| | | sign test *p*-value | 0.213 | 0.003 *** | 0.003 *** | <0.001 *** | 0.022 ** | 0.836 | 0.407 | 0.061 * | <0.001 *** |

This table presents the means of cumulative abnormal spread changes around new ratings and rating withdrawals between during time intervals $[\tau_1, \tau_2]$. Sample includes Moody's (S&P) rating events above B3 (B-) from October 1, 2007 to March 18, 2014 with CDS spread data available in the event window [-90, 90]. Null hypothesis for new ratings and rating withdrawals is under Student's *t*-test: CASC-mean = 0 and under Wilcoxon sign test: % of positive CASC = 0.5. * indicates $p<0.10$, ** indicates $p<0.05$, *** indicates $p<0.01$.

## 5.2. Multivariate analysis results

This section presents the results of my multivariate analysis focusing on upgrades and downgrades. First, I study the announcement window [-1, 1] around the rating events using a sample of events across agencies. I test whether the CDS markets react to the effects on announcement and whether the effects differ from one agency to another. Second, along with the initial announcement window, I also analyze the market reaction on pre- and post-announcement time intervals close to the rating event. I run the second set of regressions on event samples within and across agencies.

Table 6 presents the regression results for the announcement effect regressions. Around upgrades, a significant abnormal change during the event window [-1, 1] is unobservable for both Moody's and S&P. A similar result applies to downgrades as well, and thereby the results provide evidence against *H1*, but for *H2*, as agency specific variables show same results within the announcement window. Interestingly, the spreads in the Japanese CDS market seem to be significantly higher than in the other Asian markets (at a 5%-level) around downgrades, but not around upgrades. The coefficient of $JAP_t$ around downgrades suggests that Japanese credit default swaps have 0.459 basis points higher spreads than non-Japanese CDSs.

**Table 6. CDS market regressions across agencies – focus on announcement window [-1, 1]**

| Dep. Var. $\Delta CDS$ | Moody's or S&P upgrades | | | | Moody's or S&P downgrades | | | |
|---|---|---|---|---|---|---|---|---|
| | Coeff. | Robust st. error | *p*-value | | Coeff. | Robust st. error | *p*-value |
| *ΔI* | 0.692 | 0.196 | <0.001 *** | *ΔI* | 0.819 | 0.044 | <0.001 *** |
| *UM* | -0.355 | 0.442 | 0.422 | *DM* | 0.160 | 0.752 | 0.831 |
| *USP* | 0.882 | 1.257 | 0.483 | *DSP* | 1.720 | 1.456 | 0.237 |
| *JAP* | -0.197 | 0.202 | 0.328 | *JAP* | 0.459 | 0.225 | 0.041 ** |
| Const | 0.026 | 0.106 | 0.810 | Const | -0.163 | 0.195 | 0.403 |
| *n* | 17738 | | | *n* | 27512 | | |
| *R^2* | 0.138 | | | *R^2* | 0.130 | | |

This table presents the results of regressing daily raw CDS spread changes $\Delta CDS_{i,t}$ on daily CDS spread index change $\Delta I_t$. During the announcement window [-1, 1], dummy $UM_{i,t}$ ($DM_{i,t}$) takes the value 1 around a Moody's upgrade (downgrade) and dummy $USP_{i,t}$ ($DSP_{i,t}$) takes the value 1 around an S&P upgrade (downgrade). Dummy variable $JAP_i$ controls for region-specific differences and takes the value 1 if the CDS reference entity is Japanese. The pooled regressions are conducted using the Huber/White/sandwich estimator of variance. Sample includes Moody's (S&P) rating events above B3 (B-) from October 1, 2007 to March 18, 2014 with CDS spread data available in the event window [-90, 90]. * indicates $p<0.10$, ** indicates $p<0.05$, *** indicates $p<0.01$.

