An Empirical Analysis of Changes in the Relative Timeliness of Issuer-Paid vs. Investor-Paid Ratings[†]

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We investigate the lead-lag relationships between issuer- and investor-paid credit rating agencies, in the aftermath of the regulatory reforms undertaken in the U.S. between 2002 and 2006 —including watch list inclusions and outlooks. First, we find that the lead effect of investor-paid over issuer-paid credit rating agencies has weakened: in recent years, causality has turned bi-directional. Second, when changes in outlooks are included, we find evidence of a less conservative behavior by issuer-paid agencies, when compared to their rating behavior. Third, stock prices manifest statistically significant abnormal reactions to downgrades of all agencies; however, abnormal negative returns are significantly higher for investor-paid downgrades. Our results support the hypothesis that when issuer-paid agencies have seen their market power threatened by tighter regulations, they have felt incentives to improve the quality and timeliness of their ratings. However, event studies show that markets still price stocks under the assumption that investor-paid rating actions carry superior information.

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1. Introduction

During the last decade the three big issuer-paid credit rating agencies (Fitch, Moody's, and Standard and Poor's, i.e., the long-time nationally recognized statistical rating organizations), have been exposed to major criticisms for their lack of timeliness in predicting imminent bankruptcies (see, e.g., Morgenson, 2008). A fact often invoked to support the need for tighter regulations is that the three leading credit rating agencies (CRAs) maintained investment-grade ratings only days before chapter 11 was filed by a number of bond issuers in notorious default cases such as WorldCom, Enron and, more recently, Lehman Brothers.[2] At the same time, a small niche of new and dynamic CRAs, paid only by investors (e.g., Egan & Jones Ratings and Rapid Ratings), have built a good reputation for accurate and economically valuable ratings.[3] As it has been widely discussed in the literature (see e.g., Cantor and Packer, 1994; Johnson, 2004; Langohr and Langohr, 2009), while issuer-paid CRAs extract fees directly from the issuers of bonds, investor-paid CRAs are only compensated by the final users of their ratings, such as institutional investors.

The prevailing explanation in the academic literature for the alleged failure of the classical NRSROs rating agencies to respond to investors' needs by providing timely revisions of their assessments of the creditworthiness of firms is that the conflicts of interest implicit in the compensation structure of issuer-paid CRAs would often (especially in the case of downgrades) advise them to change their ratings only when material and widely confirmed information becomes available, which is often well-after important events have occurred (see e.g., Johnson, 2004; Beaver et al., 2006).[4] Other authors have argued in favour of a relationship-based explanation, whereby rated firms with a long rating history would receive better ratings than

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[2] Moody's (S&P) put Lehman Brothers' "A2" ("A") rating on watch on September 10, 2008 (September 09, 2008). On the day Lehman announced its bankruptcy filing (September 15, 2008), Moody's downgraded Lehman by ten notches to B3 (non-investment grade) and placed it on review for possible further downgrades. Similarly, S&P downgraded Lehman on the same day from A to SD (selective default).

[3] Although it has maintained its original investor-paid fee model, Egan & Jones has gained NRSRO status on December 21, 2007. Therefore, because our empirical analysis focuses on the 1997-2007 sample, the key distinction is between the issuer- vs. investor-paid models, and not the opposition NRSROs vs. non-NRSROs.

[4] It has been argued that reputational concerns would discourage CRAs from engaging in short-term opportunistic behaviours in spite of the issuer-paid fee model: the on-going value of the business would depend so strongly on continued investor confidence in the reliability of ratings, that no fee could be important enough to jeopardize it. For However a number of academic papers have reported evidence that the business model affects on average the *level* of ratings, see e.g., Jiang et al. (2012) and Xia (2010). Our paper focuses on the effects of such model on the (relative) *timeliness or rating changes*.

firms with shorter histories (see e.g., Mählmann, 2011). Finally, a few papers (e.g., Ahmed, 2010; Becker and Milbourn, 2011; Ekins et al., 2011; Opp et al. , 2012) have emphasized that the oligopolistic structure of the rating industry represents a complementary reason for the lack of incentives to respond to the demand of promptly updated ratings (see Bolton et al., 2012; Doherty et al., 2012). Although the fine details of these alternative explanations differ (see the discussion in Mathis, McAndrews and Rochet, 2009; Bar-Isaac and Shapiro, 2013), they all boil down to statements that to traditional, issuer-paid NRSROs the expected reputational cost of delaying the revision of a rating may be inferior to the expected benefits, in terms of profitable business with bond issuers.

As a result of this debate on the alleged shortcomings of the prevailing rating business model, since 2002 both the U.S. Congress and the Security and Exchange Commission (SEC) have held hearings on the certification process of the Nationally Recognized Statistical Ratings Organizations (NRSROs), and have discussed the possibility of overhauling the regulatory framework that applies to CRAs, in order to increase transparency and promote competition (SEC 2002; SEC 2003; SEC 2007; SEC 2008). These discussions have led to the appearance of specific references to CRAs in the Sarbanes-Oxley Act (2002), to the Credit Rating Agency Reform Act (2006), and more recently to the Dodd-Frank Wall Street Reform Act (2010). Moreover, since 2009 in Europe there has been a debate concerning the opportunity to tightly regulate CRAs (see e.g., Katz et al., 2009), while the European Central Bank has identified a list of credit assessment sources accepted within the Euro-system.

Because the alleged lack of timeliness has been the most widely debated accusation moved to CRAs in the aftermath of the 2002-2003 SEC hearings (see Altman and Rijken, 2004; Cantor and Mann, 2006; Liu et al., 2010), our paper investigates the evolution over time of the comparative (relative) timeliness of investor-paid and issuer-paid agencies. As shown by Beaver, Shakespeare and Soliman (2006, henceforth BSS), credit rating changes by issuer-paid CRAs were initially found to be led by changes by investor-paid agencies. However, recent changes in regulations and increased investor scrutiny mean that a late response to any deterioration in the creditworthiness of a firm has had an increasing reputational cost effect. Consequently, because the regulatory reforms and increased "political" pressures are likely to have modified the relative incentives of CRAs, affecting the likelihood of reputational losses exceeding the costs of currying favour with bond issuers, we expect issuer-paid agencies to have improved their rating timeliness. Therefore, with reference to corporate bond ratings, in this paper we perform formal statistical tests of structural change in the lead-lag relationship between issuer- and investor-paid CRAs.

More generally, we present the first systematic comparison of the timeliness of the rating "actions" performed by the leading investor-paid CRA, Egan & Jones Ratings (EJR), and three issuer-paid agencies, i.e., Fitch, Moody's, and S&P. Given our goals, we extend the sample of rating actions in the earlier literature, to a full decade (July 1997 to December 2007) that includes a number of regulatory reforms that are likely to have affected the CRAs' incentives. Moreover, our notion of rating "action" is not limited to outright rating revisions (as in Johnson, 2004, BSS, 2006, Cheng and Neamtiu, 2009), but encompasses the early warning system represented by watch list inclusions and changes in outlooks that are routinely implemented by both issuer- and investor-paid agencies. We incorporate changes in outlooks because issuer-paid agencies may behave less conservatively in changing outlooks than ratings, since rating changes are directly tied to their regulatory status (Bannier et al., 2010).

We start by establishing lead-lag relationships involving the two kinds of rating agencies—issuer- vs. investor-paid CRAs. To gain power, we use two alternative methods, a Granger causality analysis and an ordered-probit framework. The intuition behind this range of econometric testing frameworks is that, if prior rating actions of one CRA help to forecast the subsequent actions of another agency but the same is not true the other way around, one can claim to have isolated a one-way lead-lag relationship. Ordered probit models have been widely used to model rating changes due to their ordinal nature since Ederington (1985) and has been more recently exploited by Güttler and Wahrenburg (2007) and Alsakka and Gwilym (2010). This type of econometric approach not only allows us to jointly analyze relative timeliness across downgrades and upgrades, but also to assess the differences in the probabilities of rating actions by one agency, based on the magnitude of prior actions by one or more other CRAs. Because we conjecture that the structure of lead-lag relationships has changed over time, we use a Chow-style breakpoint approach to test for instability in the parameters of the model. Moreover, we perform a set of event studies to test whether market reactions are in line with the results of our lead-lag analysis and to assess the economic value, if any, of the change in timeliness between issuer- and investor-paid CRAs.

Our results show that the unidirectional lead-lag relationships for downgrades reported by previous studies (e.g., Johnson, 2004; BSS, 2006), in which investor-paid CRA actions Granger-cause the actions of issuer-paid agencies, turn into bi-directional relationships when we expand the period of analysis beyond June 2002, i.e., beyond the original sample period in BSS (2006). Downgrades by Fitch, Moody's, and S&P significantly increase the probability of EJR downgrades over a time window of up to four months. On the other hand, downgrades by EJR significantly increase the probability of downgrades by the other agencies over a time window of up to six months. As far as upgrade actions are concerned, the results suggest instead that EJR is still a leader and most likely to be the first mover. Upgrades by EJR significantly increase the probability of subsequent upgrades by Fitch, Moody's, and S&P in the following six months.

Next, we incorporate changes in outlooks and watch lists for all rating agencies. The bi-directional lead-lag relationship between issuer- and investor-paid agencies is preserved for downgrades but it changes for upgrades. For downgrades, there is overwhelming evidence of a bi-directional lead-lag relation, albeit EJR leads for up to six months while issuer-paid agencies for about four to five months. For upgrades, EJR leads Fitch and S&P, but is (weakly) led by Moody's. We also find evidence that issuer-paid changes in outlooks increase the probability of the investor-paid agency's changes in ratings, which lends support to a less-conservative behaviour of issuer-paid agencies in changes in outlooks than in ratings.

However, the watershed of June 2002 has been exogenously imposed in the empirical results in the first part of the paper discussed so far, a natural choice advised by the fact that 2002 was the end of the sample period in BSS (2006). To remove any undue influence of such a choice of an exogenous break date, we also test for instability in the parameters of our dynamic ordered probit. We find evidence of breaks in the parameters for all pairs of CRAs involving EJR and each of the issuer-paid agencies. Our results show that, even though the dates of the breaks are different across pairs, all of the breaks that we manage to isolate have in common that they were preceded by the changes in regulations recalled above.

Because the first two steps of our research design are exquisitely statistical in their nature, one may wonder what the economic effects of these breaks in the lead-lag relationships have been. As a last step in our investigation we therefore examine the link between abnormal stock returns and ratings, outlooks and watch list inclusions. Here, we separate rating actions into two categories: *unconditional* actions, when the rating change is not preceded by actions by any of the other agencies, over an event window of one month; *conditional* actions, if there has been a previous rating decision in the same direction (i.e., an upgrade or a downgrade) by another CRA. For unconditional downgrades, abnormal negative returns are found for all CRAs, but these turn out to be the largest for downgrades issued by the investor-paid agency. For unconditional upgrades, abnormal positive returns only appear in the case of EJR. For conditional downgrades, we observe significant abnormal returns only in the case of the investor-paid agency. This suggests that EJR's downgrades that follow an earlier downgrade by an issuer-paid CRA are considered to carry additional information not already discounted in previous downgrades. In the case of conditional upgrades, we do not find significant abnormal positive returns as a result of the actions by any of the agencies. Additional event studies separately performed on the pre-2002 and post-2002 samples indicate that while prior to 2002, abnormal stock reactions to conditional actions (in particular, downgrades) are significantly greater than reactions to unconditional changes when EJR follows earlier actions undertaken by investor-paid CRAs, the differences in abnormal returns triggered by conditional vs. unconditional actions do not yield strong evidence in the post-2002 sample.

The literature that compares issuer- and investor-paid CRAs has only recently emerged and is still fairly sparse. Johnson (2004) finds that EJR leads Standard & Poor's ratings' changes around the investment grade boundary. BSS (2006) and Strobl and Xia (2012) have extended Johnson's analysis to the entire ratings ladder but only compare Moody's, S&P, and EJR rating changes. They all find a considerable lead effect of EJR over Moody's and S&P.[5]

To our knowledge, our paper is the first examination of the instability in the parameters affecting the differences in timeliness and in the implied economic value between the most celebrated investor-paid agency (EJR) and the big three, issuer-paid CRAs. Prior research has only compared single pairs of investor-paid and issuer-paid CRAs (Johnson, 2004), with emphasis on the Moody's-EJR pair (BSS, 2006; Bruno et al., 2012) or specific thresholds such as the investment grade (Johnson, 2004). Additionally, we incorporate data from the early

[5] A literature has estimated leads and lags among issuer-paid CRAs only: Jewell and Livingston (1999) compare the ratings of Fitch, Moody's and S&P finding that the sample of companies rated by Fitch enjoys higher ratings than those not rated by Fitch when Moody's and S&P do not agree; Güttler (2011) examines the intensity of Moody's and S&P's actions conditional on each other, and finds that Moody's lags S&P. Milidonis (2013) finds that in the insurance sector, EJR leads bond and financial strength ratings by Fitch and S&P.

warning system whose changes often precede rating actions, i.e., watch lists or outlooks, using the approach by Gande and Parsley (2005). Finally, compared to the existing literature, we expand our sample beyond 2002 to include data up to December 2007.

The three most closely related papers are Bruno, Cornaggia, and Cornaggia (2012), Cheng and Neamtiu (2009), and Jorion, Liu, and Shi (2005).[6] It is useful to discuss them in isolation to emphasize our contributions. Bruno et al. have exploited the change in EJR's status to NRSRO in December 2007 to test the importance of SEC certification relative to the compensation structure in determining rating behaviour. They test the null hypothesis that—if the timeliness and accuracy of EJR's ratings are a function of the differential compensation structure—one should expect the differences in rating properties to persist following EJR's NRSRO designation; if on the contrary market participants simply change how *they* use EJR's credit ratings after the ratings became sanctioned for regulatory compliance, this change should prompt EJR to produce ratings with informational properties similar to those produced by Moody's. Bruno et al. find that EJR's ratings remained timelier than Moody's in the aftermath of December 2007. However, the analysis in Bruno et al. (2012) mostly concerns the period after December 2007, it focuses on EJR taking Moody's behaviour as given, and assumes the absence of any breaks in the distorted incentives by the traditional issuer-paid NRSROs before 2007. On the contrary, our paper tests the hypothesis that the three issuer-paid CRAs may have been affected by the change in regulations between 2002 and 2006, and our data extend to outlooks and watch list inclusions.[7]

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[6] Using S&P ratings, Blume et al. (1998) have studied whether investment-grade standards had changed between 1978 and 1995. They concluded that the deterioration in the credit quality of US firms was driven by the stricter standards employed by S&P. With reference to Moody's ratings, Lucas and Lonski (1992) had found that the ratio of long-term downgrades to upgrades deteriorated from an 1.17 average in the 1970s to 2.17 in the 1980s, and reached a record 4.93 in 1990. Alp (2013) finds that until 2002, rating standards above investment grade are stricter than those below investment grade, but after 2002, rating standards become tighter for both categories. However, none of these papers had formally examined the presence of structural breaks in the relative timeliness of issuer- vs. investor-paid ratings.

[7] Kisgen and Strahan (2010) have studied the existence of a structural break in the rating policies of Dominion Bond Rating Service (DBRS) after it received the NRSRO designation in 2003. They report that after 2003, the market relied on DBRS ratings more for regulatory compliance, but not for information about credit quality. However, like the Big Three, DBRS is an issuer-paid CRA and there is no evidence to suggest that its ratings were ever more timely or accurate than those of the Big 3 before 2003. With respect to the level of ratings, Jiang et al. (2012) find that when S&P switched from an investor-paid to an issuer-paid compensation model, an inflation in their ratings was observed, especially in the sample of firms where the conflict of interest related to compensation would be stronger.

Cheng and Neamtiu (2009) investigate whether and how the three historical, issuer-paid NRSROs have changed their rating methods in response to the recent increase in regulatory pressure, including the 2002 reforms and the proposals that have subsequently led to the Credit Rating Agency Reform Act of 2006. Using data on defaults and ratings for a 1997-2005 sample, they find that the three issuer-paid CRAs downgraded defaulting bonds earlier and assigned significantly closer-to-default (i.e., more timely) ratings in the post-2002 reform. However, their focus is exclusively on potential changes in the behaviour of Fitch, Moody's, and S&P, without any comparison with investor-paid CRAs. In our paper we explicitly leverage our tests on the relative timeliness and informativeness of issuer- vs. investor-paid.

Finally, Jorion et al. (2005) test for breaks in the information content of credit ratings, even though their focus is on the effects of the Regulation Fair Disclosure (Reg FD) implemented by the SEC in 2000.[8] Their hypothesis is that, if rating agencies obtained almost exclusive access to selective information that could no longer be disclosed to other agents, the information content of, and the stock price reaction to, ratings should have grown after Reg FD came into effect. Using the post-FD period after November 2000 and up to December 2002 as a post-break sample for the three big issuer-paid NRSROs, Jorion et al. find that, relative to the pre-Reg FD period, downgrades were associated with greater falls in stock prices, and upgrades were associated with greater increases. Although we also pursue the idea that the quality and statistical properties of ratings may have been subject to structural instability caused by a change in regulations, our paper focuses on a later set of regulatory reforms, uses a different battery of tests compared to Jorion et al.'s event studies, and tests hypotheses concerning the relative timeliness and informativeness of issuer- vs. investor-paid CRAs.

The remainder of the paper is organized as follows. The next section provides background information on CRAs and the recent evolution of the regulatory framework that underlies the key hypotheses tested in the paper. Section 3 presents the hypotheses development and the methodology. Section 4 describes our data and sample selection criteria. Section 5 reports our empirical results. Section 6 concludes.

[8] Reg FD requires that U.S. public companies that intentionally disclose material, non public information to a select group also disclose it simultaneously to the public.

2. Institutional Background

Credit rating agencies are intermediaries that help reduce the information asymmetries among investors and firms. Originally, CRAs were funding their costs by selling their public information-based rating manuals. The Great Depression triggered a number of changes in financial regulation that forever altered the relationship between the U.S. bond market and the CRAs. In 1936, eager to encourage banks to invest only in safe bonds, regulators (the Comptroller of the Currency, even though under the Federal Reserve Act, the same regulations would also govern all Federal Reserve member banks) prohibited them from investing in speculative securities, as determined by the rating agencies that existed at the time (Fitch, Moody's, and Standard & Poor's). From this moment on, banks were obliged to rely on the judgements of certain "recognized" agencies, whose opinions had implicitly acquired the force of law. Later, in 1975, this recognized category became known as the NRSRO group, when the SEC imposed minimum capital requirements on broker-dealers, based on their ratings. Concurrently, the agencies with this certification changed their business model, funding themselves by selling their ratings predominantly to issuers instead of investors (see White, 2010, for a discussion of the key driving motives). Since then, two different types of agencies have existed: those paid by the investors and the NRSROs, all of which have been mostly funded by issuers' fees. Over time, regulators and capital markets have come to increasingly rely on NRSRO ratings. Today, these ratings are not only used for valuation purposes but are also featured in the capital adequacy rules enforced over banks, in federal and state legislations, and also extensively in financial contracts (see Asquith et al., 2005). On the contrary, the ratings issued by investor-funded CRAs are only used for valuation purposes. Therefore, once a CRA is granted NRSRO status, its influence increases substantially because any new public debt issuance has to be rated by at least one NRSRO. Nevertheless, until 2007 there were no clear regulatory requirements that a CRA had to meet in order to qualify as a NRSRO: It was simply stated that the agency's ratings should be widely used (i.e., an agency had to be "nationally recognized", based on the vague wording of the 1975 SEC's regulation) and considered reliable by their users. Until 2003 there were only three NRSROs: Fitch, Moody's, and S&P.[9] However, at the time of our writing, ten CRAs have been granted NRSRO

[9] During the 25 years that followed the SEC's 1975 creation of the NSRO category, the SEC approved only four firms as additional NRSROs: Duff & Phelps in 1982; McCarthy, Crisanti & Maffei in 1983; IBCA in 1991; and

status.[10] Although the compensation regimes (issuer- vs. investor-paid) now vary among the existing NRSROs, the three most important agencies (Fitch, Moody's and S&P) are still entirely paid by bond issuers.

Since the financial debacles of Enron (December 2, 2001) and WorldCom (July 21, 2002), there has been sustained pressure from the U.S. Senate and the House of Representatives to review the regulatory system that applies to CRAs. The main concerns were the conflict of interests plaguing issuer-paid CRAs, an alleged lack of competence, and a potential deficiency of regulatory oversight in what has been for a long time a self-regulated industry. The process of reform started in 2002 with the Sarbanes-Oxley (SOX) Act, which required the SEC to scrutinize and monitor the role of CRAs.[11] Additionally, on September 29, 2006, the Credit Rating Agency Reform Act was passed by the U.S. Congress to enhance transparency and competition in ratings. The Rating Agency Act also gave the SEC the authority to implement registration, recordkeeping, financial reporting and oversight rules with respect to registered CRAs. The new rules became effective on June 26, 2007. Finally, in 2011 the Dodd-Frank Act has mandated further measures from the SEC regarding NRSROs, related to NRSROs' annual reporting of methodologies and rating assumptions, and potential conflicts of interest.

3. Hypotheses Development and Empirical Methodology

3.1 Hypotheses

The main difference between the two types of CRAs lies in the source of the compensation they receive for their services. While a first, newcomer group is composed of CRAs that are paid by investors and that only respond to the final users of the ratings, a traditional block is composed of CRAs that are paid by the issuers of bonds. However, all CRAs naturally have two audiences. On the one hand, the issuers (that pay to be rated) are interested in receiving good

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Thomson BankWatch in 1992. However, mergers among the entrants and with Fitch caused the number of NRSROs to return to the original number of three before the end of 2000.

[10] These are A.M. Best Company, Inc., DBRS Inc., Egan-Jones Rating Company, Fitch, Inc., Japan Credit Rating Agency, Ltd., Kroll Bond Rating Agency, Moody's Investors Service, Inc., Morningstar Credit Ratings, LLC, Rating and Investment Information, Inc., Standard & Poor's Ratings Services.

[11] The SOX of July 2002 required the SEC to investigate the role of CRAs in securities' markets with respect to (1) information flow in the credit rating process; (2) potential conflicts of interest; (3) alleged anticompetitive or unfair practices, and (4) potential regulatory barriers to entry into the credit rating business (see SEC, 2003).

ratings and enjoying a low cost of debt as a result. On the other hand, investors and regulators look at these ratings in order to make their decisions. The historical behaviour of agencies that are paid by the issuers suggests that they may be more conservative in changing their ratings, in the sense that they would need to observe substantial evidence of a deterioration (or an improvement) in the financial health of a company before they would downgrade (upgrade), see BSS (2006). Therefore, although this does not represent a formal hypothesis tested in what follows (because a literature exists that has documented this finding), our paper is generally concerned with the fact that investor-paid rating agencies' changes lead the changes in ratings produced by issuer-paid agencies.[12] As a result, in Section 4 we report results that corroborate this lead-lag relationship but we organize the conjectures that follow in the form of a set of hypotheses concerning the *changes*, if any, in strength of such relationships observed over time (after July 2002) as well as their economic effects/value.

On the one hand, a number of papers (e.g., Watts, 2003) have imputed the tendency of issuer-paid CRAs to be conservative and sluggish to their implicit regulatory responsibilities that still characterize the U.S. legislation even after the 2002 and 2006 reforms and because these responsibilities mostly (or only) relate to official rating changes, we expect any differential timeliness to affect much more ratings than they do with outlooks and inclusions in watch lists. On the other hand, one contribution of our paper consists of the extension of standard results concerning official rating changes to data also concerning outlook revisions and watch list inclusions. It is therefore important to gain a preliminary grasp for the state of any lead-lag relationships concerning these categories of CRA actions, before proceeding any further. This leads to our first hypothesis:

H1: For both upgrades and downgrades, the lead-lag relationships that link the changes in ratings by investor-paid to issuer-paid credit rating agencies are stronger than those that link the outlooks and watch list inclusions by investor- to issuer-paid agencies.

Equivalently, because a number of laws, statutes, and contracts (especially at the investment grade boundary, see Asquith et al., 2005) attribute a special role to NRSRO-originated ratings but there are many less (or no) provisions that give the same special status to NRSRO-

[12] In this perspective, we revisit the hypothesis in BSS2006 using a longer sample period characterized by several regulatory changes, and also employ a larger sample of issuer-paid CRAs (i.e. we add Fitch and S&P to BSS analysis that was limited to Moody's vs. EJR).

originated outlooks and watch list inclusions, given that the NRSRO group of CRAs has been and still is dominated by the issuer-paid business model, we expect issuer-paid CRAs to be able to afford an inferior conservative bias when it comes to changes in outlooks and watch lists.[13] To our knowledge, this specific hypothesis concerning the relative timeliness including outlooks and watch lists has not been tested before.

However, H1 does not exploit in any way the recent overhaul of the regulations concerning CRAs to foster our understanding of their incentives and operational mechanisms. On the one hand, as we have summarized in section 2, a series of changes in regulations over the recent years (possibly to also include the Dodd-Frank Act in 2011) has made it increasingly costly for issuer-paid CRAs to delay making changes. On the other hand, the literature has clearly shown that, with reference to 1999-2002 data, investor-paid CRAs (e.g., EJR in BSS, 2006) lead issuer-paid CRAs by up to 6 months for upgrades and by 1-4 months for downgrades and that issuer-paid downgrades fail to Granger cause investor-paid CRA downgrades. However, the increased legislative and regulatory pressure on all CRAs to increase the timeliness and accuracy of their ratings is expected to have reduced this previously ascertained "distance" in favour of investor-paid over traditional issuer-paid CRAs. Equivalently, we expect issuer-paid CRAs to have modified their behaviour to increase their relative timeliness when compared to investor-paid CRAs (which of course, does not rule out ex-ante that the former may have progressively overtaken the latter).

H2: The lead-lag relationship between investor- and issuer-paid rating agencies has changed after 2002, and the effects of such a break ought to be visible by the end of 2006, when a further wave of reforms was enacted; this effect is likely to be stronger in the case of downgrades.

Equivalently, the post-2002 (but pre-NRSRO status acquisition by EJR) regulatory changes may have affected Moody's, S&P, and Fitch more than EJR because during our sample the Big Three CRAs were issuing ratings carrying legal and contractual value and this was not the case for EJR that was simply serving the interests of investors. In fact, even leaving aside the

[13] It may be objected that some degree of conservatism should be expected also of outlooks and watch lists because NRSROs will appreciate these are signals of potential, future rating changes. A few papers (see Ellul et al., 2011) have emphasized how hasty downgrades may trigger self-fulfilling market reactions due to the role played by NRSRO ratings (e.g., broker-dealers use them to determine the amount of collateral to hold against derivatives exposure), thus making NRSROs cautious. However, one may counter the fact that—especially when pressured to improve their timeliness and accuracy (see Cheng and Neamtiu, 2009)—NRSROs may actively use outlooks and watch lists to weaken the perception that they may be deficient.

differential compensation schemes between the traditional NRSROs and EJR, while the former function requires of ratings to be stable over time and advises a degree of prudence in the case of downgrades that may contrast with the objectives of the wave of reforms concerning the industry that has occurred between 2002 and 2006, the latter function should have been affected to a latter extent. Although we do not formalize this aspect in a separate proposition, H1 and H2 also jointly imply that the differential timeliness of investor- vs. issuer-paid CRAs may have faded to zero around the middle of the past decade as far as outlooks and watch lists are concerned.

Finally, the majority of the investors have probably realized that rating changes by issuer-paid agencies are not as timely as those by investor-paid CRAs. Even though our H2 implies that this situation has changed over time, it remains important to assess whether there is any significant difference in the reactions of the stock market to the rating changes of the two types of CRAs. Moreover, if there is a difference, we want to determine whether such differential stock price reactions may generate any predictable economic value that an investor may exploit. BSS (2006) concluded that investor-paid (EJR) rating upgrades (downgrades) have a significantly larger positive (negative) contemporaneous abnormal return than issuer-paid rating changes (Moody's) do.[14] The contribution of our paper consists in differentiating between two types of rating changes: *Unconditional* changes are those for which, in a predefined period prior to the change *t*, there are no other rating-related events or announcements. In these cases, the reaction of the stock market at time *t* may be imputed to the rating change announcement of the agency.[15] *Conditional* rating changes are changes that produce effects that only compound over earlier rating-related events or announcements within a fixed period of time before time *t*. In particular, given a pair of agencies (one issuer- vs. one investor-paid), conditional changes are those advertised by one CRA that are preceded by a rating-event in the same direction by the other CRA. Here a possible intuition is that investors may interpret the second rating change as a reinforcement of the news given by the first

[14] There is an early literature that has studied the effect of rating changes on abnormal stock returns using pre-2003 data, when the market shares of issuer-paid NRSROs were close to 100%. Holthausen and Leftwich (1986) found that rating changes by Moody's are information events but that abnormal returns result from downgrades but not for upgrades. Also Griffin and Sanvicente (1982) and Dichev and Piotroski (2001) conclude that bond rating changes affect common stock prices, contrary to the evidence in Pinches and Singleton (1978).

[15] This concept echoes Holthausen and Leftwich's (1986) and Stickel's (1986) definition of "clean" rating announcements, although their analysis focuses on the contaminating effects of the release of firm-specific information as reported in the Wall Street Journal.

change and, therefore, their reaction may be stronger than in the case of an unconditional change (see also Halek and Eckles, 2010, with reference to insurance ratings). This is the reason of our requirement that when multiple rating-related events occur over a short period of time and are used in our empirical tests, these should all point in the same direction, carrying news of similar "sign" to investors. Given these definitions, our two hypotheses regarding market reactions to CRA actions are:

H3: Stock markets display a significantly greater abnormal reaction to the unconditional rating changes issued by investor-paid agencies than they do for those by issuer-paid agencies.

H4: Abnormal stock market reactions to conditional rating changes will be significantly stronger than reactions to unconditional rating changes.

Notice that H4 refrains from ranking stock market responses to issuer- vs. investor-paid conditional rating changes. This is because conditional changes often imply—especially in the last part of our sample, as a result of the elevated activism of investor-paid CRAs—a strong tendency of rating events determined by both types of CRAs to mix, making a distinction between "pure" investor-paid and pure issuer-paid conditional rating events rather subtle and essentially non informative.

Finally, the same logic that has led us to write and test H2 supports a test for breaks in the reaction of stock markets to rating changes:

H5: Abnormal stock market reaction differentials between unconditional and conditional rating changes have declined after 2002.

We expect this empirical result to hold because of the regulatory reforms having affected the business model of NRSROs. Assuming the reforms were successful over time, NRSRO actions should have become increasingly informative, thus making all CRA actions on average more informative, including investor-paid CRA actions. As a result, the differential market reaction to unconditional vs. conditional actions should have declined, the investors being less "needy" of conditional reassurances from a range of CRAs before implementing their trading strategies. Moreover, if the actions of issuer-paid CRAs became increasingly informative, this should have also reduced the spread between abnormal reactions to the unconditional rating changes by investor-paid agencies vs. those by issuer-paid ones.

3.2 Empirical Methodology

To test the hypotheses above, we use four different tests: (1) we investigate (relative) timeliness using a Granger causality test; (2) we estimate a range of ordered probit models to test whether, in pair-wise comparison, upgrades and downgrades in ratings (and outlooks) by one CRA predict—or are predicted by—the probability of another CRA's actions; (3) we use a Chow-type framework to test for instability in the parameters of the ordered probit model and infer whether any breaks in the relationships of interest have occurred as a response to the changes in legislation; (4) we perform event studies to assess the stock market reaction to conditional and unconditional rating changes on the day of the event.

3.2.1 VAR-Based Granger Causality Tests

Hypotheses 1-2 concern the relative timeliness of the rating changes of different CRAs. All of our analyses compare one investor-paid rating agency (EJR), with each one of the main issuer-paid agencies (Fitch, Moody's, and Standard & Poor's). For each of the three pairs, we test whether the investor-paid CRA leads the issuer-paid one. To do so, we rely on standard Granger causality tests (Granger, 1969). This methodology analyses separately the lead-lag relationships of upgrades and downgrades for the sample of changes in ratings, and for the complete sample of changes in ratings and outlooks. We say that one CRA Granger-causes the other if the first agency's rating changes help to predict those of the second (i.e., the direction of the temporal relationship matters). For each pair of agencies, we estimate eight logistic regressions: two for upgrades in ratings and two for downgrades in ratings. Then, we also incorporate changes in outlooks in the sample and re-estimate the same four logistic regressions. To save space, below we explain the methodology using changes in ratings (i.e., equations 1-4), even though in later empirical tests we also incorporate changes in outlooks.

The four dependent variables used to test lead-lag relationships are indicator variables which take the value of one if there has been a change at time *t*, and zero otherwise. Two separate regressions are estimated for upgrades and two for downgrades. The unrestricted models include lagged values of the dependent variable up to 6 months, as well as the other agency's lagged indicator variables for another 6 lags:

$$AgencyADown_t = \alpha_0 + \sum_{j=1}^{6} \alpha_j AgencyADown_{t-j} + \sum_{j=1}^{6} \beta_j AgencyBDown_{t-j} + \varepsilon_t \quad (1)$$

$$AgencyBDown_t = \alpha_0 + \sum_{j=1}^{6} \alpha_j AgencyADown_{t-j} + \sum_{j=1}^{6} \beta_j AgencyBDown_{t-j} + \varepsilon_t \quad (2)$$

$$AgencyAUp_t = \alpha_0 + \sum_{j=1}^{6} \alpha_j AgencyAUp_{t-j} + \sum_{j=1}^{6} \beta_j AgencyBUp_{t-j} + \varepsilon_t \quad (3)$$

$$AgencyBUp_t = \alpha_0 + \sum_{j=1}^{6} \alpha_j AgencyAUp_{t-j} + \sum_{j=1}^{6} \beta_j AgencyBUp_{t-j} + \varepsilon_t, \quad (4)$$

where, by convention we set agency A to correspond to an investor-paid CRA and agency B is one of the issuer-paid CRAs. The restricted models only incorporate lagged values of the dependent variable as explanatory variables. The logistic models in equations (1)-(4) can be used to test relative timeliness in the following way. If agency A Granger-causes agency B to perform a rating change, then we expect the (or at least some of the) α coefficients to be positive and statistically significant in equations (2) and (4) while the β coefficients are not significant. At the same time, we expect the β coefficients in equations (1) and (3) not to be significant, i.e., investor-paid CRAs lead issuer-paid agencies but not the reverse. On the other hand, if agency B Granger-causes agency A, we expect the (or at least some of the) β coefficients to be positive and significant in equations (1) and (3). In various forms, our hypotheses 1-2 state that investor-paid CRAs should lead issuer-paid CRAs. We test for Granger causality by calculating a log-likelihood ratio test (LRT) that compares the explanatory power of the restricted vs. the unrestricted models, where in the restricted model the explanatory variables are only the one- through six-month lagged values of the dependent variable.[16] For example, the restricted versions of (1) and (3) are:

$$AgencyADown_t = \alpha_0 + \sum_{j=1}^{6} \alpha_j AgencyADown_{t-j} + \varepsilon_t \quad (5)$$

[16]The log-likelihood ratio statistic is calculated as $LRT = -2(L(\theta_0|x) - L(\theta_1|x))$, where $L(\theta_0|x)$ corresponds to the log-likelihood of the restricted model and $L(\theta_1|x)$ is the log-likelihood of the unrestricted model. The probability distribution of the statistic is approximately chi-square with degrees of freedom equal to the difference between the degrees of freedom of the unrestricted and the restricted models.

$$AgencyAUp_t = \alpha_0 + \sum_{j=1}^{6} \alpha_j AgencyAUp_{t-j} + \varepsilon_t \qquad (6)$$

Statistical significance at conventional size levels (i.e., the probability of type I error of incorrectly rejecting) of the LRT indicates that the additional explanatory variables in the unrestricted model (i.e., lagged values of the potentially leading CRA), Granger-cause changes in the dependent variable, because they improve over the log-likelihood value of the restricted model more than what may be attributed to pure chance.

3.2.2 Ordered Probit-Based Causality Tests

There is a second Granger-like causality method that we use to assess potential lead-lag relationships between pairs of CRAs. Following Güttler and Wahrenburg (2007) and Alsakka and Gwilym (2010), we use an ordered probit model that duly accounts for the discrete and ordinal nature of rating changes.[17] To examine whether the lead-lag relationships between CRAs differ with regard to upgrades and downgrades, the following models are estimated with agency A as the follower and agency B as a potential leader:

$$\Delta R_{i,t}^{*A} = \sum_{h=1}^{6} \beta_h^1 D\_up_{i,t-h}^{B} + \sum_{h=1}^{6} \beta_h^2 D\_dw_{i,t-h}^{B} + \varepsilon_i;\ \varepsilon_i \sim N(0,1) \qquad (7)$$

$$\Delta R_{i,t}^{*B} = \sum_{h=1}^{6} \beta_h^1 D\_up_{i,t-h}^{A} + \sum_{h=1}^{6} \beta_h^2 D\_dw_{i,t-h}^{A} + \varepsilon_i;\ \varepsilon_i \sim N(0,1) \qquad (8)$$

where $\Delta R_{i,t}^{*}$ is an unobserved latent variable linked to the observed ordinal response categories $\Delta R_{i,t}$, which refer to a rating change by agency A in equation (7) or by agency B in equation (8) for firm *i* on day *t*. As a first step, to identify changes in ratings, as common in the literature, we transform the letter ratings of all CRAs into a numeric credit rating (CR) scale (see Appendix A). We employ four different classes of rating changes: ≤-2, -1, +1, ≥+2, that is, a downgrade of two or more notches, a downgrade of one notch, an upgrade of one notch, and an upgrade of two or more notches, respectively.[18]

[17] Only formal probit (or tobit) methods provide consistent estimators for categorical variables, as discussed by Alsakka and Gwilym (2010).

[18] Upgrades and downgrades in excess of 2 two notches are sufficiently rare to make this four-bin ordered probit the largest estimable model with our data; see section 4 for additional details.

The independent variables $D\_up_{i,h}$ ($D\_dw_{i,h}$) correspond to dummy variables that take the value of one if a senior unsecured bond was upgraded (downgraded) by the potential leading agency within a predefined window of time h. Here h=1 refers to the interval 1-30 days, h=2 to 31-60 days, h=3 to 61-90 days, h=4 to 91-120 days, h=5 to 121-150 days, and h=6 to the interval 151-180 days, prior to the change in rating of firm i at time t, by the potential follower agency; the dummies are set to zero otherwise.