Table 7 summarizes the regression results for upgrades; first presenting them for the joint sample (Moody's and S&P) followed by the results of the agency-specific samples (Moody's or S&P). Contrary to the univariate event study around upgrades, the multivariate analysis does not show a significant CDS market reaction on announcement. I addition, the statistically insignificant coefficient of $U101_t$, offers partial evidence against *H1*. On the agency specific level, neither S&P nor Moody's upgrades generate abnormal spread changes during any of the event windows on and around the rating announcement, which offers some support for *H2*. The estimation suggests that no market-specific differences occur around upgrades of any rating agency.

**Table 7. CDS market regressions within and across agencies around an upgrade announcement**

| Dep. Var. ΔCDS | Moody's or S&P | | | Moody's | | | S&P | | |
|---|---|---|---|---|---|---|---|---|---|
| | Coeff. | Robust st. error | *p*-value | Coeff. | Robust st. error | *p*-value | Coeff. | Robust st. Error | *p*-value |
| ΔI | 0.692 | 0.150 | <0.001 *** | 1.002 | 0.009 | <0.001 *** | 0.510 | 0.299 | 0.088 * |
| U72 | -0.204 | 0.314 | 0.515 | -0.052 | 0.631 | 0.935 | -0.287 | 0.368 | 0.436 |
| U101 | 0.628 | 1.021 | 0.538 | -0.454 | 0.458 | 0.320 | 0.828 | 1.209 | 0.493 |
| U27 | -0.223 | 0.427 | 0.601 | 0.137 | 0.405 | 0.735 | -0.498 | 0.513 | 0.332 |
| JAP | -0.199 | 0.218 | 0.323 | -0.385 | 0.555 | 0.480 | -0.185 | 0.172 | 0.282 |
| Const | 0.040 | 0.111 | 0.715 | 0.004 | 0.086 | 0.962 | 0.076 | 0.142 | 0.592 |
| *n* | 17738 | | | 3439 | | | 14299 | | |
| *R^2* | 0.138 | | | 0.451 | | | 0.062 | | |

This table presents the results of regressing daily raw CDS spread changes $\Delta CDS_{i,t}$ on daily CDS spread index change $\Delta I_t$. There are three event time dummies: $U72_t$ takes the value 1 during the event interval [-7, -2], $U101_t$ takes the value 1 during the event interval [-1, 1] and $U27_t$ takes the value 1 during event interval [2, 7]. Dummy variable $JAP_i$ controls for region-specific differences and takes the value 1 if the CDS reference entity is Japanese. The pooled regressions are conducted using the Huber/White/sandwich estimator of variance. Sample includes Moody's (S&P) rating events above B3 (B-) from October 1, 2007 to March 18, 2014 with CDS spread data available in the event window [-90, 90]. * indicates $p<0.10$, ** indicates $p<0.05$, *** indicates $p<0.01$.

Table 8 exemplifies the estimation results of CDS spread changes around downgrades. I estimate the raw spread changes with a joint sample (Moody's or S&P), as well as with the single-agency samples. Neither the combined results, nor the agency-specific results show evidence that downgrades would result in abnormal spread changes in the Asian CDS markets on announcement, offering evidence against *H1*. However, a uniform result across agencies offers supportive evidence for *H2*. As the coefficients of $D72_t$ and $D27_t$ remain insignificant within and across agencies, the multivariate analysis results suggest the CDS markets do not react significantly during event windows close to the announcement. The control variable $JAP_i$ suggest that, at the 5%-level, Japanese credit defaults swap spreads are 0.453 basis points above the spreads of non-Japanese CDSs' in the [-90, 90] event window around a rating downgrade.