The unobserved latent variable $\Delta R^{*}_{i,t}$ is linked to our predefined and observed response rating change categories by the following measurement model:

$$\Delta R_{i,t} = \begin{cases} -2 \ (i.e. \, downgrade \ of \ two \ or \ more \ notches) & if \ \Delta R^{*}_{i,t} \leq \mu_1 \\ -1 \ (i.e. \, downgrade \ of \ one \ notch) & if \ \mu_1 < \Delta R^{*}_{i,t} \leq \mu_2 \\ 1 \ (i.e. \, upgrade \ of \ one \ notch) & if \ \mu_2 < \Delta R^{*}_{i,t} \leq \mu_3 \\ 2 \ (i.e. \, upgrade \ of \ two \ or \ more \ notches) & if \ \mu_3 < \Delta R^{*}_{i,t} \end{cases} \tag{9}$$

where μ_m represents thresholds to be estimated by maximum likelihood, along with the β parameters, subject to the constraint $\mu_1 < \mu_2 < \mu_3$. In addition, we calculate the marginal effects of past upgrades and downgrades to estimate the economic significance of each explanatory variable.

In order to incorporate in our analysis outlooks and watch list inclusions, we create a new variable named *comprehensive credit rating* following Gande and Parsley (2005), (CCR; see Appendix B). CCR is a modified version of the CR scale in Appendix A (panel A), based on one additional category created to be placed between the letter ratings used by CRAs to allow us to incorporate watch list inclusions, and changes in outlooks (see Appendix A, panel B). The new CCR variable increases (decreases) by one when a CRA action implies a negative (positive) watch list inclusion (exclusion) or one outlook change. CCR increases (decreases) by two when a full upgrade (downgrade) in rating takes place. To incorporate outlooks and watch lists in the Granger-causality vector autoregressive framework in (1)-(4), similarly to the case of changes in ratings only, we treat upgrades and downgrades separately.[19]

The predictive ordered-probit framework in equations (7)-(8) allows us to separate changes in ratings from changes in outlooks and watch list movements. For the dependent variable we denote upgrades (downgrades) in ratings of one notch by +2 (-2), and more than one notch by

[19] For upgrades (downgrades), we assume that positive (negative) changes in outlooks and ratings carry the same weight, since logistic regressions restrict the level of analysis to a binary framework.

+4 (-4). For positive changes in outlooks with (without) a simultaneous upgrade in rating, we use +3 (+1). Similarly for negative changes in outlooks with (without) a simultaneous downgrade in rating, we use -3 (-1). Therefore even numbers are comparable to the values used without watch list movements and changes in outlook data.

3.2.3 Structural Instability Tests

As a consequence of the regulatory framework changes that have occurred between 2002 and 2006, in H2 we express our prior belief that issuer-paid CRAs may have altered their behaviour and improved their timeliness in recent years. Accordingly, we also expect to find changes over time in the parameters of the statistical models described in sections 3.2.1-3.2.2 to test for causality. A first, rudimental step to assess the existence of breaks in the parameters that capture timeliness scores and lead-lag relationships, consists in splitting our sample in two distinct periods with the purpose of comparing sub-sample estimation results: the pre-regulatory overhaul (henceforth, PRE) period July 1997 - June 2002; the post-regulatory change (henceforth, POST) period July 2002 - December 2007. The PRE sample conveniently corresponds to BSS data. A formal way to check whether an econometric model is stable over time is to test whether the parameters of two different regressions, corresponding to disjoint sub-periods, are equal. This procedure is normally known as a partial F-test or Chow test (Chow, 1960) of structural change. However, because in this paper we are using ordered probit maximum likelihood methods instead of simple OLS, we use a log-likelihood ratio test as an alternative to the partial F-test.[20] An LRT of structural instability is based on a comparison of the maximized log-likelihood of two alternative models: one unrestricted and in which the instability is captured by interaction effects triggered by a "break dummy" (b_t); another model is restricted not to include any dummies. The two models are presented in equations (10) and (11), respectively:

$$\Delta \mathrm{R}_{i,t}^{*A} = \sum_{h=1}^{6} \beta_h^1 D\_up_{i,t-h}^{B} + \sum_{h=1}^{6} \beta_h^2 D\_dw_{i,t-h}^{B} + \sum_{h=1}^{6} \beta_h^3 \cdot b_t \cdot D\_up_{i,t-h}^{B} + \sum_{h=1}^{6} \beta_h^4 \cdot b_t \cdot D\_dw_{i,h}^{B} + \varepsilon_i;\ \varepsilon_i \sim N(0,1) \tag{10}$$

[20] In the case of linear regressions, the likelihood ratio and the F (Wald) tests are asymptotically equivalent.

$$\Delta \mathrm{R}_{i,t}^{*A} = \sum_{h=1}^{6} \beta_h^1 D\_up_{i,t-h}^{B} + \sum_{h=1}^{6} \beta_h^2 D\_dw_{i,t-h}^{B} + \varepsilon_i;\ \varepsilon_i \sim \mathrm{N}(0,1) \tag{11}$$

where $\mathrm{b_t}$ is a dummy variable that takes the value one prior (or in correspondence to) the POST period and zero otherwise. In the POST period, $b_{\mathrm{t}} = 1$ affects also the coefficients that load $\Delta \mathrm{R}_{i,t}^{*A}$ onto lagged values of previous rating changes that occurred during the PRIOR period, i.e., the regulatory reforms affects the way in which $\Delta \mathrm{R}_i^{*A}$ responds to previous rating changes. The null hypothesis to be tested can be formulated as $\mathrm{H_0}\colon \beta_1^3 = \cdots = \beta_6^4 = 0$.

The LRT-based test described above assumes that the break date is known, as it is typical of Chow tests. We argue that, in the light of our background work on the evolution of the institutional landscape concerning the regulation of CRAs, such as an assumption may be reasonable, similarly to what has been argued by Cheng and Neamtiu (2009). However, econometric methods have been developed to perform tests similar in spirit to the known-data Chow-style tests when the break date is unknown. As a result, we also iteratively compute the LRT break test in correspondence to each potential break date between January 1999 and December 2007, where both dates are strictly contained in our overall sample period to allow us the chance to meaningfully estimate the restricted model in equation (10).

3.2.4 Event Studies

If rating actions were completely uninformative, then no abnormal returns should be observed as a result of a rating change affecting the issuing company, because all of the information known at the time of the rating change should have already been reflected in the prices of the securities of the company. This principle should concern both corporate bonds to which financial ratings directly refer to, and stocks, as several studies have shown that stock returns are a significant predictor of bankruptcy (Shumway, 2001). Therefore abnormal stock returns may reflect the informational content of a credit rating change that has not yet been impounded into prices. Hence, stock returns can be used to make inferences about the timeliness of rating changes with respect to information that is relevant to equity pricing. Thus, we conduct a series of stock market event studies to investigate hypotheses 3-5.

To perform our event studies, we first form pairs of CRAs and select only those rating changes that concern firms covered by both agencies during a specific period of time. Each pair always includes one investor-paid credit rating agency (EJR) and one of the issuer-paid CRAs. Once we have selected a pair, we define two types of rating changes: conditional and unconditional (see section 2 for a detailed definition). In essence, a conditional rating change is a rating upgrade (downgrade) that occurs at time t and that has being preceded by a rating upgrade (downgrade) in the prior 30 days. Because our event studies are based on pairs that reflect the structure of hypotheses 4-6 above, we eliminate any observations that concern ratings that are "polluted" by actions performed by CRAs outside the specific pair under investigation (i.e., confounding events in the sense of Stickel, 1986) as they may lead to erroneous inferences. An unconditional rating change is instead an upgrade (downgrade) by a CRA that has not been preceded by any rating changes by other CRAs within the previous 30 days. Robustness checks reveal that the choice of a 30-day window is immaterial for the results in section 5.

We follow standard methodology (see e.g., Campbell et al., 1997) in performing our event studies. The market model for security i at time t is given by

$$\mathrm{R}_{i,t} = \alpha_i + \beta_i R_{m,t} + \varepsilon_{i,t}. \tag{12}$$

This regression can be equivalently re-written as a compact regression, $\mathrm{R}_i = X_i \phi_i + \varepsilon_i$.[21] Given an estimate of ϕ_i, the abnormal returns are calculated as $\hat{\varepsilon}_i = \widetilde{\mathrm{R}}_i - \widetilde{X}_i \hat{\phi}_i$, where $\widetilde{\mathrm{X}}_\mathrm{i} = \left[\iota\ \widetilde{\mathrm{R}}_\mathrm{m}\right]'$ is a $[T_1, T_2]$, $(\mathrm{T}_2 - \mathrm{T}_1 + 1)$ x 2 matrix that refers to the $(\mathrm{T}_2 - \mathrm{T}_1 + 1)$-long event window that spans the period between T_1 and T_2.

Our first set of event studies examines the differences in abnormal returns for unconditional rating changes comparing investor- vs. issuer-paid CRAs. In fact, our experiments set $\mathrm{T}_1 = \mathrm{T}_2$ so that the event-imputed abnormal returns are measured only in correspondence to the day of the rating action. The estimation window starts 255 trading days before the event day and ends 46 days before the event day, i.e., L = 46 and $\mathrm{T}_1 = \mathrm{T}_0 + 255$. Because some literature (see, e.g., Armitage, 1995) has shown that the power of event studies is slightly stronger when the market factor returns are measured from an equally-weighted index, the vector R_m consists of

[21] $\mathrm{R}_\mathrm{i} = \left[\mathrm{R}_{\mathrm{iT_0+1}} \dots \mathrm{R}_{\mathrm{iT_1}}\right]'$ is the vector of returns sampled during the estimation window $[T_0, T_1\text{-}1\text{-}L]$, $\mathrm{X}_\mathrm{i} = [\iota\ \mathrm{R}_\mathrm{m}]'$ is a matrix $(\mathrm{T}_1 - \mathrm{T}_0 - L)$ x 2 with a vector of ones in the first column and the vector of market returns in the second column, and $\phi_\mathrm{i} = [\alpha_\mathrm{i}\ \mathrm{Q}_\mathrm{i}]'$ is the (2 x 1) vector of parameters. Under rather general conditions, ordinary least squares (OLS) is a consistent estimation procedure for the market model parameters.

the Center for Research in Security Prices (CRSP) equally-weighted portfolio.[22] Higher abnormal returns (in absolute value) with a sign identical to the change in rating for investor- vs. issuer-paid CRAs, imply that the informational content of the former CRAs exceeds those of the latter, so that this is evidence that the investor-paid business model is more informative than the issuer-paid model is. Our second set of event studies tests instead whether the market reactions due to conditional and unconditional rating changes by the same CRA have the same abnormal impact.

4. Sample Selection

Following prior studies such as Johnson (2004) and Beaver et al. (2006), we use EJR as a representative of investor-paid CRAs. The main reason for selecting this agency is that it covers a range of rated firms comparable to that of the issuer-paid NRSROs, unlike other investor-paid agencies that tend to specialize in certain industries.[23] As representatives of the issuer-paid CRAs, we focus on the "big three NRSROs", i.e., Moody's, S&P, and Fitch. [24]

In our analysis, we assume that a firm's credit risk is best reflected by senior unsecured ratings (similarly to BSS, 2006). Therefore our sample covers senior unsecured credit rating *changes* made by Moody's, S&P, Fitch and EJR over the period from July 17, 1997 through December 21, 2007.[25] The start date of our analysis corresponds to the first rating action by EJR and so it represents an objective initial date. The end date corresponds to the date on which EJR was granted the status of NRSRO (with SEC release No. 57031). We end our sample on this date because the behaviour of EJR ratings may have experienced significant alterations after NRSRO status was granted. In essence—although this also represents an interesting

[22] As a robustness check, we have performed all event studies afresh using the value-weighted CRSP index to compute abnormal returns and used a range values for $\kappa \equiv T_1 - T_2$, κ=3 and 5, to define symmetric windows. Results are unchanged and available from the authors upon request.

[23] Additional reasons to focus our analysis of investor-paid ratings on EJR are given in BSS (2006). For instance, during the 2002-2006 process of regulatory overhaul, EJR has participated in both the CRA hearings in Congress and at the SEC. Since its foundation in 1995, EJR has been rating more than 1,300 companies in the industrial, financial, and the service sectors. EJR market their ratings via a subscription service on Bloomberg.

[24] The credit rating business remained dominated by these three CRAs, with a market share in excess of 80% as of the end of 2009. Although its market share is not completely negligible, ratings by A.M. Best are not used because not available before 2005 and because A.M. Best mostly rates firms in the insurance industry.

[25] Data from July 1997 to June 2002 were generously provided by BSS (2006). We would like to thank Egan Jones Ratings for granting us access to their data since 1999. The entire dataset was then obtained by merging the data provided by BSS with those collected from EJR. Any duplicate observations were deleted.

research question worth further pursuing—one cannot rule out that after 2008 EJR may have obeyed new but mixed incentives because, although its business model remained strictly investor-paid, it came to also enjoy the rents from the market power exercised by NRSROs (see Kisgen and Strahan, 2010; White, 2010, for related comments). Hence, it seems safe to limit our analysis to 2007 data. Note that our sample is however considerably longer than the July 1996 - June 2002 data set used by BSS (2006), because we explicitly pursue the investigation of potential breaks in timeliness and informativeness of (relative) EJR ratings between 2002 and 2006. In this respect, if the reform of the rating industry undertaken by the U.S. regulators and legislators after 2002 had reached completion by 2006, to use data for all of 2006 and 2007 should give our structural break tests sufficient power to isolate any (parametric) instability.

The sample of EJR bond rating actions comprises 24,800 observations. EJR bases its activity on five categories of actions: initial ratings (first coverage), upgrades, downgrades, affirms (when EJR reviews but maintains the currently assigned rating), and drops (when they stop their coverage of a firm). In this study we are interested only in rating changes (upgrades and downgrades) and therefore we delete initial ratings, affirms and drops, leaving 5,016 credit rating changes by EJR over the period.

We obtain Moody's, S&P and Fitch rating actions from the Mergent Fixed Income Securities Database (FISD). We choose only those actions corresponding to senior unsecured bonds, which leaves us with 28,875 rating actions for our sample period. Table 1 reconciles our data sources. Note that because EJR started publishing outlooks only in July 1999, when outlooks and watch list movements are taken into account, two years of data on rating changes are lost as the sample starts then in July 1999.

Insert Table1 here

As all of our CRAs also provide watch list data, we create two different samples, one without watch list data and the other also incorporating them. In the first case we select only rating changes, narrowing down the sample to 5,518 observations. Then we removed, for all firms, all rating actions that occurred before EJR's first investor-paid action to make temporal coverage homogeneous, narrowing the sample to 4,890 observations. For the second case we also include those actions that imply a change in the variable CCR built in section 3, for a total of 6,333 observations. We then assign to each rating a numerical value, and reconcile Moody's, Fitch and S&P's schemes, according to the conversion in Appendix A. We carry out our analyses by comparing EJR with each of the issuer-paid CRAs, one at the time. To select the corresponding rating changes for each pair, we look for those firms covered by both agencies. To perform the event studies, we also require that each firm has a permanent security identification number (PERMNO) assigned, so to retrieve daily prices from CRSP.

5. Results

5.1 Descriptive statistics

Tables 2 and 3 present descriptive statistics and detailed distributions of rating changes (excluding watch list data) for each of the three pairs formed by matching EJR with the three major NRSROs during our sample period. Table 2 presents the summary statistics of rating levels to which the rating agencies change their ratings to. It is clearly visible that EJR performs more rating changes than any of the issuer-paid CRAs do. For example, EJR changes ratings almost three times more often than Fitch does (2249 vs. 751), and more than twice than Moody's and S&P do (3279 vs. 1462 and 3397 vs. 1504, respectively). The mean rating from EJR in each of the two tables is close to 10 and the median is exactly 10, which corresponds to the lowest investment-grade rating (see Appendix A, first panel). Interestingly, each of the issuer-paid CRAs have weakly higher mean and median ratings, with means of 11 or higher in two of the three cases (the distance is smaller when Fitch is compared to EJR). Even though the difference is modest, this implies that issuer-paid CRAs assign a lower average (median) rating to the same firms covered by EJR. Table 3 shows details on the distribution of ratings by each CRA, generally confirming the impression that in general EJR is not more severe than the three big issuer-paid ones are. Because the differences in the average scores or distributions fail to appear major or to support the idea that the investor-paid business model yields in any way more restrictive ratings, then any differences must arise from the relative speed with which these are adjusted over time, or from the nature of such adjustments, which is what sections 5.2-5.4 explicitly test.[26]

[26] We note some differences between our data and the data used by BSS (2006): (a) BSS data start in July 1996 while ours begin in July 1997; (b) we match issuer-paid data from Fitch, Moody's and S&P from Mergent FISD, while BSS use *only* Moody's data from Moody's Corporate Bond Default Database. Hence results over BSS sample period can be expected to be similar but never completely identical.

Insert Table 2 and 3 here.

5.2 Timeliness tests: logistic regressions

We begin our analysis by using a Granger causality test to determine which agency in each pair leads and which follows, if any. We perform the analysis for three different sub-samples: (a) July 1997- June 2002 (the PRE sample); (b) July 2002- December 2007 (the POST sample); (c) the full sample, July 1997- December 2007. We test for Granger causality by calculating a likelihood ratio test (LRT) that compares an unrestricted model (that incorporates previous upgrades/downgrades of the other credit agency), against the explanatory power of a restricted model (where rating changes only depend on previous upgrades/downgrades of the same CRA). Table 4 shows the LRT results "between" (i.e., both from and to) EJR and each of the three issuer-paid CRAs. The upper half of each panel includes only ratings change data; the lower part of each panel additionally incorporates watch list movements and outlooks.

Starting with the period July 1997 - June 2002 and the pair EJR/ Fitch, Table 4 yields a test statistic of 1.58 (p-value of 0.45), which shows that EJR's upgrades are not Granger-caused by Fitch's previous upgrades. In other words, Fitch upgrades in a range of 1 to 180 days prior to an upgrade by EJR do not help to predict EJR's upgrades. In the second column we report test statistics for the opposite and more interesting causality channel, i.e. the case in which investor-paid CRA upgrades may forecast the issuer-paid ones. We reject the hypothesis that Fitch is not Granger caused by EJR, as over the same sample the LRT is 18.68 (p-value of 0.00). The same lead-lag relationship applies to downgrades, as we fail to find evidence of Fitch Granger-causing EJR (p-value is 0.25) but we strongly reject the null of Granger-causality from EJR to Fitch (p-value < 0.00). These results confirm the results in BSS: at least in the PRE sample, the investor-paid business model could produce much timelier ratings than the issuer-paid model did.

Insert Table 4

When we repeat this analysis again with reference to the EJR/Fitch pair, but with reference to the second sub-period (June 2002 - December 2007, columns 3 and 4 in table 4), we find that the lead-lag relationship between EJR and Fitch turns now bi-directional if one tests using a 5% test size, in the sense that for both upgrades and downgrades, while one can now reject the null of EJR not Granger-causing Fitch with p-values below 0.00 (the LRT statistics are 48.4 and 46.4 for upgrades and downgrades, respectively), we can also reject the hypothesis of Fitch not Granger-causing EJR with p-values between 1 and 5 percent (the exact p-values are 0.04 and 0.02 for upgrades and downgrades, respectively).