**Table 8. CDS market regressions within and across agencies around a downgrade announcement**

| Dep. Var. ΔCDS | Moody's or S&P | | | | Moody's | | | | S&P | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | Coeff. | Robust st. error | *p*-value | | Coeff. | Robust st. error | *p*-value | | Coeff. | Robust st. error | *p*-value | |
| ΔI | 0.819 | 0.044 | <0.001 | *** | 1.021 | 0.037 | <0.001 | *** | 0.654 | 0.075 | <0.001 | *** |
| D72 | 0.383 | 0.542 | 0.479 | | 0.941 | 0.932 | 0.313 | | 0.239 | 0.647 | 0.712 | |
| D101 | 1.092 | 0.926 | 0.238 | | 0.245 | 0.765 | 0.749 | | 1.730 | 1.459 | 0.236 | |
| D27 | -0.457 | 0.690 | 0.508 | | -0.664 | 0.661 | 0.316 | | -0.190 | 1.071 | 0.859 | |
| JAP | 0.453 | 0.225 | 0.044 | ** | 0.528 | 0.411 | 0.199 | | 0.439 | 0.271 | 0.105 | |
| Const | -0.157 | 0.195 | 0.422 | | -0.256 | 0.384 | 0.505 | | -0.135 | 0.223 | 0.544 | |
| *n* | 27512 | | | | 11041 | | | | 16471 | | | |
| *R^2* | 0.130 | | | | 0.215 | | | | 0.079 | | | |

This table presents the results of regressing daily raw CDS spread changes $\Delta CDS_{i,t}$ on daily CDS spread index change $\Delta I_t$. There are three event time dummies: $D72_t$ takes the value 1 during the event interval [-7, -2], $D101_t$ takes the value 1 during the event interval [-1, 1] and $D27_t$ takes the value 1 during event interval [2, 7]. Dummy variable $JAP_i$ controls for region-specific differences and takes the value 1 if the CDS reference entity is Japanese. The pooled regressions are conducted using the Huber/White/sandwich estimator of variance. Sample includes Moody's (S&P) rating events above B3 (B-) from October 1, 2007 to March 18, 2014 with CDS spread data available in the event window [-90, 90]. * indicates $p$<0.10, ** indicates $p$<0.05, *** indicates $p$<0.01.

Similarly to Norden and Weber (2004), to test the robustness of my multivariate analysis, I run the same set of regressions using smaller subsamples of upgrades and downgrades cleared of observations of entities with extreme abnormal spread changes.[14] The results of the check show similar results with the regressions on the full sample. All event windows before, on and after the announcement day appear free of abnormal spread changes. Additionally, the coefficient of $JAP_i$ still shows significance at the 5%-level. As the results are similar to those obtained with the full event sample containing extreme spread changes, my results appear to be robust to changes in the original test settings.

[14] All events of entities with a daily abnormal spread change of over 300 or less than -300 basis points are dropped from the subsamples. All events of ENN Energy Holdings Ltd, IOI Corporation Bhd, KIA Motors Corp Nippon Sheet Glass Co Ltd, Razvedka Dobycha KazMunayGaz AO, Road King Infrastructure Ltd, Shinsei Bank Ltd, Tokyo Electric Power Co Inc and Wan Hai Lines Ltd are omitted from the robustness check sample.

## 6. Discussion of results

### 6.1 Univariate event study results

My univariate analysis results seem to be in line with the past findings on CDS market reaction to credit rating changes to a great extent. This suggests that the effect I observe in the Asian CDS markets is consistent with the market reaction past studies have observed, mainly in the North American and European markets.

My findings on announcement effects around downgrades are very similar to Hull et al. (2004) and Norden and Weber (2004). The observed market reaction to downgrades offers partial support for *H1*, as downgrades spark a positive spread change during the rating announcement window, suggesting that the rating change brings new information to the markets on announcement. On the other hand, the result showing that CDS markets start to anticipate future downgrades as early as 90 days prior to the rating event, suggests that the CDS markets already adapt to the publicly available information before announcement and the observed announcement effect is generated by the non-public content of the rating. The results also imply that the informational content of Moody's and S&P downgrades is viewed differently by the Asian CDS market, as an S&P downgrade results in a CDS spread increase of 5.338 basis points on announcement while the market reaction to a Moody's downgrade remains statistically insignificant. This finding provides evidence against *H2* and denotes the lack of informational content of Moody's downgrades in the Asian markets.

My analysis also suggests that significant abnormal spread changes are observed around upgrade announcements, offering support for *H1*. The results differ from the early CDS market studies, but the increased informational content can be explained by investors' increasing interest in positive announcements (e.g. Micu et al., 2006; Finnerty et al., 2013) or by a strengthened credit rating process (Finnerty et al., 2013). The CDS markets start to anticipate positive rating events around 30 days prior to the rating event, which is later than the observed anticipation of downgrades. This is similar to Finnerty et al. (2013), who conclude that upgrades are not as anticipated as downgrades are. As previous literature does not support my findings of the abnormal spread changes observed 61 days after an upgrade, I discuss on a speculative note, that they may already be contamination by another market event rather than a reaction to the announced upgrade. My finding, that the announcement effect to upgrades is of similar sign and magnitude for both Moody's and S&P, provides evidence for *H2*, and suggests that the Asian CDS markets view the informational content of Moody's and S&P upgrade announcements to be similar.

My results, suggesting that markets anticipate UTIG well before the actual announcement, and that the market reaction to DTNIG starts only after the rating announcement, are both against *H1*. However, instead of concluding on the lack of information on announcement, the findings can be explained to result from regulatory constraints usually set at Moody's Baa3 (S&P BBB-). The constraints may only allow institutional investors to invest in entities rated above the threshold grade, and thereby changes around the investment threshold grade generate more trading volume around the announcement than other rating events do (Steiner and Heinke, 2001). Micu et al. (2006) also note the price pressure regarding rating events occurring close to the threshold grade and they conclude that CDS markets react with highest spread changes around the threshold ratings. However, due to fairly small samples of rating events to and from the investment grade, the topic would require further study to fully draw conclusions on the CDS markets' reaction to threshold grade rating changes.

My exploratory study on new ratings and rating withdrawals also yields results that provide evidence against *H1*, as the CDS markets show significant changes before and after a new rating, and before a rating withdrawal is announced. As the announcement of a new rating or a rating withdrawal does not contain public information, anticipation of these events does not seem rational and a market reaction could only be expected on or after the announcement. Also, due to the shortcomings of the "Not rated" –index, drawing significant conclusions on the results does not seem relevant. However, on a speculative note, rather than resulting from the actual rating announcement, the observed abnormal reaction may actually capture the spread movement resulting from other common factors to entities seeking a new rating or having their rating withdrawn. An entity seeking a new rating is willing to have its creditworthiness publicly announced to the market i.e. when raising new financing, which could imply that the entity is performing exceptionally well overall and seeking a rating is more a result of the abnormal performance rather than vice versa. Similarly, rating withdrawals usually occur around bankruptcies and corporate reorganizations and thereby, instead of an announcement effect, the abnormal changes around rating withdrawals may actually be the market effect of an anticipated acquisition or a default.

### 6.2 Multivariate analysis results

Similarly to Norden and Weber (2004), my multivariate analysis results differ slightly from the results obtained with the univariate methods. As coefficients of the announcement day dummy variables are insignificant, the regression results do not show any evidence for *H1*, whereas the univariate Student's *t*-test and Wilcoxon sign-test also provide some evidence for a significant announcement effect to downgrades and upgrades. Also, as the coefficients of the event time dummies taking value 1 during the period [-7, -2] remain largely insignificant, the results suggest that the Asian CDS markets adjust to the information on changes in default risk already seven trading days before the actual upgrade or downgrade announcement is made. Similarly to the univariate analysis results, as the markets do not show a lagging reaction to the rating announcement in the event windows following an upgrade or a downgrade announcement, the Asian CDS markets seem to have incorporated all new default risk information already on and before the actual announcement. As my regression results are similar across rating agencies and neither an upgrade nor a downgrade result in abnormal spread changes on rating announcement, the multivariate analysis results provide some evidence for *H2*.