However, when the same analysis is applied to the entire sample, the results start being different across upgrades vs. downgrades. In the former case, we find additional evidence of bi-directional linkages, which signals that the strong leading position of EJR over Fitch has sufficiently weakened after 2002 to cast doubts on the overall findings in BSS (however related to Moody's, but see below for directly comparable evidence). In the latter case, for downgrades, while it is the case that EJR leads Fitch (p-value is < 0.001), there is marginal evidence that also Fitch actions forecast EJR's (p-value is 0.11).

The results in table 4 are similar for the pair EJR/S&P (panel III): for upgrades, in the PRE period EJR actions Granger-cause S&P's (with p-values well below 0.001), but S&P actions hardly predict EJR's. In the POST period, the bi-directional relationship is instead pervasive as rating actions by both CRAs predict subsequent actions by the other CRA. This finding extends to the full-sample period. However, it is obvious that in the case of the EJR/S&P, the bi-directional nature of the linkages existed already in the PRE sample in the case of downgrades. Finally, results for the pair EJR/Moody's in table 4 (panel II) are qualitatively similar to those already described, but statistical significance in lead-lag relationships now prevails in all tests. The only difference between the PRE and POST samples is that while in the POST sample bi-directional linkages emerge once more with p-values inferior to 0.001, in the PRE sample the power of Moody's in predicting subsequent rating actions by EJR is occasionally weaker (and yet significant at conventional levels).[27]

All in all, this is a first important empirical finding: while BSS had found on their 1997-2002 sample that investor-paid CRA rating changes led issuer-paid rating actions but the opposite did not apply, when the analysis is extended to the 2002-2007 period, no uni-directional causality obtains in the sense that the issuer-paid rating changes now also predict the investor-paid decisions. This is consistent with the idea that the timeliness differential between EJR and the traditional issuer-paid NRSROs may have weakened as a result of the new institutional

[27] The fact that by construction our sample starts a year later than the sample in BSS (2006) and also the different data sources for issuer-paid agencies may explain the deviations from the results in BSS.

background created by the legislative reforms of 2002. Equivalently, while BSS-type results obtain in the PRE sample, they instead fail to emerge during the POST sample, which in its turn is consistent with our hypothesis 2.

The lower part of panels I-III in table 4 concerns the case in which outlook and watch list actions are included in the analysis which, at least to our knowledge, is new in the context of an analysis of the links between investor and issuer-paid CRAs. In this case, we implicitly assume that outlooks carry the same weight as formal rating actions, which is of course a simplistic hypothesis. Interestingly, results change significantly and important differences emerge between upgrades and downgrades. For instance, in panel I, in the PRE period while when it comes to downgrades we uncover that while EJR leads Fitch (p-value < 0.001), the opposite is not true (p-value exceeds 0.05), in the case of upgrades we have the opposite result. In the POST regulation overhaul period, results change: for downgrades we find once more evidence of bi-directional links as in the case in which only ratings were incorporated in the analysis; for upgrades there is now evidence of EJR Granger-causing Fitch. Full-sample results confirm those from the POST period: unidirectional Granger causality appears only for upgrades, when EJR causes Fitch and not the other way around.

Panels II and III of table 4 have instead a homogeneous structure after outlooks and watch lists are included: when it comes to downgrades, there is no evidence of Granger-causality, in the sense that causality is always bi-directional, with EJR actions predicting subsequent S&P and Moody's actions, and vice versa; all these lead-lag relationships are estimated to occur with p-values that are systematically inferior to 0.001. These patterns emerge both in the PRE and in the POST sample periods as well as in the full sample, so—consistently with our hypothesis 1—there is no indication of a differential in relative timeliness of downgrades when negative outlooks and watch list movements are included in the analysis. This may be due to the fact that while the ratings of the three "big" issuer-paid NRSROs play a key quasi-legal role in a number of statutes and regulations and are used as indexation parameters in a range of contracts, this does not apply to outlook and watch lists, so that issuer-paid CRAs may have actually always behaved as aggressively as investor-paid CRAs do.

5.3 Timeliness results: ordered probit analysis

We resort to ordered probit regressions to assess the lead-lag relationships between CRAs as well as *the impact of the size of a rating change* by the potential CRA leader on the probability of a rating change by the potential follower. As in section 5.2, the same estimation steps are undertaken without (table 5) and with (table 6) information on outlooks and watch list movements. Panels I through III in table 5 present results for the lead-lag relationships between EJR and each of the issuer-paid CRAs for our entire sample.[28] Panel I of table 5 shows the results for the pair EJR/Fitch. As for a potential causal link from Fitch to EJR, the only statistically significant variable in the case of upgrades corresponds to changes between 121 to 150 days prior to an EJR rating change. The estimated coefficient is 0.84 with a p-value of 0.02. The coefficient is positive meaning that a previous, even though distant upgrade by Fitch will increase the probability of an upgrade by EJR. However, the evidence of Fitch leading EJR is much stronger when it comes to downgrades: in this case, 4 out of 5 estimated coefficients, with reference to Fitch's actions undertaken between 1 and 120 days before, are statistically significant (with p-values < 0.001) and positively affect the probability of an EJR downgrade; the marginal impacts to the right of table 5 show that these effects are economically large.

The bottom section of panel I shows that EJR past rating changes strongly affect the probability of subsequent Fitch actions: for both upgrades and downgrades, 5 out of 6 estimated coefficients are statistically significant (four with a p-value of 1% or less) and they all have the expected sign. Some of the marginal effects imply considerable economic impact: for example, the value of 32.0 corresponding to the variable Δ-down (h=1) by EJR within the last 30 days means that if this variable turns from zero to one (signalling that EJR has downgraded a firm over the past month), then there is a 32% higher probability that Fitch will also follow up with an aggressive two-notch downgrade. Interestingly, because the adjacent marginal effect is -8.4, a recent EJR downgrade implies a decrease by 8.4% in the probability of a Fitch one-notch downgrade within a single month. This may imply that Fitch follows over the short run EJR's aggressive moves in a very aggressive way, almost applying a multiplier to

[28] Separate tables for the PRE (July 1997 - May 2002) and POST (June 2002 - December 2007) sub-samples are shown in Appendix C.

EJR's decisions.[29] Moreover, still looking at marginal effects, the biggest probability of Fitch following a previous EJR's upgrade corresponds to a rather delayed response to actions undertaken between 91 to 120 days before. Yet, for downgrades, marginal effects imply that the largest probability that Fitch downgrades occurs as a reaction to EJR downgrades over the previous 30 days, which signals a pronounced reactivity by Fitch.

Insert Table 5 here

Panel II of table 5 yields results that are qualitatively similar to panel I, but that concern instead the pair Moody's/EJR: while past upgrades by Moody's hardly affect the probability of rating changes by EJR (with the only exception of the [0, 30] days interval lag), past downgrades by Moody's do affect current EJR downgrades. As one would expect in the light of earlier evidence and panel I, EJR leads both in the upgrades and the downgrades. In the case of recently decided downgrades by EJR we find that the probability of a two-notch Moody's downgrade within 30 days increases by 26% (the respective result for a Moody's downgrade within 30 days in Section I is 22%). In the case of upgrades the probabilities are larger when EJR upgraded in the previous 1 to 150 days and probabilities are also larger for a one notch upgrade by Moody's. For downgrades, it is more likely we witness a Moody's downgrade when EJR has just downgraded the same stock in the previous 30 days. In panel III of table 5 the evidence turns instead bi-directional both for upgrades and downgrades, in the sense that although it remains obvious—both in economic terms and in terms of statistical significance of the associated coefficients—that past EJR rating actions affect subsequent decisions by S&P, there are weaker and yet accurately estimated indications that also past upgrades and downgrades by S&P affect EJR's actions.

Hence while EJR appears to be an obvious leader in upgrades (at least as far as Fitch and Moody's are concerned), the order probit analysis gives evidence of a bi-directional relationship for downgrades that involves all pairs of issuer- vs. investor-paid CRAs. This is clearly inconsistent with the earlier findings by BSS (2006) for the PRE 2002 period: over the full sample, investor-paid downgrade decisions fail to lead those enacted by issuer-paid ones. However, it remains the case that even for downgrades the magnitude of the economic

[29] This pattern of Fitch aggressively reacting (multiple notching) to EJR's downgrades applies to all lags investigated. However, it does not characterize upgrades, in the sense that Fitch responds in the same direction as any type of previous upgrade by EJR.

(marginal) effects that connect issuer-paid to investor-paid rating changes are somewhat stronger than those involving the reverse link. For instance if we focus on the average, marginal effect of all past EJR upgrades associated to statistically significant coefficients (at 5%) onto the probability of subsequent upgrades by issuer-paid CRAs, we have a sizeable +10.5% increase in probability; the analogous number for downgrades is 9.0%. The similarly computed average marginal probability effects of all past issuer-paid upgrades (downgrades) on subsequent EJR upgrades (downgrades) are 7.7% and 8.2% only.[30]

Yet, these conclusions partially inconsistent with the classical assumption that investor-paid CRA actions should lead issuer-paid actions, have been reached with reference to the full-sample period. Hypothesis 2 leads us to check whether this novel result may depend on the stronger incentives for issuer-paid CRAs to produce timely ratings as a consequence of the regulatory reforms of 2002-2006. A table with structure similar to table 5 in Appendix C, that reports parameter estimates from an ordered probit fitted on a 2002-2007 sample reveals that this the case: the bi-directional dynamic linkages between issuer- and investor-paid CRAs got stronger during and after the intense scrutiny that the industry has undergone. For instance, in the case of the much researched pair EJR/Moody's, while before 2002, Moody's upgrades hardly affected the probability of subsequent EJR upgrades, after 2002 the effect becomes powerful, as it concerns not only the most recent among Moody's decisions, but also those in the [31, 60] prior days interval (see table C1, panel II). At the same time, for instance for the EJR/S&P pair, we observe that some of the predictive power from past EJR to subsequent S&P actions is lost, as the corresponding estimated parameters stop being significant (see table C1, panel III). This is an indication that the lead-lag dynamic relationship between issuer- and investor-paid CRAs has become increasingly symmetrical after 2002. Of course, these comparisons between table 5 and the tables in Appendix C do not represent a formal test of the occurrence of breaks in parameters, for which we defer to section 5.4.

Next, the ordered probit allows us to deal with hypothesis 1, i.e., that any lead-lag dynamic relationship that puts investor-paid CRAs in a leading position weakens when the data are extended to also include outlooks and watch list movements. Therefore in table 6 we propose

[30] Moreover, the predictive power of an EJR downgrade for issuer-paid downgrades is always estimated to last for at least 6 months, while in the case of issuer-paid agencies varies, and usually the implied predictive power only lasts between 4 and 6 months, depending on the issuer-paid CRA under examination.

the same estimation exercise as in table 5, but with outlook and watch list decisions accounted for. Our hypothesis 1 finds all the support that it can find, given earlier evidence that in our full sample, hypothesis 1 is inconsistent with the evidence. Any leadership of investor-paid CRAs over issuer-paid ones further weakens when outlooks and watch list movements are taken into account. However, the bi-directional characterization in which all CRAs jointly influence one another already reported in table 5 is strengthened.

Panels I-III of table 6 may be laid on top of the corresponding panels of table 5 finding few differences (excluding that marginal effects can now be computed for a wider range of rating actions). In panel I, we find again that Fitch and EJR move as a pair, with rather complex leads and lags, when it comes to downgrades, while EJR mostly leads Fitch in upgrades. In panel II, the result is identical but now, compared to both panel I and with panel II of table 5, the bi-directional relationship between EJR and Moody's becomes stronger and mostly arising from the [0, 60]-day interval. In panel III, the evidence is similar to panel III of table 5, but the evidence of EJR and S&P influencing each other in complex ways has gotten more robust.[31]

Insert Table 6

Table 6 reveals details of the reaction of CRAs to the actions of other CRAs that were not visible before. For instance, in panel I we note (see the row of total marginal effects for each pair inside the panels) that while the most likely reaction by EJR to Fitch is to assign a consistently signed outlook to the same firm, the most likely reaction by Fitch to an EJR action consists of an equally-signed full rating change. The evidence is qualitatively identical in panel III, concerning the pair EJR/S&P. This confirms earlier impressions that if there is anything to the lead-lag result reported by BSS, this may apply to the pairs EJR/Fitch and EJR/S&P. In fact, panel II of table 6 shows that EJR and the remaining issuer-paid CRA, Moody's, come close to stand on an equal footing with EJR: in the same way in which the most likely reaction by EJR to a Moody's action is to assign a consistently signed outlook to the same firm, the likely reaction by Moody's to EJR consists of a coherent outlook assignment.

These remarks allude to the existence of further "multiplier effects" in table 6: even when EJR also follows other CRAs—as it is always the case with downgrades, and to a large extent

[31] We also estimated the unrestricted versions of equations (7) and (8), which include lagged values of changes in ratings by the CRA used as the dependent variable. Results are robust to this specification.

also upgrades, at least when Moody's and S&P are involved—its reaction appears to be "measured". Consider for instance the pair EJR/Moody's in panel II of table 6. Moody's is an appropriate choice because it seems the issuer-paid CRA for which the relationship with EJR is more symmetric. In the table we see that a recent downgrade by Moody's is unlikely to cause a more-than-proportional two-notch downgrade by EJR, as the probabilities of a "-4" adjustment are on average 7% only. Yet, a recent downgrade by EJR is usually likely to cause a more-than-proportional two-notch downgrade by Fitch or S&P. This means that tables 5 and 6 emphasize that EJR retains some form of (admittedly, weak) lead over the issuer-paid CRAs, but only through the multiple notching strength of issuer-paid reactions.[32]

We also perform LRT (Granger-causality) tests based on ordered probit models. Specifically, we use an LRT test that compares a restricted model (equations (7) and (8)) with an unrestricted model that incorporates as explanatory variables lagged upgrades/downgrades from the same CRA that gives the dependent variable events. To save space, the results are not tabulated but are available upon request. For instance, when outlooks and watch list inclusions are accounted for, we detect a bi-directional relationship at the 1% level across all pairs of CRAs, except the EJR-Fitch pair over the period 1997-2002 , which gives indication of weaker bi-directional linkages at the 5% p-value only.

5.3 Endogenous Instability Tests

We now turn to formal instability tests. The results of the causality tests described in section 5.1 suggest that there is a structural change in the parameters of our equations, as we estimate different lead-lag relationships before and after 1 June 2002.[33] We apply the log-likelihood ratio test introduced in section 3.2.2. The results are shown in Table 7 in three panels, one for

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[32] Appendix D reports estimates of ordered probit models identical to those in table 6 based on POST regulatory overhaul (2002-2007) data only, apart from some episodic loss of statistical significance due the smaller sample size (e.g., in the case of Moody's, the number of observations on upgrades declines from a total of 6,233 in table 6 to 4,258; the number of observations on downgrades declines from a total of 2,107 in table 6 to 1,371). The only difference is that over the second sub-sample all CRAs seem to have become increasingly rapid in incorporating other CRA's actions in their own decisions, in the sense that a few indicators measuring prior rating actions between 90 and 150 days turn insignificant. Although this has not been formalized in section 2, this finding is consistent with the joint implications of hypotheses 1 and 2: if lead-lag relationships favouring investor-paid CRAs are structurally weak because issuer-paid CRAs were already exploiting outlooks and watch list to achieve higher timeliness, then we do not expect dramatic effects from the regulatory changes of 2002-2006.

[33] The results in section 5.2 are also partially supportive of the presence of a break, even though the results in Appendices C and D are more mixed in their economic implications.

each of the three pairs, EJR/Fitch, EJR/Moody's, and EJR/S&P. Within each panel, each column shows the values of the LRT for the case of one CRA acting as a rating follower.[34] In the flank of the table we have listed a range of quarterly dates—between Jan. 2001 and October 2007—in correspondence to which we have computed the Chow-type LRT. In the table, we emphasize the dates in correspondence to which the null of structural change is rejected with a p-value of 1% or lower.

Insert Table 7

Table 7 shows three stark results. First and foremost, there is clear evidence of breaks in the lead-lag relationships captured by our logistic regression model. In the case of both the EJR/Fitch and EJR/Moody's pairs, such a break is dated as occurring in July 2003. The fact that for two different pairs, the dating results turn out to be endogenously identical is remarkable. Moreover, such a date, only a few months after the new legislation passed in 2002 became effective, is sensible and consistent with our institutional background overview in section 2. In particular, July 2003 shortly follows the Congress hearings on the reform of the rating industry held in April 2003. However, in the case of the EJR/S&P pair, two different and rather "late" dates trigger the LRT test to reject the null of parameter stability: January and April 2007. On the one hand, these two dates are adjacent, which may sensible if our test statistic incurs into problems in allocating the break to a precise time. On the other hand, the fact that the breakpoint date turns out to follow by almost 4 years the date identified in the case of the remaining two issuer-paid NRSROs may be reason for concern. Even though it may be conceivable that a complex and progressive set of regulatory reforms may have taken some time to filter through our data, it can be argued that the fact that the dates fail to match the exact timing of the regulatory milestones identified in section 2 may be normal: CRAs are adverse to sudden changes to their methodologies; the re-evaluation of the firms they rate is a process that is likely to take several quarters, in principle as long as a full rating cycle concerning each of the rated firms. In any event, our tests signal the occurrence of a break in July 2004 when Moody's leads EJR. Because section 5.2 has shown that the relationship between Moody's and EJR tends to be "more symmetrical" than other pairs are, this may be

[34] To save space, we only present results based on rating changes, although they are robust to the inclusion of outlooks and watch list movements. Complete estimation results are available upon request.

taken as evidence that—in the measure in which the EJR/Moody's connection is actually largely bi-directional—this has been affected as well by the new institutional environment.

5.4 Event Studies

Our final step is to analyse abnormal stock returns that are due to rating changes, in order to assess the reactions by markets to announcements by both types of CRAs and whether these reactions are affected by the unconditional/conditional nature of rating actions. In what follows, we present results for both the case in which outlooks and watch lists are excluded (tables 8-10) and when they are included (table 11).