Interestingly, as the coefficient $JAP_i$ is significantly greater than zero, the multivariate analysis implies that around a rating downgrade, the Japanese CDS spreads quote approximately 0.459 basis points higher that the non-Japanese CDS spreads. Li et al. (2006) find that Japanese markets react stronger to American ratings that to ratings by local agencies, whereas Ferri et al. (2013) observe the exact opposite finding in the Korean market. If other emerging Asian markets react to foreign ratings similarly to the Korean market, this national difference in the relevance of American rating agencies could explain my finding by suggesting that the American agencies' rating methods evaluate Japanese credit better, and thereby fail to detect some factors of creditworthiness relevant in the other Asian markets. An alternative explanation could be offered by the finding of Yamori et al. (2006), suggesting that international agencies assign Japanese entities with lower credit ratings than the local rating agencies, which would then lead to higher CDS spread levels overall.

## 7. Conclusion

In this thesis, I have studied the CDS market's response to credit rating announcements. I focus my study on the Asian CDS markets during the period between October 1, 2007 and March 18, 2014 and construct a final uncontaminated sample of 373 Moody's and S&P issuer rating changes. As both, an entity's current credit rating, and the corresponding CDS spread, act as measures of credit risk, I question whether a credit rating change brigs new information to the CDS markets, or has the information about the change in an entity's creditworthiness already been incorporated in the CDS spreads before announcement. I hypothesize that, given the large presence of the international rating agencies, all rating actions generate a market reaction and that the markets react similarly to both Moody's and S&P ratings.

To test my hypotheses, I use traditional event study methods and regression analysis, using both joint (Moody's and S&P) and agency specific (Moody's or S&P) samples. First, I analyze the mean cumulative abnormal CDS spread changes, $\overline{CASC}_{\tau1,\tau2}$, during the event time window [-90, 90], also 90 business days before and after the rating announcement day zero. I test $\overline{CASC}_{\tau1,\tau2}$ with two univariate tests to analyze the statistical and economic significance of the spread changes during 9 shorter event windows around the rating announcement. Second, I estimate two pooled cross-sectional regression models, where I regress raw CDS spread changes on CDS index changes and add event time dummy variables to the model to analyze the market reaction close to the rating announcement as well as the difference in market reaction across rating agencies.

My analyses yield the following key results; first, my findings suggests that both upgrades and downgrades bring new information to the markets and thereby provide evidence for *H1*, as a significant market reaction on upgrade and downgrade announcements are observed in the Asian CDS markets. While the effect for upgrades is uniform between Moody's and S&P, the markets show no reaction to Moody's downgrades on announcement. Second, although rating announcements bring new information to the markets, the pending rating events are also largely anticipated by the market participants. CDS spreads start to incorporate new default risk information as early as 90 prior to a downgrade announcement and 30 days prior to an upgrade announcement, suggesting that markets adjust to the public information prior to the rating event and only react to the non-public information on announcement. Third, I find the markets to react strongly around rating events between investment- and non-investment grade, which suggests that the rating related investment thresholds generate abnormal changes in the CDS spreads.

Topics for further research include extending the sample of this thesis to consider rating reviews and outlooks by the American companies as well as the ratings of their Asian counterparts (e.g. Japanese R&I or Chinese Xinhua Far East). This would allow for a thorough comparison between the international and local credit rating agencies. Another topic to consider is to further test the volatile market reaction around rating changes close to the investment threshold grade, by observing trading volume changes in different markets and analyzing the reaction in trading volumes to a pending credit rating change.