Table 8 shows the results of event studies for rating changes obtained on the overall, 1997-2007 sample. To be able to test hypothesis 3 from section 2, we focus again on firms that are simultaneously rated by pairs of CRAs, EJR vs. one of three issuer-paid. Interestingly, the market shows uniformly significant abnormal returns in the case of EJR rating actions, both upgrades and downgrades (these are the abnormal return statistics that in the table appear in the rows labelled as "unconditional" and in correspondence to pairs in which EJR is listed first). The reactions to issuer-paid actions are smaller but statistically significant in the case of Moody's and S&P's downgrades but never significant for upgrades and for Fitch's (these are the abnormal return statistics that in the table appear in the rows labelled as "unconditional" and in correspondence to pairs in which the issuer-paid CRA is listed first). This finding is consistent with the ample evidence in the literature of an asymmetry in reactions to downgrades vs. upgrades (see Ederington and Goh, 1998; Hand et al., 1992; Holthausen and Leftwich, 1986). In the table, we mark with the symbols ΔΔΔ, ΔΔ, or Δ pairs of investor- vs. issuer-paid CRAs for which the null of a test of differences in mean is rejected with p-values of less than 1%, 5%, or 10%, respectively.[35] Consistently with hypothesis 3, in table 8 for unconditional rating actions, the stock market abnormal reaction on the day of a rating announcement by EJR is always significantly in excess of the reaction to any of the three issuer-

[35] In tables 8-11, it is possible for the same pair of CRAs to generate different "unconditional abnormal returns" (e.g., in table 8 we have -3.82% for the pair EJR/Fitch but -2.23% for the pair Fitch/EJR, in the case of downgrades) because these concern a different number of companies (e.g., 1,238 and 392 for the pair in the EJR/Fitch example): the former number is defined as the companies covered by an issuer-paid CRA on which EJR expresses a later rating action in the same direction; the latter as the companies covered by EJR on which an issuer-paid CRA expresses a later rating action in the same direction.

paid NRSROs, with p-values always below 1%. The evidence is stronger in the case of upgrades, when the p-values of mean difference tests are below 0.001 and the differences are economically large, 2.42% (a day) in favour of EJR over Moody's, 2.60% for EJR over S&P, and 2.15% for EJR over Fitch. Although the differences are smaller in the case of downgrades (2.26% in favour of EJR over Moody's, 1.70% for EJR over S&P, and 1.59% for EJR over Fitch), they remain highly statistically significant.[36] Therefore, there is economic value in trading on the basis of CRA actions and the resulting returns are significantly larger when these are based on the rating actions of investor-paid agencies.[37]

Insert Table 8 here

Table 8 partially validates our hypothesis 4: abnormal stock market reactions to conditional rating changes are significantly stronger than reactions to unconditional ones. However, as stated in our initial conjecture, this holds only in the case of downgrades, for which the confirmatory effect of a later action on the basis of some earlier decision may be stronger, and when the second downgrade comes from EJR, in the wake of an earlier action undertaken by an issuer-paid CRA. In particular, an EJR downgrade conditioned on an earlier Moody's (Fitch's) downgrade yields an abnormal return of -7.41% (-7.50%) against a reaction to unconditional actions of -3.74% (-3.82%) and the difference is significant with a p-value of 0.004 (0.078). However, such a result does not obtain when issuer-paid actions have been preceded by investor-paid actions (in fact, in the case of upgrades, the difference has the wrong sign in two cases out of three). In other words, market participants think that when an issuer-paid agency makes a conditional follower-type change, it is reacting to EJR, but on the contrary when EJR makes a conditional downgrade, the market assumes they are unveiling new information. The results are in fact opposite when it comes to upgrades: even though most tests of paired mean difference yield p-values of 10% or lower (with one exception), the effects of unconditional

[36] Although hypothesis 3 concerns unconditional rating actions, table 8 also shows that stock markets display a significantly greater abnormal reaction to the *conditional* downgrades of investor-paid CRAs than they do for issuer-paid agencies. For instance, EJR downgrades that condition on earlier Moody's actions cause an average abnormal return of -7.41%; Moody's downgrades that condition on earlier EJR actions cause an average abnormal return of -2.25%; the 5.16% difference is highly statistically significant, as indicated by the three ΔΔΔ symbols. Results are similar for S&P and Fitch, although they are weakly statistically significant in the latter case. However, the bottom panel of table 8 reveals no significant differences in reactions to conditional upgrades.

[37] Our findings related to issuer-paid CRAs are consistent with previous research, such as BSS (2006), Dichev and Piotroski (2001), and Hand et al. (1992) who explain these results with the asymmetric loss functions and with concerns for the reputational effects of downgrades that would be typical of issuer-paid CRAs.

upgrades are stronger than the effects of conditional ones. This means that while EJR's downgrade decisions carry so much value—either of a confirmatory type as discussed in section 2, or at least in excess of the information previously revealed by the issuer-paid CRA covering the same stocks—to lead to a conditional impact that exceeds the unconditional impact of an EJR rating action alone, in general this is not the case for upgrades, at least those in which EJR is the follower.[38] In this sense, hypothesis 4 finds only partial validation on our data.

Tables 9 and 10 show the results of tests performed along the same lines as in table 8, but in this case applied to two distinct sub-samples: table 9 concerns the PRE, 1997-2002 sub-sample; table 10 concerns instead the POST, 2002-2007 sample. Consistently with hypothesis 5, the goal is to show that abnormal stock market reaction differentials between issuer- and investor-paid actions have declined after 2002. Table 9 contains results that are qualitatively similar to table 8, with the only marginal difference that some of the statistical significance is weakened by the smaller sample size backing these new findings. In particular, abnormal stock reactions to conditional downgrades are significantly greater than reactions to unconditional changes. In the case of downgrades, an EJR action that follows either Moody's or Fitch's tends to produce further abnormal returns (3.75% and 6.10%, respectively) with p-values of 0.029 and 0.081.[39] On the opposite, the differences in abnormal reactions (with particular emphasis on those in which EJR is a follower) fail to yield much evidence in the POST sample, when we find it is the issuer-paid agencies that often "move" the market in risk-adjusted ways conditionally on earlier EJR actions and in excess of unconditional upgrades (this happens in the case of Moody's). Therefore, the 2002-2006 wave of regulatory reforms have exercised an effect on the ability of EJR downgrades to create economic value, especially the value that we have found to be accessible in a conditional form, i.e., in spite of earlier downgrades assigned by the three traditional issuer-paid NRSROs in advance of EJR. Although this is limited to downgrades, this confirms hypothesis 5.

Insert Tables 9 and 10 here

[38] However, when S&P and Fitch upgrade a stock after EJR, they obtain abnormal returns that exceed their (in fact, modest) abnormal returns that follow actions with p-values between 1 and 10%.

[39] The reduced sample size prevents us from performing all the tests and from obtaining adequate statistical significance in the case of upgrades, that, as commented in table 2, tend to occur less frequently.

Table 11 is the analog of table 8, when outlooks and watch list inclusions are taken into account. The table is virtually indistinguishable from table 8 and, if anything, results get a bit stronger. This is sensible because event studies consist of techniques to measure the economic value/stock price reaction to rating actions, and it is well-known (e.g. Hand, Holthausen and Leftwitch, 1992) that outlooks and watch lists convey timely and valuable information. Also in this case, the market shows uniformly significant abnormal returns in the case of EJR rating actions, both upgrades and downgrades (these are the abnormal return statistics that in the table appear in the rows labelled as "unconditional" and in correspondence to pairs in which EJR is listed first). In a majority of cases, the unconditional impact of EJR actions is larger when outlooks are included but the differential effects are economically small. The reactions to issuer-paid actions are smaller but statistically significant in the case of issuer-paid downgrades but never significant for upgrades. Consistently with hypothesis 4, for unconditional actions, the stock market abnormal reaction on the day of a rating announcement by EJR is significantly in excess of the reaction to any of the three issuer-paid NRSROs, with p-values always below 1%. The evidence is equally strong in the case of upgrades, when the p-values of mean difference tests are always inferior to 0.001 and the differences are economically large. All in all, table 11 confirms hypotheses 4 and 5 also when our data are expanded to include outlooks and watch lists.

Insert Table 11 here

In unreported tests, we have also constructed tables that mimic the structure of tables 9 and 10, but in this case including outlooks and watch list data. Our conclusions are not significantly affected: while in the PRE sample, results are qualitatively similar to table 11—for instance, abnormal reactions to conditional downgrades are significantly greater than reactions to unconditional changes when EJR follows earlier actions undertaken by investor-paid CRAs—the differences in the abnormal reactions triggered by conditional and unconditional rating actions fail to yield any strong evidence in the POST sample.

6. Conclusion

We have found empirical evidence supporting the argument that changes in the legislative and regulatory framework of the credit rating industry over the last decade have modified the lead-lag relationship between issuer-paid and investor-paid CRAs. Using a range of econometric tools, we find that the strong lead-lag relationship reported in previous papers (e.g., Johnson, 2004; BSS, 2006) has subsided, especially in the case of downgrades. The strong negative reputational effect of the untimely downgrading of a firm in financial distress has made issuer-paid agencies more responsive. Although we have evidence of improvements in timeliness by issuer-paid agencies, our marginal effect analysis proves that they are still conservative to the point that in an ordered probit framework there is strong evidence that downgrades by EJR may often increase the probability of future downgrades by two or more notches by one of the Big Three NRSROs, which may be taken as an indication of sluggish reaction to new information.

We provide evidence of instability or, equivalently, of breaks in the parameters of our vector auto-regressive and ordered probit models used to estimate the existence of causal links in credit rating changes. Our instability tests back our intuition on the sources of such instability, based on the differences in the lead-lag relationships before and after June 2002. More interesting is the fact that the tests often show evidence of instability usually three or four months after key events related to the modification of the legislation applied to CRAs.

We then perform a battery of classical event studies aimed at shedding light on the market perceptions on (the market value of) the lead-lag relationship between the two types of CRAs. In the case of unconditional ratings, the market reacts significantly to the upgrades and downgrades of the investor-paid CRA but only reacts significantly to the downgrades of the issuer-paid agencies. This suggests that the latter are only timely in the case of downgrades, while the former are always timely. We also find that, for unconditional downgrades, those issued by investor-paid CRAs imply significantly larger negative abnormal returns than those by issuer-paid agencies. This suggests that, even though our analysis implies that after 2002 causality may be manifesting itself in a bi-directional fashion, stock market abnormal returns show that investor-paid CRAs are still leading issuer-paid ones in an economic perspective. Moreover, abnormal stock market reactions to conditional downgrade actions—i.e., actions that are preceded by actions in the same direction in the previous 30 trading days—are significantly higher than reactions to unconditional downgrades. We suggest that the market may perceive that when an issuer-paid agency makes a conditional change, it is reacting to EJR, but on the contrary when EJR makes a conditional downgrade, the market assumes they are revealing new information.

Additional event studies separately performed on the PRE (1997-2002) and POST (2002-2006) regulatory overhaul sub-samples indicate that consistently with the breakpoint analysis applied to (vector auto-) regressive and ordered probit techniques, also finds instability in the strength and economic significance of event studies. In particular, while in the PRE sample stock market reactions to conditional actions (in particular, downgrades) are significantly greater than reactions to unconditional changes when EJR follows earlier actions by investor-paid CRAs, the differences in the reactions triggered by conditional vs. unconditional rating actions fail to yield any strong evidence in the POST sample.

Finally, the informational content of investor-paid CRAs diminishes when we incorporate outlook and watch lists but it is still statistically significant. This is consistent with our hypothesis that the lead relationships that link the changes in ratings by investor- to issuer-paid CRAs are stronger than those that link the outlooks and watch lists by investor- to issuer-paid agencies. Such weaker differences between the issuer- and the investor-paid models become even weaker after the 2002-2006 reforms.

Even though this evidence of a progressively shrinking divide between the historical NRSROs that abide to the traditional issuer-paid model and the new investor-paid model—here incarnated by Egan Jones Ratings—is suggestive of the fact that the wave of regulatory reforms undertaken in the U.S. after 2002 in the wake of the Enron and WorldCom scandals has been eventually successful, our paper has not reported conclusive evidence on either the optimality of the speed at which such previously documented differences have been vanishing or on the welfare implications of such a process of measured-pace "homogenization". In the light of the recent debate on the alleged shortcoming in CRAs conducts and practices during the Great Financial Crisis, many would be tempted to argue that the industry may need a further tightening of the regulations enforced, probably to be implemented at a faster pace than the 2002-2006 period investigated in this paper. Moreover, even though our paper has no evidence against or in favour of this conjecture, one may argue that when the issuer-paid CRAs start performing more similarly—in terms of lead-lag relationship or economic value of their announcements—to investor-paid CRAs, this might be due to the latter (especially in the case these are granted NRSRO status, as it occurred to EJR in late 2007) becoming increasingly less timely and accurate, given the same poor performance by issuer-paid CRAs. We leave this intriguing conjecture for future research (but see evidence in Bruno et al., 2012). However, our more benign interpretation of issuer-paid CRAs having become better at rating corporate bonds is consistent with the findings in Baghai et al. (2012) and Jorion et al. (2008) who have found that issuer-paid CRAs have tightened their standards over time and issued progressively less favourable ratings, which may be interpreted as indication of growing quality.

Finally, our evidence on market reactions to ratings undertaken by different types of CRAs has emphasized a few results consistent with alternative econometric methods, which we see as rather persuasive in the light of the differences between the nature of regression or probit models when compared to simpler event studies. Yet, only actual trading rules and profitability tests (i.e., computing the simulated, abnormal returns on a trading strategy based on conditional rating actions, under plausible transaction costs) may eventually give us a final understanding of the true, economic value of different rating models. Additionally, it would be interesting to see if the same results uncovered in this paper hold on bond market data. A series of bond event studies seems to be a natural extension of this research such as those performed in Weinstein (1977) or May (2010), in spite of the well-know issues with handling high-frequency bond return data and models (see e.g., Hotchkiss and Ronen, 2002).

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Table 1 Sample collection and reconciliation details for period 1997 - 2007

		Number of rating actions
"Investor-paid" credit rating agency (Egan Jones Ratings)		
Number of observations from EJR		24,800
Less: initial ratings, affirms and drops and outlooks		-19,784
Final "investor-paid" sample		5,016
"Issuer-paid" credit rating agencies		
Fitch, Moody's and S&P sample		28,875
Less: initial ratings, affirms, drops and outlooks		-23,357
Less: rating actions before the first action from "investor-paid" agencies		-628
Final "issuer-paid" sample		4,890
Aggregate sample with available PERMNO in CRSP		9,906
Pairs of "Investor-paid" and "Issuer-paid" agencies		
Firms rated by both EJR and Fitch		3,000
Rating changes by Fitch	751	
Rating changes by EJR	2,249	
Firms rated by both EJR and Moody's		4,741
Rating changes by Moody's	1462	
Rating changes by EJR	3,279	
Firms rated by both EJR and S&P		4,901
Rating changes by S&P	1,504	
Rating changes by EJR	3,397	

This table shows the steps followed in constructing the sample of changes in ratings for the investor-paid and issuer-paid credit rating agencies. The second part of our analysis that requires the inclusion of changes in outlooks, follows the same logic as above.

Table 2 Summary statistics for rating levels after a rating change of senior unsecured bonds.

	N	Mean	Std Dev	25% Percentile	Median	75% Percentile
EJR	2,249	10.05	3.58	8	10	12
Fitch	751	10.62	4.04	8	10	13
EJR	3,279	10.39	3.64	8	10	13
Moody's	1,462	11.35	3.86	9	11	14
EJR	3,397	10.38	3.62	8	10	13
S&P	1,504	11.42	4.32	8	11	14

Rating categories for all credit ratings agencies have been transformed to numerical ratings from 1-22 as explained in Appendix A. Numbers shown represent the rating level assigned by each rating agency after the rating action. A value of 10 is the lowest investment grade rating level across all rating agencies. A value of 11 is the highest non-investment grade rating level. The Investor-paid agency (EJR) is paired with each of the three issuer-paid agencies (Fitch, Moody's and S&P), by focusing only on the firms rated by both agencies.

Table 3 Distribution of rating levels after a rating change of senior unsecured bonds by each pair of Investor- and Issuer-paid rating agencies.

	Firms rated by both EJR and Fitch				Firms rated by both EJR and Moody's				Firms rated by both EJR and Standard & Poor's			
	Number of observations		Percentage of total		Number of observations		Percentage of total		Number of observations		Percentage of total	
Rating	EJR	Fitch	EJR	Fitch	EJR	Moody's	EJR	Moody's	EJR	S&P	EJR	S&P
1	0	3	0.0	0.4	1	5	0.0	0.3	1	6	0.0	0.4
2	1	4	0.0	0.5	2	8	0.1	0.5	2	4	0.1	0.3
3	22	5	1.0	0.7	31	10	0.9	0.7	30	12	0.9	0.8
4	50	16	2.2	2.1	63	19	1.9	1.3	67	22	2.0	1.5
5	101	19	4.5	2.5	127	23	3.9	1.6	133	49	3.9	3.3
6	153	49	6.8	6.5	191	56	5.8	3.8	199	72	5.9	4.8
7	233	70	10.4	9.3	290	92	8.8	6.3	296	101	8.7	6.7
8	261	76	11.6	10.1	350	128	10.7	8.8	362	127	10.7	8.4
9	289	95	12.9	12.6	400	147	12.2	10.1	414	155	12.2	10.3
10	264	79	11.7	10.5	390	195	11.9	13.3	407	175	12.0	11.6
11	200	65	8.9	8.7	330	129	10.1	8.8	341	143	10.0	9.5
12	162	61	7.2	8.1	260	134	7.9	9.2	272	90	8.0	6.0
13	138	56	6.1	7.5	230	105	7.0	7.2	239	93	7.0	6.2
14	110	36	4.9	4.8	176	102	5.4	7.0	185	99	5.4	6.6
15	102	21	4.5	2.8	159	85	4.8	5.8	165	84	4.9	5.6
16	67	24	3.0	3.2	119	58	3.6	4.0	123	71	3.6	4.7
17	8	18	0.4	2.4	8	57	0.2	3.9	8	55	0.2	3.7
18	36	19	1.6	2.5	63	37	1.9	2.5	63	38	1.9	2.5
19	0	5	0.0	0.7	43	32	1.3	2.2	0	30	0.0	2.0
20	27	10	1.2	1.3	28	33	0.9	2.3	44	23	1.3	1.5
21	14	10	0.6	1.3	18	7	0.5	0.5	28	16	0.8	1.1
22	11	10	0.5	1.3	0	0	0.0	0.0	18	39	0.5	2.6
Total	2,249	751	100	100	3,279	1,462	100	100	3,397	1,504	100	100

Pairs of agencies are formed between the investor-paid representative (EJR) and the three representatives of issuer-paid credit rating agencies. Ratings categories for all credit ratings agencies have been transformed to numerical ratings from 1-22 as explained in Appendix A. The investor-paid agency (EJR) is paired with each of the three issuer-paid agencies (Fitch, Moody's and S&P), by focusing only on the firms rated by the pair of agencies. Numbers shown represent the rating level assigned by each rating agency after a rating action. The highest rating by all agencies is 1 and the lowest is 22. A value of 10 is the lowest investment grade rating level across all rating agencies. A value of 11 is the highest non-investment grade rating level.

Table 4 Likelihood ratio tests of Granger-causality relationships between investor- and issuer-paid agencies.

Panel I: Fitch-EJR							
No Oulooks		1997-2002		2002-2007		1997-2007	
		EJR caused by Fitch	Fitch caused by EJR	EJR caused by Fitch	Fitch caused by EJR	EJR caused by Fitch	Fitch caused by EJR
Upgrades	LRT	1.58	18.68	13.14	48.37	14.16	57.48
	p-value	0.45	0.00	0.04	0.00	0.03	0.00
	Observations	953	913	2,045	2,007	3,000	2,956
Downgrades	LRT	7.80	27.30	14.68	46.38	10.33	68.28
	p-value	0.25	0.00	0.02	0.00	0.11	0.00
	Observations	955	955	2,045	2,045	3,000	3,000
With Outlooks		1999-2002		2002-2007		1999-2007	
		EJR caused by Fitch	Fitch caused by EJR	EJR caused by Fitch	Fitch caused by EJR	EJR caused by Fitch	Fitch caused by EJR
Upgrades	LRT	14.03	4.30	9.22	35.69	9.95	32.62
	p-value	0.00	0.64	0.16	0.00	0.13	0.00
	Observations	2,289	2,266	4,204	4,144	6,544	6,472
Downgrades	LRT	12.39	89.61	23.60	85.71	30.44	165.05
	p-value	0.05	0.00	0.00	0.00	0.00	0.00
	Observations	2,340	2,340	4,204	4,204	6,544	6,544
Panel II: Moody's-EJR							
No Outlooks		1997-2002		2002-2007		1997-2007	
		EJR caused by Moody's	Moody's caused by EJR	EJR caused by Moody's	Moody's caused by EJR	EJR caused by Moody's	Moody's caused by EJR
Upgrades	LRT	21.42	32.30	24.69	57.90	44.32	83.03
	p-value	0.00	0.00	0.00	0.00	0.00	0.00
	Observations	1,575	1,545	3,166	3,166	4,741	4,741
Downgrades	LRT	14.87	73.65	47.90	86.34	55.04	161.66
	p-value	0.02	0.00	0.00	0.00	0.00	0.00
	Observations	1,575	1,575	3,166	3,166	4,741	4,741
With Outlooks		1999-2002		2002-2007		1999-2007	
		EJR caused by Moody's	Moody's caused by EJR	EJR caused by Moody's	Moody's caused by EJR	EJR caused by Moody's	Moody's caused by EJR
Upgrades	LRT	14.21	6.04	5.80	37.01	12.46	36.11
	p-value	0.03	0.42	0.45	0.00	0.05	0.00
	Observations	3,541	3,383	6,167	6,167	9,708	9,708
Downgrades	LRT	39.31	189.65	51.83	151.43	83.48	324.77
	p-value	0.00	0.00	0.00	0.00	0.00	0.00
	Observations	3,541	3,541	6,167	6,167	9,708	9,708
Panel III: S&P-EJR							
No Outlooks		1997-2002		2002-2007		1997-2007	
		EJR caused by S&P	S&P caused by EJR	EJR caused by S&P	S&P caused by EJR	EJR caused by S&P	S&P caused by EJR
Upgrades	LRT	10.02	55.47	49.88	45.63	66.21	79.53
	p-value	0.12	0.00	0.00	0.00	0.00	0.00
	Observations	1,565	1,551	3,337	3,337	4,902	4,902
Downgrades	LRT	22.71	47.96	25.58	81.02	45.91	121.69
	p-value	0.00	0.00	0.00	0.00	0.00	0.00
	Observations	1,565	1,565	3,337	3,337	4,902	4,902
With Outlooks		1999-2002		2002-2007		1999-2007	
		EJR caused by S&P	S&P caused by EJR	EJR caused by S&P	S&P caused by EJR	EJR caused by S&P	S&P caused by EJR
Upgrades	LRT	11.50	4.44	14.30	24.90	15.18	22.07
	p-value	0.07	0.62	0.03	0.00	0.02	0.00
	Observations	3,435	3,346	6,186	6,186	9,621	9,621
Downgrades	LRT	43.15	135.57	39.19	159.96	74.07	284.89
	p-value	0.00	0.00	0.00	0.00	0.00	0.00
	Observations	3,435	3,435	6,186	6,186	9,621	9,621

The logistic regressions estimated in this table correspond to equations (1) to (4) of the paper. The log-likeleehood ratio test is calculated as follows: $LR=-2ln(L(UM)/L(RM))=2(ll(RM)-ll(UM))$ where: RM represents the restricted model, UM represents the unrestricted model, L() denotes the likelyhood of the respective model, and ll() the natural logarithms of the likelyhood of the models. The statistic is distributed chi-squared with degrees of freedom equal to the difference in number of degrees of freedom between the two models.

Table 5 Lead-lag relationship between "Issuer-paid" and "Investor-paid" credit rating agencies

Panel I: Lead-lag relationships between Fitch an Egan-Jones

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects %			
						<=-2	-1	1	>=2
Section I: Egan & Jones as rating follower									
Δ-Up by Fitch h=1	0.35	1.29		0.20	7.0	-4.5	-9.5	9.4	4.6
Δ-Up by Fitch h=2	-0.03	-0.12		0.91	0.6	0.5	0.7	-0.9	-0.3
Δ-Up by Fitch h=3	0.04	0.17		0.87	0.7	-0.6	-0.9	1.1	0.4
Δ-Up by Fitch h=4	0.17	0.42		0.67	3.4	-2.5	-4.4	4.9	2.0
Δ-Up by Fitch h=5	0.84	2.26	*	0.02	16.0	-7.5	-24.4	16.4	15.5
Δ-Up by Fitch h=6	0.36	1.53		0.13	7.2	-4.6	-9.9	9.6	4.8
Δ-Dw by Fitch h=1	-0.86	-5.23	**	0.00	14.0	21.8	6.1	-23.6	-4.4
Δ-Dw by Fitch h=2	-0.62	-4.65	**	0.00	10.8	14.1	7.5	-17.8	-3.8
Δ-Dw by Fitch h=3	-0.67	-3.75	**	0.00	11.5	15.9	7.1	-19.2	-3.9
Δ-Dw by Fitch h=4	-0.66	-4.40	**	0.00	11.3	15.5	7.2	-18.8	-3.8
Δ-Dw by Fitch h=5	0.23	1.11		0.27	4.6	-3.2	-6.0	6.5	2.8
Δ-Dw by Fitch h=6	-0.33	-1.71		0.09	6.1	6.5	5.7	-9.7	-2.5
Observations	2,216				Total	8.9	49.2	37.2	4.7
Section II: Fitch as rating follower									
Δ-Up by Egan & Jones h=1	0.48	2.78	**	0.01	8.9	-11.9	-5.9	10.3	7.5
Δ-Up by Egan & Jones h=2	0.79	3.00	**	0.00	15.0	-16.7	-13.2	14.9	15.1
Δ-Up by Egan & Jones h=3	0.59	2.06	*	0.04	11.1	-13.6	-8.5	12.1	10.1
Δ-Up by Egan & Jones h=4	0.94	3.70	**	0.00	17.9	-18.3	-17.5	16.1	19.7
Δ-Up by Egan & Jones h=5	0.23	0.93		0.35	4.1	-6.3	-1.9	5.1	3.0
Δ-Up by Egan & Jones h=6	0.72	3.03	**	0.00	13.7	-15.9	-11.5	14.2	13.3
Δ-Dw by Egan & Jones h=1	-0.90	-7.61	**	0.00	16.0	32.0	-8.4	-17.4	-6.2
Δ-Dw by Egan & Jones h=2	-0.40	-3.03	**	0.00	6.6	13.2	-1.3	-8.5	-3.4
Δ-Dw by Egan & Jones h=3	-0.29	-1.93		0.05	4.6	9.3	-0.4	-6.2	-2.6
Δ-Dw by Egan & Jones h=4	-0.60	-3.88	**	0.00	10.4	20.9	-4.1	-12.2	-4.5
Δ-Dw by Egan & Jones h=5	-0.63	-4.34	**	0.00	11.1	22.1	-4.6	-12.8	-4.7
Δ-Dw by Egan & Jones h=6	-0.36	-2.10	*	0.04	6.0	12.0	-1.1	-7.7	-3.1
Observations	731				Total	22.4	49.6	22.3	5.6

This table reports the results of ordered probit estimation of equations (7) and (8), using data from each pair of agencies. The sample period is 17 July 1997 to 21 December 2007. The dependent variables $\Delta R^{A}_{i,t}$ and $\Delta R^{B}_{i,t}$ represent comprehensive rating changes by EJR and Fitch, respectively, for firm i at time t. **,* denotes if the coefficients are statistically significant at 1 and 5 percent, respectively.

Table 5 (continued)

Panel II: Lead-lag relationships between Moody's and Egan-Jones

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects %			
						<=-2	-1	1	>=2
Section I: Egan-Jones as rating follower									
Δ-Up by Moody's h=1	0.55	4.11	**	0.00	10.9	-6.4	-15.3	13.2	8.5
Δ-Up by Moody's h=2	0.23	1.48		0.14	4.5	-3.3	-5.7	6.2	2.7
Δ-Up by Moody's h=3	0.21	1.00		0.32	4.2	-3.1	-5.3	5.9	2.5
Δ-Up by Moody's h=4	0.24	1.62		0.11	4.9	-3.5	-6.2	6.7	3.0
Δ-Up by Moody's h=5	0.20	0.96		0.34	4.0	-2.9	-5.0	5.5	2.4
Δ-Up by Moody's h=6	0.36	1.58		0.12	7.1	-4.7	-9.5	9.4	4.8
Δ-Dw by Moody's h=1	-0.85	-9.33	**	0.00	14.2	22.0	6.5	-23.9	-4.6
Δ-Dw by Moody's h=2	-0.58	-5.88	**	0.00	10.4	13.5	7.3	-17.1	-3.8
Δ-Dw by Moody's h=3	-0.51	-3.83	**	0.00	9.3	11.5	7.1	-15.1	-3.4
Δ-Dw by Moody's h=4	-0.32	-2.66	**	0.01	6.0	6.5	5.6	-9.6	-2.5
Δ-Dw by Moody's h=5	-0.29	-2.01	*	0.05	5.5	5.8	5.2	-8.7	-2.3
Δ-Dw by Moody's h=6	-0.23	-1.73		0.08	4.4	4.4	4.4	-6.9	-1.9
Observations	3,227				Total	9.4	47.3	38.4	4.9
Section II: Moody's as rating follower									
Δ-Up by Egan & Jones h=1	0.66	4.32	**	0.00	12.5	-15.1	-9.9	13.9	11.2
Δ-Up by Egan & Jones h=2	0.66	3.38	**	0.00	12.5	-15.0	-10.0	13.8	11.2
Δ-Up by Egan & Jones h=3	0.81	5.26	**	0.00	15.4	-17.1	-13.7	15.8	15.1
Δ-Up by Egan & Jones h=4	0.76	5.06	**	0.00	14.6	-16.6	-12.6	15.3	13.9
Δ-Up by Egan & Jones h=5	0.69	3.68	**	0.00	13.2	-15.6	-10.8	14.3	12.0
Δ-Up by Egan & Jones h=6	0.43	3.08	**	0.00	8.1	-11.0	-5.2	9.8	6.4
Δ-Dw by Egan & Jones h=1	-0.74	-8.10	**	0.00	13.0	26.0	-5.6	-15.4	-5.0
Δ-Dw by Egan & Jones h=2	-0.66	-6.74	**	0.00	11.5	23.1	-4.6	-13.9	-4.6
Δ-Dw by Egan & Jones h=3	-0.48	-4.74	**	0.00	8.2	16.4	-2.1	-10.6	-3.7
Δ-Dw by Egan & Jones h=4	-0.40	-3.76	**	0.00	6.8	13.6	-1.3	-9.1	-3.3
Δ-Dw by Egan & Jones h=5	-0.32	-2.79	**	0.01	5.3	10.5	-0.6	-7.2	-2.7
Δ-Dw by Egan & Jones h=6	-0.60	-5.56	**	0.00	10.6	21.2	-4.0	-12.9	-4.2
Observations	1,390				Total	22.8	48.3	23.7	5.2

Panel III: Lead-lag relationships between Standard & Poor's and Egan-Jones

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects %			
						<=-2	-1	1	>=2
Section I: Egan & Jones as rating follower									
Δ-Up by S&P h=1	0.91	6.61	**	0.00	16.9	-7.9	-26.0	15.8	18.0
Δ-Up by S&P h=2	0.41	2.47	*	0.01	8.0	-5.0	-11.1	10.3	5.8
Δ-Up by S&P h=3	0.72	5.63	**	0.00	13.9	-7.2	-20.6	15.0	12.7
Δ-Up by S&P h=4	0.28	1.64		0.10	5.6	-3.8	-7.4	7.6	3.6
Δ-Up by S&P h=5	0.54	3.14	**	0.00	10.5	-6.0	-15.0	12.7	8.3
Δ-Up by S&P h=6	0.42	2.03	*	0.04	8.2	-5.1	-11.4	10.5	5.9
Δ-Dw by S&P h=1	-0.81	-9.24	**	0.00	13.7	20.0	7.3	-22.7	-4.6
Δ-Dw by S&P h=2	-0.72	-6.70	**	0.00	12.4	17.4	7.5	-20.6	-4.3
Δ-Dw by S&P h=3	-0.44	-3.48	**	0.00	8.2	9.3	7.0	-13.1	-3.2
Δ-Dw by S&P h=4	-0.53	-4.23	**	0.00	9.6	11.7	7.4	-15.6	-3.6
Δ-Dw by S&P h=5	-0.24	-1.86		0.06	4.6	4.5	4.6	-7.1	-2.0
Δ-Dw by S&P h=6	-0.31	-2.40	*	0.02	5.8	6.0	5.6	-9.1	-2.5
Observations	3,344				Total	9.0	47.7	38.3	5.0
Section II: S&P as rating follower									
Δ-Up by Egan & Jones h=1	0.43	2.54	*	0.01	7.9	-10.5	-5.3	9.9	6.0
Δ-Up by Egan & Jones h=2	0.39	2.09	*	0.04	7.1	-9.7	-4.6	9.0	5.3
Δ-Up by Egan & Jones h=3	0.61	3.53	**	0.00	11.5	-13.7	-9.3	13.4	9.6
Δ-Up by Egan & Jones h=4	0.66	4.26	**	0.00	12.4	-14.3	-10.4	14.1	10.6
Δ-Up by Egan & Jones h=5	0.83	4.59	**	0.00	15.8	-16.5	-15.1	16.5	15.1
Δ-Up by Egan & Jones h=6	0.76	7.29	**	0.00	14.3	-15.7	-13.0	15.6	13.1
Δ-Dw by Egan & Jones h=1	-0.68	-7.68	**	0.00	11.6	23.2	-4.8	-14.1	-4.4
Δ-Dw by Egan & Jones h=2	-0.47	-4.53	**	0.00	7.9	15.8	-2.1	-10.3	-3.4
Δ-Dw by Egan & Jones h=3	-0.45	-4.68	**	0.00	7.5	15.1	-1.9	-9.9	-3.3
Δ-Dw by Egan & Jones h=4	-0.44	-4.81	**	0.00	7.2	14.4	-1.6	-9.6	-3.2
Δ-Dw by Egan & Jones h=5	-0.39	-3.07	**	0.00	6.3	12.7	-1.2	-8.6	-2.9
Δ-Dw by Egan & Jones h=6	-0.41	-3.39	**	0.00	6.7	13.3	-1.4	-8.9	-3.0
Observations	1,450				Total	21.4	51.1	22.7	4.8

This table reports the results of ordered probit estimation of equations (7) and (8), using data from each pair of agencies. The sample period is 17 July 1997 to 21 December 2007. The dependent variables $\Delta R^{A}_{i,t}$ and $\Delta R^{B}_{i,t}$ represent rating changes by EJR and Moody's or EJR and S&P, respectively, for firm i at time t. **,* denotes if the coefficients are statistically significant at 1 and 5 percent, respectively.

Table 6 Lead-lag relationship between "issuer-paid" and "investor-paid" credit rating agencies including outlooks and watch list data

Panel I: Lead-lag relationships between Egan-Jones and Fitch

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects %							
						<=-4	-3	-2	-1	1	2	3	>=4
Section I: Egan-Jones as rating follower													
Δ-Up by Fitch h=1	0.11	0.47		0.64	1.1	-1.1	-0.8	-1.7	-0.9	1.0	1.8	0.7	1.0
Δ-Up by Fitch h=2	0.34	1.77		0.08	3.4	-2.6	-2.2	-5.2	-3.4	2.3	5.3	2.3	3.6
Δ-Up by Fitch h=3	-0.09	-0.43		0.67	0.9	1.0	0.7	1.3	0.5	-1.0	-1.4	-0.5	-0.6
Δ-Up by Fitch h=4	0.27	1.50		0.13	2.7	-2.2	-1.8	-4.1	-2.5	2.0	4.2	1.8	2.6
Δ-Up by Fitch h=5	0.40	2.12	*	0.03	4.0	-3.0	-2.5	-6.1	-4.3	2.4	6.2	2.8	4.5
Δ-Up by Fitch h=6	0.27	1.68		0.09	2.7	-2.2	-1.8	-4.2	-2.6	2.0	4.3	1.8	2.7
Δ-Dw by Fitch h=1	-0.63	-5.73	**	0.00	5.9	10.4	5.4	8.0	-0.9	-8.7	-8.6	-2.7	-2.9
Δ-Dw by Fitch h=2	-0.58	-4.91	**	0.00	5.4	9.1	4.9	7.5	-0.5	-7.8	-8.0	-2.5	-2.8
Δ-Dw by Fitch h=3	-0.49	-3.97	**	0.00	4.6	7.4	4.1	6.7	0.1	-6.5	-7.0	-2.2	-2.5
Δ-Dw by Fitch h=4	-0.71	-6.47	**	0.00	6.7	12.2	6.0	8.5	-1.7	-9.8	-9.3	-2.8	-3.0
Δ-Dw by Fitch h=5	0.11	0.65		0.51	1.1	-1.1	-0.8	-1.8	-0.9	1.0	1.8	0.7	1.0
Δ-Dw by Fitch h=6	-0.25	-1.71		0.09	2.4	3.2	2.0	3.7	0.8	-3.1	-3.8	-1.3	-1.6
Observations	4,511				Total	5.0	5.2	17.1	27.5	23.5	14.6	3.7	3.6
Section II: Fitch as rating follower													
Δ-Up by Egan & Jones h=1	0.46	3.17	**	0.00	4.5	-9.6	-2.3	-6.3	0.5	1.9	9.5	0.5	5.8
Δ-Up by Egan & Jones h=2	0.64	3.18	**	0.00	6.2	-11.9	-3.0	-9.5	-0.1	2.0	12.7	0.7	9.2
Δ-Up by Egan & Jones h=3	0.51	2.01	*	0.04	5.0	-10.2	-2.5	-7.3	0.3	1.9	10.4	0.5	6.8
Δ-Up by Egan & Jones h=4	0.46	2.95	**	0.00	4.6	-9.5	-2.3	-6.5	0.4	1.8	9.5	0.5	5.9
Δ-Up by Egan & Jones h=5	0.40	2.33	*	0.02	4.0	-8.5	-2.0	-5.4	0.6	1.7	8.3	0.4	4.9
Δ-Up by Egan & Jones h=6	0.73	4.13	**	0.00	7.1	-13.1	-3.4	-11.3	-0.6	2.0	14.2	0.8	11.3
Δ-Dw by Egan & Jones h=1	-0.81	-7.67	**	0.00	7.3	26.0	2.8	0.3	-5.6	-5.1	-13.7	-0.5	-4.4
Δ-Dw by Egan & Jones h=2	-0.56	-5.31	**	0.00	5.2	17.0	2.2	1.6	-3.5	-3.5	-10.0	-0.4	-3.4
Δ-Dw by Egan & Jones h=3	-0.41	-3.32	**	0.00	3.9	12.0	1.7	1.8	-2.4	-2.5	-7.6	-0.3	-2.7
Δ-Dw by Egan & Jones h=4	-0.43	-3.36	**	0.00	4.1	12.8	1.8	1.7	-2.6	-2.7	-8.0	-0.3	-2.8
Δ-Dw by Egan & Jones h=5	-0.47	-4.06	**	0.00	4.5	14.2	2.0	1.7	-2.9	-3.0	-8.7	-0.3	-3.0
Δ-Dw by Egan & Jones h=6	-0.38	-2.83	**	0.01	3.6	11.1	1.6	1.7	-2.2	-2.4	-7.1	-0.3	-2.5
Observations	1,067				Total	17.4	5.8	30.7	14.4	9.2	17.9	0.5	4.1

This table reports the results of ordered probit estimation of equations (7) and (8), using data from each pair of agencies. The sample period is 17 July 1997 to 21 December 2007. The dependent variables $\Delta R^{A}_{i,t}$ and $\Delta R^{B}_{i,t}$ represent comprehensive rating changes (rating changes plus watchlist or outlooks) by EJR and Fitch respectively for firm i at time t. **,* denotes if the coefficients are statistically significant at 1 and 5 percent, respectively.