# Appendix: List of reference entities in the final CDS data sample

**BAHRAIN** (1 entity)
Arab Banking Corporation BSC

**CHINA** (3 entities)
Agricultural Bank of China Ltd
Bank of China Ltd
Industrial and Commercial Bank of China Ltd

**HONG KONG** (22 entities)
Agile Property Holdings Ltd
Bank of East Asia Ltd
Cheung Kong Holdings Ltd
China Mobile Ltd
China Overseas Land & Investment Ltd
Citic Resources Holdings Ltd
CLP Holdings Ltd
CNOOC Ltd
ENN Energy Holdings Ltd
Greentown China Holdings Ltd
Henderson Land Development Co Ltd
Hopson Development Holdings Ltd
Hysan Development Co Ltd
Kerry Properties Ltd
Li & Fung Ltd
Road King Infrastructure Ltd
Shanghai Industrial Urban Development Group Ltd
Shimao Property Holdings Ltd
SRE Group Ltd
Swire Pacific Ltd
Titan Petrochemicals Group Ltd
Wharf Holdings Ltd

**INDIA** (12 entities)
Axis Bank Ltd
Bharti Airtel Ltd
Canara Bank Ltd
ICICI Bank Ltd
IDBI Bank Ltd
Indian Oil Corpn Ltd
NTPC Ltd
Power Finance Corporation Ltd
Reliance Industries Ltd
State Bank of India
Tata Motors Ltd
Tata Power Company Ltd

**INDONESIA** (3 entities)
Bank Negara Indonesia Persero Tbk PT
Indosat Tbk PT
XL Axiata Tbk PT

**ISRAEL** (1 entity)
Teva Pharmaceutical Industries Ltd

**JAPAN** (133 entities)
Acom Co Ltd
Aeon Co Ltd
AIFUL Corp
Aisin Seiki Co Ltd
Ajinomoto Co Inc
Aozora Bank Ltd
Asahi Glass Co Ltd
Asahi Group Holdings Ltd
Asahi Kasei Corp
Bank of Kyoto Ltd
Bank of Yokohama Ltd
Bridgestone Corp
Brother Industries Ltd
Canon Inc
Casio Computer Co Ltd
Central Glass Co Ltd
Central Japan Railway Co
Chubu Electric Power Co Inc
Chugoku Electric Power Co Inc
Citizen Holdings Co Ltd
Cosmo Oil Co Ltd
Credit Saison Co Ltd
Dai-ichi Life Insurance Co Ltd
Daikin Industries Ltd
Daiwa House Industry Co Ltd
Daiwa Securities Group Inc
Denso Corp
Dentsu Inc
DIC Corp
eAccess Ltd
East Japan Railway Co
Ebara Corp
Eisai Co Ltd
Fuji Electric Co Ltd
Fuji Heavy Industries Ltd
Fujitsu Ltd
Furukawa Electric Co Ltd
Hankyu Hanshin Holdings Inc
Hitachi Capital Corp
Hitachi Ltd
Hitachi Metals Ltd
Hokkaido Electric Power Co Inc
Hokuriku Electric Power Co
Honda Motor Co Ltd
IHI Corp
Isuzu Motors Ltd
Itochu Corp
Japan Airlines Co Ltd
Japan Real Estate Investment Corp
Japan Tobacco Inc
JFE Holdings Inc
Kajima Corp
Kansai Electric Power Co Inc
Kao Corp
Kawasaki Heavy Industries Ltd
Kawasaki Kisen Kaisha Ltd
KDDI Corp
Keikyu Corp
Keio Corp
Kintetsu Corp
Kirin Holdings Co Ltd
Kobe Steel Ltd
Komatsu Ltd
Kuraray Co Ltd
Kyushu Electric Power Co Inc
Makita Corp
Marubeni Corp
Marui Group Co Ltd
Mazda Motor Corp
Mitsubishi Chemical Holdings Corp
Mitsubishi Corp
Mitsubishi Electric Corp
Mitsubishi Estate Co Ltd
Mitsubishi Heavy Industries Ltd
Mitsubishi Materials Corp
Mitsubishi UFJ Financial Group Inc
Mitsui & Co Ltd
Mitsui Chemicals Inc
Mitsui Fudosan Co Ltd
Mitsui O.S.K. Lines Ltd
Mizuho Financial Group Inc
Nagoya Railroad Co Ltd
Nankai Electric Railway Co Ltd
NEC Corp
Nikon Corp
Nippon Electric Glass Co Ltd
Nippon Paper Industries Co Ltd
Nippon Sheet Glass Co Ltd
Nippon Steel & Sumitomo Metal Corp
Nippon Telegraph And Telephone Corp
Nippon Yusen KK
Nishimatsu Construction Co Ltd
Nissan Motor Co Ltd
Nomura Holdings Inc
NSK Ltd
NTT Docomo Inc
Obayashi Corp
Odakyu Electric Railway Co Ltd
Oji Holdings Corp
Okinawa Electric Power Co Inc
Omron Corp
Oriental Land Co Ltd
Orix Corp
Osaka Gas Co Ltd
Panasonic Corp
Pioneer Corp
Ricoh Co Ltd
Ricoh Leasing Co Ltd
Sapporo Holdings Ltd
Sharp Corp
Shinsei Bank Ltd
Shiseido Co Ltd
Softbank Corp
Sojitz Corp
Sony Corp
Sumitomo Chemical Co Ltd
Sumitomo Corp
Sumitomo Electric Industries Ltd
Suzuki Motor Corp
Taiheiyo Cement Corp
Tobu Railway Co Ltd
Tohoku Electric Power Co Inc
Tokyo Electric Power Co Inc
Tokyo Gas Co Ltd
Tokyu Corp
Toppan Printing Co Ltd
Toray Industries Inc
Toshiba Corp
Toyota Motor Corp
Toyota Tsusho Corp
West Japan Railway Co
Yamaha Motor Co Ltd
Yokogawa Electric Corp

**KAZAKHSTAN** (7 entities)
Alliance Bank AO
ATFBank AO
Bank TsentrKredit AO
BTA Bank AO
Halyk Bank AO
Kazkommertsbank AO
Razvedka Dobycha KazMunayGaz AO

**MALAYSIA** (7 entities)
Genting Bhd
IOI Corporation Bhd
Malayan Banking Bhd
MISC Bhd
Public Bank Bhd
Telekom Malaysia Bhd
Tenaga Nasional Bhd

**PHILIPPINES** (3 entities)
Globe Telecom Inc
Philippine Long Distance Telephone Co
Universal Robina Corp

**QATAR** (3 entities)
Commercial Bank of Qatar QSC
Doha Bank QSC
Ooredoo QSC

**SAUDI ARABIA** (3 entities)
Riyad Bank SJSC
Samba Financial Group
Saudi British Bank

**SINGAPORE** (11 entities)
Capitaland Ltd
China Fishery Group Ltd
DBS Bank Ltd
DBS Group Holdings Ltd
Flextronics International Ltd
Jardine Matheson Holdings Ltd
Noble Group Ltd
Oversea-Chinese Banking Corporation Ltd
Singapore Telecommunications Ltd
STATS ChipPAC Ltd
United Overseas Bank Ltd

**SOUTH KOREA** (17 entities)
Hyundai Motor Co
Industrial Bank of Korea
KCC Corp
KIA Motors Corp
Korea Electric Power Corp
Korea Gas Corp
KT Corp
LG Electronics Inc
LG Uplus Corp
POSCO
Samsung Electronics Co Ltd
Shinsegae Co Ltd
SK Broadband Co Ltd
SK Hynix Inc
SK Innovation Co Ltd
SK Telecom Co Ltd
Woori Finance Holdings Co Ltd

**TAIWAN** (4 entities)
Cathay Financial Holding Co Ltd
CTBC Financial Holding Co Ltd
Fubon Financial Holding Co Ltd
Wan Hai Lines Ltd

**THAILAND** (2 entities)
Bangkok Bank PCL
Thai Oil PCL

**TURKEY** (1 entity)
Turkiye Is Bankasi AS

**UNITED ARAB EMIRATES** (7 entities)
Abu Dhabi Commercial Bank PJSC
Abu Dhabi National Energy Co PJSC
Aldar Properties PJSC
DP World Ltd
Dubai Islamic Bank PJSC
Mashreqbank PSC
National Bank of Abu Dhabi PJSC

**TOTAL** (240 entities)

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