Table 6 (continued)

Panel II: Lead-lag relationships between Egan-Jones and Moody's

	Coefficients	z-statistic		p-value	Average |Change	<=-4	-3	-2	-1	1	2	3	>=4
						Marginal effects %							
Section I: Egan-Jones as rating follower													
Δ-Up by Moody's h=1	0.50	3.67	**	0.00	4.9	-4.0	-2.9	-7.2	-5.3	2.3	7.2	3.7	6.2
Δ-Up by Moody's h=2	0.36	2.62	**	0.01	3.6	-3.3	-2.3	-5.3	-3.5	2.2	5.4	2.6	4.1
Δ-Up by Moody's h=3	0.26	1.71		0.09	2.5	-2.5	-1.7	-3.8	-2.2	1.9	3.9	1.8	2.6
Δ-Up by Moody's h=4	0.19	1.32		0.19	1.8	-1.9	-1.2	-2.8	-1.5	1.5	2.8	1.3	1.8
Δ-Up by Moody's h=5	0.14	0.96		0.34	1.4	-1.5	-1.0	-2.1	-1.0	1.2	2.2	0.9	1.3
Δ-Up by Moody's h=6	0.36	2.22	*	0.03	3.6	-3.2	-2.2	-5.3	-3.4	2.2	5.4	2.6	4.0
Δ-Dw by Moody's h=1	-0.70	-9.66	**	0.00	6.6	13.2	5.5	7.8	-1.5	-9.5	-9.1	-3.0	-3.3
Δ-Dw by Moody's h=2	-0.61	-6.79	**	0.00	5.8	11.1	4.8	7.2	-0.9	-8.3	-8.2	-2.8	-3.0
Δ-Dw by Moody's h=3	-0.36	-3.80	**	0.00	3.5	5.6	2.8	4.9	0.5	-4.6	-5.2	-1.9	-2.2
Δ-Dw by Moody's h=4	-0.39	-4.37	**	0.00	3.7	6.1	3.0	5.1	0.5	-4.9	-5.5	-2.0	-2.3
Δ-Dw by Moody's h=5	-0.26	-2.82	**	0.01	2.6	3.8	2.0	3.7	0.7	-3.2	-3.9	-1.5	-1.7
Δ-Dw by Moody's h=6	-0.15	-1.58		0.11	1.5	2.0	1.1	2.2	0.6	-1.7	-2.3	-0.9	-1.1
Observations	6,233				Total	6.0	5.3	17.1	26.0	23.5	14.4	4.0	3.8
Section II: Moody's as rating follower													
Δ-Up by Egan & Jones h=1	0.75	6.15	**	0.00	7.2	-10.9	-5.0	-9.7	-3.4	5.0	10.2	3.6	10.2
Δ-Up by Egan & Jones h=2	0.31	2.34	*	0.02	3.0	-5.8	-2.3	-3.7	0.0	3.0	4.4	1.3	3.0
Δ-Up by Egan & Jones h=3	0.64	4.49	**	0.00	6.2	-9.9	-4.4	-8.3	-2.3	4.7	8.9	3.0	8.1
Δ-Up by Egan & Jones h=4	0.72	5.64	**	0.00	6.9	-10.6	-4.8	-9.3	-3.1	4.9	9.8	3.5	9.5
Δ-Up by Egan & Jones h=5	0.57	3.77	**	0.00	5.5	-9.2	-4.0	-7.2	-1.6	4.5	8.0	2.6	6.8
Δ-Up by Egan & Jones h=6	0.34	2.40	*	0.02	3.2	-6.3	-2.5	-4.1	-0.1	3.2	4.8	1.5	3.4
Δ-Dw by Egan & Jones h=1	-0.69	-9.45	**	0.00	6.7	19.5	4.3	3.1	-6.4	-7.9	-7.6	-1.8	-3.2
Δ-Dw by Egan & Jones h=2	-0.59	-7.56	**	0.00	5.8	16.5	3.8	3.0	-5.3	-6.9	-6.7	-1.6	-2.9
Δ-Dw by Egan & Jones h=3	-0.43	-4.78	**	0.00	4.2	11.2	2.9	2.8	-3.3	-4.9	-5.1	-1.2	-2.3
Δ-Dw by Egan & Jones h=4	-0.49	-5.56	**	0.00	4.8	13.2	3.2	2.9	-4.1	-5.7	-5.7	-1.4	-2.5
Δ-Dw by Egan & Jones h=5	-0.44	-4.85	**	0.00	4.3	11.7	3.0	2.8	-3.5	-5.1	-5.2	-1.3	-2.3
Δ-Dw by Egan & Jones h=6	-0.47	-5.56	**	0.00	4.6	12.5	3.1	2.8	-3.8	-5.4	-5.5	-1.3	-2.4
Observations	2,107				Total	13.8	7.9	21.2	25.8	15.2	10.7	2.1	3.3

Panel III: Lead-lag relationships between Egan-Jones and Standard & Poor's

	Coefficients	z-statistic		p-value	Average |Change	<=-4	-3	-2	-1	1	2	3	>=4
						Marginal effects %							
Section I: Egan-Jones as rating follower													
Δ-Up by S&P h=1	0.62	5.38	**	0.00	6.0	-4.5	-3.4	-8.8	-7.2	1.8	8.5	4.7	8.9
Δ-Up by S&P h=2	0.39	2.89	**	0.00	3.8	-3.4	-2.4	-5.7	-3.8	2.2	5.7	2.8	4.6
Δ-Up by S&P h=3	0.53	3.95	**	0.00	5.2	-4.1	-3.1	-7.7	-5.8	2.1	7.5	4.0	7.1
Δ-Up by S&P h=4	0.23	1.13		0.26	2.2	-2.2	-1.5	-3.4	-1.9	1.7	3.4	1.6	2.4
Δ-Up by S&P h=5	0.22	1.63		0.10	2.2	-2.2	-1.5	-3.3	-1.9	1.7	3.4	1.5	2.3
Δ-Up by S&P h=6	0.07	0.46		0.64	0.7	-0.8	-0.5	-1.1	-0.5	0.6	1.1	0.5	0.7
Δ-Dw by S&P h=1	-0.73	-10.34	**	0.00	6.9	13.9	5.7	8.0	-1.8	-9.9	-9.4	-3.1	-3.5
Δ-Dw by S&P h=2	-0.53	-6.07	**	0.00	5.0	9.1	4.2	6.6	-0.2	-7.0	-7.3	-2.5	-3.0
Δ-Dw by S&P h=3	-0.36	-4.01	**	0.00	3.4	5.5	2.8	4.8	0.6	-4.4	-5.1	-1.9	-2.3
Δ-Dw by S&P h=4	-0.47	-4.98	**	0.00	4.4	7.8	3.7	6.0	0.1	-6.0	-6.5	-2.3	-2.7
Δ-Dw by S&P h=5	-0.23	-2.43	*	0.02	2.2	3.2	1.7	3.2	0.7	-2.6	-3.3	-1.3	-1.6
Δ-Dw by S&P h=6	-0.26	-2.79	**	0.01	2.5	3.7	2.0	3.6	0.7	-3.0	-3.8	-1.4	-1.8
Observations	6,348				Total	5.9	5.2	17.1	25.9	23.4	14.5	4.0	4.0
Section II: S&P as rating follower													
Δ-Up by Egan & Jones h=1	0.45	3.22	**	0.00	4.4	-8.8	-3.2	-5.6	0.8	2.7	7.2	1.6	5.3
Δ-Up by Egan & Jones h=2	0.44	3.19	**	0.00	4.3	-8.6	-3.1	-5.4	0.8	2.6	7.0	1.5	5.2
Δ-Up by Egan & Jones h=3	0.54	3.77	**	0.00	5.2	-10.0	-3.7	-7.0	0.4	3.0	8.6	1.9	6.9
Δ-Up by Egan & Jones h=4	0.66	4.72	**	0.00	6.3	-11.5	-4.5	-9.0	-0.2	3.2	10.3	2.4	9.2
Δ-Up by Egan & Jones h=5	0.60	3.71	**	0.00	5.8	-10.8	-4.1	-8.1	0.1	3.1	9.5	2.2	8.1
Δ-Up by Egan & Jones h=6	0.53	3.47	**	0.00	5.1	-9.8	-3.7	-6.8	0.5	2.9	8.4	1.9	6.7
Δ-Dw by Egan & Jones h=1	-0.65	-8.37	**	0.00	6.1	19.7	3.6	1.2	-5.9	-5.1	-8.6	-1.4	-3.5
Δ-Dw by Egan & Jones h=2	-0.50	-5.63	**	0.00	4.8	14.8	3.0	1.5	-4.4	-4.0	-6.9	-1.1	-2.9
Δ-Dw by Egan & Jones h=3	-0.37	-4.26	**	0.00	3.6	10.3	2.3	1.7	-2.9	-2.9	-5.2	-0.9	-2.3
Δ-Dw by Egan & Jones h=4	-0.49	-5.88	**	0.00	4.7	14.2	2.9	1.6	-4.2	-3.8	-6.7	-1.1	-2.8
Δ-Dw by Egan & Jones h=5	-0.42	-4.45	**	0.00	4.0	12.0	2.6	1.6	-3.5	-3.3	-5.9	-1.0	-2.5
Δ-Dw by Egan & Jones h=6	-0.44	-4.68	**	0.00	4.2	12.7	2.7	1.6	-3.7	-3.5	-6.1	-1.0	-2.6
Observations	1,969				Total	16.2	8.3	26.9	19.7	10.4	13.0	1.7	3.8

This table reports the results of ordered probit estimation of equations (7) and (8), using data from each pair of agencies. The sample period is 17 July 1997 to 21 December 2007. The dependent variables $\Delta R^A_{i,t}$ and $\Delta R^B_{i,t}$ represent comprehensive rating changes (rating changes plus watchlist or outlooks) by EJR and Moody's or EJR and S&P, respectively, for firm i at time t. **,* denotes if the coefficients are statistically significant at 1 and 5 percent, respectively.

Table 7 Log-likelihood ratio test for breaks in the coefficients of the ordered probit model

	Panel 1: Fitch-EJR pair Potential rating follower		Panel 2: Moody's-EJR pair Potential rating follower		Panel 3: S&P-EJR pair Potential rating follower	
	Fitch	EJR	Moody's	EJR	S&P	EJR
Break date	LR test	LR test	LR test	LR test	LR test	LR test
01/01/2001	16.50	13.34	11.53	9.24	11.89	13.34
01/04/2001	12.38	13.52	16.71	13.65	14.01	13.52
01/07/2001	12.92	8.99	15.60	13.19	10.21	8.99
01/10/2001	15.16	9.21	15.20	13.21	12.07	9.21
01/01/2002	14.23	8.63	14.21	13.19	14.84	8.63
01/04/2002	16.89	10.70	16.92	13.86	20.58	10.70
01/07/2002	16.65	14.41	16.47	13.40	18.48	14.41
01/10/2002	20.18	16.83	20.00	17.07	21.51	16.83
01/01/2003	21.25	18.72	21.96	20.63	20.86	18.72
01/04/2003	22.58	18.46	24.00	21.87	19.36	18.46
01/07/2003	25.15 ***	11.48	27.40 ***	18.76	23.20	11.48
01/10/2003	16.47	10.06	25.00	19.76	21.82	10.06
01/01/2004	17.39	5.30	22.40	22.61	15.07	5.30
01/04/2004	11.20	4.70	22.66	23.24	17.09	4.70
01/07/2004	12.08	7.41	23.58	26.55 ***	16.25	7.41
01/10/2004	13.14	4.70	19.79	25.28	14.93	4.70
01/01/2005	15.91	4.72	19.80	24.36	15.05	4.72
01/04/2005	15.28	6.03	19.23	24.44	15.09	6.03
01/07/2005	14.36	7.95	18.43	23.46	17.49	7.95
01/10/2005	13.54	6.72	22.73	25.90	19.48	6.72
01/01/2006	14.30	9.36	25.50	11.92	20.21	9.36
01/04/2006	18.06	12.77	19.13	19.11	20.70	12.77
01/07/2006	17.62	18.77	17.06	20.04	24.01	18.77
01/10/2006	16.18	16.33	12.44	13.10	24.24	16.33
01/01/2007	14.52	12.72	13.99	8.00	31.48 ***	12.72
01/04/2007	18.30	10.36	19.83	8.58	28.43 ***	10.36
01/07/2007	16.98	12.75	16.80	7.60	7.41	12.75
01/10/2007	7.07	3.44	17.14	1.66	5.16	3.44

The ordered probit regressions estimated in this table correspond to equations (10) and (11) in the paper. The log-likeleehood ratio test is calculated as follows: $LR = -2ln\left(\frac{L(UM)}{L(RM)}\right) = 2\left(ll(RM) - ll(UM)\right)$ where: RM represents the restrictec model (without interaction with the break dummy), UM represents the unrestricted model (with interaction with a break dummy), L() denotes the likelihood of the respective model, and ll() represents the natural logarithms of the likelihoods of the models. The statistic is distributed chi-squared with degrees of freedom equal to the difference in number of degrees of freedom between the two models. *** denotes that the LR test is significant at the 1 percent level.

Table 8 Comparison between abnormal stock returns of conditional and unconditional rating announcements: Full sample

Panel I	Type	Credit rating agency pair	Abnormal Returns			Variance of abnormal returns	N	mean difference T-test	p-value
Downgrades	Conditional	EJR | Moody's	-7.41%	***	(ΔΔΔ)	0.018	133	-3.12	0.00
	Unconditional		-3.74%	***	(ΔΔΔ)	0.010	1,720		
	Conditional	Moody's | EJR	-2.25%	***	(ΔΔΔ)	0.008	160	-0.96	0.25
	Unconditional		-1.48%	***	(ΔΔΔ)	0.009	774		
	Conditional	EJR | S&P	-4.52%	***	(ΔΔ)	0.012	134	-0.55	0.34
	Unconditional		-3.98%	***	(ΔΔΔ)	0.010	1,782		
	Conditional	S&P | EJR	-1.70%	**	(ΔΔ)	0.010	150	0.65	0.32
	Unconditional		-2.28%	***	(ΔΔΔ)	0.012	878		
	Conditional	EJR | Fitch	-7.50%	***	(Δ)	0.023	58	-1.84	0.07
	Unconditional		-3.82%	***	(ΔΔ)	0.009	1,238		
	Conditional	Fitch | EJR	-2.62%	*	(Δ)	0.023	99	-0.24	0.39
	Unconditional		-2.23%	***	(ΔΔ)	0.014	392		
Panel II									
Upgrades	Conditional	EJR | Moody's	1.40%	**	(ΔΔ)	0.001	29	-1.93	0.06
	Unconditional		2.58%	***	(ΔΔΔ)	0.004	1,428		
	Conditional	Moody's | EJR	-0.39%		(ΔΔ)	0.000	33	-1.99	0.06
	Unconditional		0.16%		(ΔΔΔ)	0.000	387		
	Conditional	EJR | S&P	-0.23%			0.002	23	-3.04	0.01
	Unconditional		2.60%	***	(ΔΔΔ)	0.004	1,490		
	Conditional	S&P | EJR	0.82%	**		0.000	33	1.88	0.07
	Unconditional		0.00%		(ΔΔΔ)	0.001	388		
	Conditional	EJR | Fitch	-0.37%			0.003	8	-1.34	0.15
	Unconditional		2.11%	***	(ΔΔΔ)	0.003	960		
	Conditional	Fitch | EJR	1.10%	**		0.001	28	2.26	0.03
	Unconditional		-0.04%		(ΔΔΔ)	0.000	201		

This table reports the abnormal stock returns of the market to credit rating announcements. The event window is set to correspond to the announcement day. The benchmark is the CRSP equally weighted index. The sample period is 17 July 1997 to 21 December 2007. ***, **, * denote p-values from a two-sided test of zero mean abnormal returns being statistically significant different from zero at 1, 5 and 10 percent, respectively. ΔΔΔ, Δ, Δ denote p-values from a mean difference t-test for abnormal returns, between agencies in a pair, being statistically significant different from zero at 1, 5 and 10 percent, respectivelly. Mean difference t-test evaluates whether the difference in mean abnormal returns between conditional and uncoditional event studies for the same pair of credit rating agencies is statistically significant.

Table 9 Comparison between abnormal stock returns of conditional and unconditional rating announcements: Pre sample

Panel I	Type	Credit rating agency pair	Abnormal Returns			Variance of abnormal returns	N	mean difference T-test	p-value
Downgrades	Conditional	EJR - Moody's	-7.71%	***	(ΔΔ)	0.016	65	-2.32	0.03
	Unconditional		-3.96%	***	(ΔΔ)	0.009	713		
	Conditional	Moody's - EJR	-2.56%	**	(ΔΔ)	0.011	73	-0.38	0.37
	Unconditional		-2.04%	***	(ΔΔ)	0.013	291		
	Conditional	EJR - S&P	-5.19%	***		0.015	62	-0.56	0.34
	Unconditional		-4.30%	***	(ΔΔ)	0.010	736		
	Conditional	S&P - EJR	-3.67%	**		0.011	59	-0.69	0.31
	Unconditional		-2.59%	***	(ΔΔ)	0.015	275		
	Conditional	EJR - Fitch	-10.38%	***		0.025	22	-1.81	0.08
	Unconditional		-4.28%	***		0.010	527		
	Conditional	Fitch - EJR	-4.39%			0.033	42	-0.20	0.39
	Unconditional		-3.78%	***		0.018	118		
Panel II									
Upgrades	Conditional	EJR - Moody's	1.60%			0.002	7	-0.98	0.23
	Unconditional		3.27%	***	(ΔΔΔ)	0.005	331		
	Conditional	Moody's - EJR	-0.61%			0.001	5	-0.66	0.29
	Unconditional		0.12%		(ΔΔΔ)	0.001	81		
	Conditional	EJR - S&P	0.13%			0.000	2	-6.74	0.00
	Unconditional		3.44%	***	(ΔΔΔ)	0.006	346		
	Conditional	S&P - EJR	2.73%			0.001	5	1.32	0.15
	Unconditional		0.49%	*	(ΔΔΔ)	0.001	71		
	Conditional	EJR - Fitch	2.80%			0.000	1	0.26	-
	Unconditional		2.68%	***	(ΔΔΔ)	0.005	232		
	Conditional	Fitch - EJR	9.29%			0.000	1	21.00	-
	Unconditional		-0.97%	*	(ΔΔΔ)	0.001	29		

This table reports the abnormal stock returns of the market to credit rating announcements. The event window is set to correspond to the announcement day. The benchmark is the CRSP equally weighted index. The sample period is 17 July 1997 to 31 May 2002. ***, **, * denote p-values from a two-sided test of zero mean abnormal returns being statistically significant different from zero at 1, 5 and 10 percent, respectivelly. ΔΔΔ, Δ, Δ denote p-values from a mean difference t-test for abnormal returns, between agencies in a pair, being statistically significant different from zero at 1, 5 and 10 percent, respectively. Mean difference t-test evaluates whether the difference in mean abnormal returns between conditional and uncoditional event studies for the same pair of credit rating agencies is statistically significant.

Table 10 Comparison between abnormal stock returns of conditional and unconditional rating announcements: Post sample

Panel I	Type	Credit rating agency pair	Abnormal Returns			Variance of abnormal returns	N	mean difference T-test	p-value
Downgrades	Conditional	EJR - Moody's	-7.13%	***	(ΔΔ)	0.019	68	-2.06	0.05
	Unconditional		-3.59%	***	(ΔΔΔ)	0.010	1,007		
	Conditional	Moody's - EJR	-1.99%	**	(ΔΔ)	0.006	87	-0.91	0.26
	Unconditional		-1.14%	***	(ΔΔΔ)	0.006	483		
	Conditional	EJR - S&P	-3.95%	***	(ΔΔ)	0.010	72	-0.16	0.39
	Unconditional		-3.75%	***	(ΔΔΔ)	0.010	1,046		
	Conditional	S&P - EJR	-0.43%		(ΔΔ)	0.009	91	1.61	0.11
	Unconditional		-2.14%	***	(ΔΔΔ)	0.011	603		
	Conditional	EJR - Fitch	-5.73%	**		0.021	36	-0.92	0.26
	Unconditional		-3.48%	***	(ΔΔ)	0.009	711		
	Conditional	Fitch - EJR	-1.32%			0.016	57	0.14	0.39
	Unconditional		-1.57%	**	(ΔΔ)	0.012	274		
Panel II									
Upgrades	Conditional	EJR - Moody's	1.33%	**	(ΔΔ)	0.001	22	-1.66	0.10
	Unconditional		2.37%	***	(ΔΔΔ)	0.004	1,097		
	Conditional	Moody's - EJR	-0.35%		(ΔΔ)	0.000	28	-1.95	0.06
	Unconditional		0.18%	*	(ΔΔΔ)	0.000	306		
	Conditional	EJR - S&P	-0.26%			0.002	21	-2.56	0.02
	Unconditional		2.34%	***	(ΔΔΔ)	0.004	1,144		
	Conditional	S&P - EJR	0.47%			0.000	28	1.46	0.14
	Unconditional		-0.11%		(ΔΔΔ)	0.002	317		
	Conditional	EJR - Fitch	-0.82%			0.003	7	-1.33	0.15
	Unconditional		1.93%	***	(ΔΔΔ)	0.003	728		
	Conditional	Fitch - EJR	0.80%	*		0.000	27	1.62	0.11
	Unconditional		0.12%		(ΔΔΔ)	0.000	172		

This table reports the abnormal stock returns of the market to credit rating announcements. The event window is set to correspond to the announcement day. The benchmark is the CRSP equally weighted index. The sample period is 01 June 2002 to 21 December 2007. ***, **, * denote p-values from a two-sided test of zero mean abnormal returns being statistically significant different from zero at 1, 5 and 10 percent, respectively. ΔΔΔ, Δ, Δ denote p-values from a mean difference t-test for abnormal returns, between agencies in a pair, being statistically significant different from zero at 1, 5 and 10 percent, respectivelly. Mean difference t-test evaluates whether the difference in mean abnormal returns between conditional and uncoditional event studies for the same pair of credit rating agencies is statistically significant.

Table 11 Comparison between abnormal stock returns of conditional and unconditional rating announcements including outlooks and watch list from 17 July 1997 to 21 December 2007

Panel I	Type	Credit rating agency pair	Abnormal Returns			Variance of abnormal returns	N	mean difference T-test	p-value
Downgrades	Conditional	EJR | Moody's	-7.34%	***	(ΔΔΔ)	0.018	158	-3.28	0.00
	Unconditional		-3.74%	***	(ΔΔΔ)	0.011	2041		
	Conditional	Moody's | EJR	-2.94%	***	(ΔΔΔ)	0.020	188	-1.39	0.15
	Unconditional		-1.44%	***	(ΔΔΔ)	0.006	727		
	Conditional	EJR | S&P	-6.46%	***	(ΔΔ)	0.020	160	-2.28	0.03
	Unconditional		-3.89%	***	(ΔΔΔ)	0.011	2072		
	Conditional	S&P | EJR	-3.03%	**	(ΔΔ)	0.022	174	-0.54	0.34
	Unconditional		-2.40%	***	(ΔΔΔ)	0.009	798		
	Conditional	EJR | Fitch	-9.72%	***	(ΔΔ)	0.029	78	-3.19	0.00
	Unconditional		-3.52%	***	(ΔΔΔ)	0.010	1,493		
	Conditional	Fitch | EJR	-4.25%	**	(ΔΔ)	0.035	115	-1.56	0.12
	Unconditional		-1.43%	***	(ΔΔΔ)	0.008	356		
Panel II									
Upgrades	Conditional	EJR | Moody's	1.06%	*	(ΔΔ)	0.001	29	-2.80	0.01
	Unconditional		2.58%	***	(ΔΔΔ)	0.004	1,726		
	Conditional	Moody's | EJR	-0.30%		(ΔΔ)	0.000	34	-2.01	0.06
	Unconditional		0.24%	*	(ΔΔΔ)	0.001	344		
	Conditional	EJR | S&P	0.83%			0.002	27	-1.82	0.08
	Unconditional		2.59%	***	(ΔΔΔ)	0.005	1,771		
	Conditional	S&P | EJR	0.70%			0.001	26	1.13	0.21
	Unconditional		0.18%		(ΔΔΔ)	0.000	339		
	Conditional	EJR | Fitch	-0.55%			0.002	10	-1.94	0.07
	Unconditional		2.29%	***	(ΔΔΔ)	0.004	1,220		
	Conditional	Fitch | EJR	1.06%	**		0.001	30	2.10	0.05
	Unconditional		0.01%		(ΔΔΔ)	0.000	176		

This table reports the abnormal stock returns of the market to credit rating announcements. The event window is set to correspond to the announcement day. The benchmark is the CRSP equally weighted index. The sample period is 17 July 1997 to 21 December 2007. ***, **, * denote p-values from a two-sided test of zero mean abnormal returns being statistically significant different from zero at 1, 5 and 10 percent, respectively. ΔΔΔ, Δ, Δ denote p-values from a mean difference t-test for abnormal returns, between agencies in a pair, being statistically significant different from zero at 1, 5 and 10 percent, respectivelly. Mean difference t-test evaluates whether the difference in mean abnormal returns between conditional and uncoditional event studies for the same pair of credit rating agencies is statistically significant.

Appendix A

Table A1 Credit rating scales for the credit rating (CR) and comprehensive credit rating (CCR) variables

Panel I: Credit rating (CR) scale

Numerical rating	Moody's	EJR, Fitch and S&P	Numerical rating	Moody's	Ejr, Fitch and S&P
Investment grade			Non-investment grade		
1	Aaa	AAA	11	Ba1	BB+
2	Aa1	AA+	12	Ba2	BB
3	Aa2	AA	13	Ba3	BB-
4	Aa3	AA-	14	B1	B+
5	A1	A+	15	B2	B
6	A2	A	16	B3	B-
7	A3	A-	17	Caa1	CCC+
8	Baa1	BBB+	18	Caa2	CCC
9	Baa2	BBB	19	Caa3	CCC-
10	Baa3	BBB-	20	Ca	CC
			21	C	C
			22		D

Panel II: Comprehensive credit rating (CCR) scale

Numerical rating	Watch list	Moody's	EJR, Fitch and S&P	Numerical rating	Watch list	Moody's	EJR, Fitch and S&P
Investment grade				Non-Investment grade			
1		Aaa	AAA	20	Positive	Ba1	BB+
2	Negative	Aaa	AAA	21		Ba1	BB+
2	Positive	Aa1	AA+	22	Negative	Ba1	BB+
3		Aa1	AA+	22	Positive	Ba2	BB
4	Negative	Aa1	AA+	23		Ba2	BB
4	Positive	Aa2	AA	24	Negative	Ba2	BB
5		Aa2	AA	24	Positive	Ba3	BB-
6	Negative	Aa2	AA	25		Ba3	BB-
6	Positive	Aa3	AA-	26	Negative	Ba3	BB-
7		Aa3	AA-	26	Positive	B1	B+
8	Negative	Aa3	AA-	27		B1	B+
8	Positive	A1	A+	28	Negative	B1	B+
9		A1	A+	28	Positive	B2	B
10	Negative	A1	A+	29		B2	B
10	Positive	A2	A	30	Negative	B2	B
11		A2	A	30	Positive	B3	B-
12	Negative	A2	A	31		B3	B-
12	Positive	A3	A-	32	Negative	B3	B-
13		A3	A-	32	Positive	Caa1	CCC+
14	Negative	A3	A-	33		Caa1	CCC+
14	Positive	Baa1	BBB+	34	Negative	Caa1	CCC+
15		Baa1	BBB+	34	Positive	Caa2	CCC
16	Negative	Baa1	BBB+	35		Caa2	CCC
16	Positive	Baa2	BBB	36	Negative	Caa2	CCC
17		Baa2	BBB	36	Positive	Caa3	CCC-
18	Negative	Baa2	BBB	37		Caa3	CCC-
18	Positive	Baa3	BBB-	38	Negative	Caa3	CCC-
19		Baa3	BBB-	38	Positive	Ca	CC
20	Negative	Baa3	BBB-	39		Ca	CC
				40	Negative	Ca	CC
				40	Positive	C	C
				41		C	C
				42	Negative	C	C
				43		D	

Numerical conversion of the alpha-numerical ratings designations among our credit rating agencies representatives

Appendix B

This appendix explains in detail how outlooks and watch lists are incorporated into our rating change analysis. Outlooks and watch list inclusions are published by CRAs as a way to signal a possible rating change in pursue of two of their key objectives: first, to be accurate and respond in a timely fashion to relevant information; second, to stabilize their credit assessments, i.e., to change their ratings only as a response to material information that justifies such a course of action. From the overall data set concerning outlooks and watch list inclusions (these may take different values, such as positive, negative, affirms, withdrawn, initial, on watch, not on watch, undetermined), we only use the positive and negative values. The challenge here is to design a way to mix all this information with the traditional rating change data, at the same time retaining the ability to assign more weight to rating changes than to outlooks and watch lists. For simplicity, we consider in this study watch lists to be equivalents to outlooks, even though we know that the time horizon of the two types of rating signals is not the same.[40] Our way to combine the rating with the outlook/watch list information consists in the creation of a new variable called comprehensive credit rating (CCR). We assign a numerical value to this variable in accordance to the rating of the firm and additionally we add or subtract one to that value if the outlook or watch list code attached to a specific rating action is negative or positive. The value assigned for a credit rating goes from 1 to AAA to 43 to D increasing in steps of 2 for each new rating category in the ratings ladder. It is worth to notice that to maintain symmetry we assign the same numerical value to a rating to which a negative outlook is attached as to the next inferior rating to which a positive outlook is attached. For example, a rating of AA- under a negative outlook (a numerical value of 7 as a result of the AA- rating to which we add 1 because of the negative outlook) receives the same CCR scores as a rating of A+ under a positive outlook (a numerical value of 9 as a result of the A+ rating minus 1 because of the positive outlook). Table B1 shows in detail the numerical values assigned to the CCR variable according to the rating and the outlooks and/or watch list data in our possession.

[40] Outlooks are signals of the future direction of rating changes in a time frame of up to 12 months. Watch list are signals of the future direction of rating changes in a time frame of up to 3 months.

Table B1 Numerical values assigned to the comprehensive credit rating variable (CCR) that embodies credit ratings plus outlook or watchlist information

Numerical rating Investment grade	Watch list	Moody's	EJR, Fitch and S&P	Numerical rating Non-Investment grade	Watch list	Moody's	EJR, Fitch and S&P
1		Aaa	AAA	20	Positive	Ba1	BB+
2	Negative	Aaa	AAA	21		Ba1	BB+
2	Positive	Aa1	AA+	22	Negative	Ba1	BB+
3		Aa1	AA+	22	Positive	Ba2	BB
4	Negative	Aa1	AA+	23		Ba2	BB
4	Positive	Aa2	AA	24	Negative	Ba2	BB
5		Aa2	AA	24	Positive	Ba3	BB-
6	Negative	Aa2	AA	25		Ba3	BB-
6	Positive	Aa3	AA-	26	Negative	Ba3	BB-
7		Aa3	AA-	26	Positive	B1	B+
8	Negative	Aa3	AA-	27		B1	B+
8	Positive	A1	A+	28	Negative	B1	B+
9		A1	A+	28	Positive	B2	B
10	Negative	A1	A+	29		B2	B
10	Positive	A2	A	30	Negative	B2	B
11		A2	A	30	Positive	B3	B-
12	Negative	A2	A	31		B3	B-
12	Positive	A3	A-	32	Negative	B3	B-
13		A3	A-	32	Positive	Caa1	CCC+
14	Negative	A3	A-	33		Caa1	CCC+
14	Positive	Baa1	BBB+	34	Negative	Caa1	CCC+
15		Baa1	BBB+	34	Positive	Caa2	CCC
16	Negative	Baa1	BBB+	35		Caa2	CCC
16	Positive	Baa2	BBB	36	Negative	Caa2	CCC
17		Baa2	BBB	36	Positive	Caa3	CCC-
18	Negative	Baa2	BBB	37		Caa3	CCC-
18	Positive	Baa3	BBB-	38	Negative	Caa3	CCC-
19		Baa3	BBB-	38	Positive	Ca	CC
20	Negative	Baa3	BBB-	39		Ca	CC
				40	Negative	Ca	CC
				40	Positive	C	C
				41		C	C
				42	Negative	C	C
				43		D	

Numerical conversion of the alpha-numerical ratings of credit rating agencies. The scale has been modified in such a way that the different ratings schemes are comparable

Appendix C

Table C1 Lead-lag relationship between "issuer-paid" and "investor-paid" credit rating agencies: Post sample

Panel I: Lead-lag relationships between Fitch an Egan-Jones						Marginal effects %			
	Coefficients	z-statistic		p-value	Average |Change	<=-2	-1	1	>=2
Section I: Egan & Jones as rating follower									
Δ-Up by Fitch h=1	0.35	1.09		0.28	6.8	-3.8	-9.8	8.8	4.8
Δ-Up by Fitch h=2	0.05	0.19		0.85	1.0	-0.7	-1.3	1.5	0.6
Δ-Up by Fitch h=3	-0.13	-0.58		0.56	2.5	2.0	3.1	-3.8	-1.2
Δ-Up by Fitch h=4	0.06	0.13		0.90	1.1	-0.8	-1.5	1.6	0.6
Δ-Up by Fitch h=5	0.74	1.85		0.07	13.7	-6.1	-21.2	13.8	13.5
Δ-Up by Fitch h=6	0.28	1.12		0.26	5.5	-3.2	-7.8	7.4	3.7
Δ-Dw by Fitch h=1	-0.81	-4.16	**	0.00	14.6	18.5	10.6	-24.5	-4.7
Δ-Dw by Fitch h=2	-0.83	-6.08	**	0.00	14.9	19.2	10.6	-25.1	-4.7
Δ-Dw by Fitch h=3	-0.65	-2.76	**	0.01	12.0	13.7	10.3	-19.9	-4.1
Δ-Dw by Fitch h=4	-0.85	-5.66	**	0.00	15.1	19.9	10.3	-25.5	-4.7
Δ-Dw by Fitch h=5	0.12	0.48		0.63	2.4	-1.6	-3.2	3.4	1.4
Δ-Dw by Fitch h=6	-0.50	-2.27	*	0.02	9.5	9.7	9.3	-15.4	-3.6
Observations	1,485				Total	7.5	43.6	43.8	5.2
Section II: Fitch as rating follower									
Δ-Up by Egan & Jones h=1	0.57	3.40	**	0.00	11.0	-11.9	-10.2	11.6	10.4
Δ-Up by Egan & Jones h=2	0.77	2.89	**	0.00	15.0	-14.3	-15.7	13.9	16.1
Δ-Up by Egan & Jones h=3	0.01	0.04		0.97	0.3	-0.4	-0.1	0.3	0.2
Δ-Up by Egan & Jones h=4	0.86	3.41		0.00	16.6	-15.1	-18.2	14.3	18.9
Δ-Up by Egan & Jones h=5	0.27	1.06		0.29	5.1	-6.4	-3.8	6.1	4.1
Δ-Up by Egan & Jones h=6	0.50	2.09	*	0.04	9.8	-10.7	-8.8	10.6	9.0
Δ-Dw by Egan & Jones h=1	-0.88	-6.00	**	0.00	14.8	29.5	-3.5	-19.1	-6.9
Δ-Dw by Egan & Jones h=2	-0.42	-2.58	*	0.01	7.0	13.1	1.0	-9.9	-4.1
Δ-Dw by Egan & Jones h=3	-0.32	-1.77		0.08	5.5	9.7	1.3	-7.6	-3.4
Δ-Dw by Egan & Jones h=4	-0.61	-2.80	**	0.01	10.0	20.0	-0.8	-13.9	-5.2
Δ-Dw by Egan & Jones h=5	-0.68	-3.61	**	0.00	11.3	22.6	-1.7	-15.3	-5.6
Δ-Dw by Egan & Jones h=6	-0.44	-1.96		0.05	7.2	13.6	0.7	-10.2	-4.2
Observations	544				Total	18.7	47.5	27.2	6.6

This table reports the results of ordered probit estimation of equations (7) and (8), using data from each pair of agencies. The sample period is 01 June 2002 to 21 December 2007. The dependent variables $\Delta R^{A}_{i,t}$ and $\Delta R^{B}_{i,t}$ represent rating changes by EJR and Fitch, respectively, for firm i at time t. **,* denotes if the coefficients are statistically significant at 1 and 5 percent, respectively.

Table C1 (continued)

Panel II: Lead lag relationships between Moody's an Egan-Jones

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects % <=-2	-1	1	>=2
Section I: Egan & Jones as rating follower									
Δ-Up by Moody's h=1	0.51	2.92	**	0.00	9.7	-5.3	-14.1	11.5	7.9
Δ-Up by Moody's h=2	0.34	2.10	*	0.04	6.6	-4.0	-9.3	8.6	4.7
Δ-Up by Moody's h=3	-0.02	-0.11		0.92	0.5	0.4	0.6	-0.7	-0.2
Δ-Up by Moody's h=4	0.33	1.93		0.05	6.4	-3.9	-8.9	8.3	4.5
Δ-Up by Moody's h=5	0.08	0.35		0.73	1.7	-1.2	-2.1	2.4	0.9
Δ-Up by Moody's h=6	0.20	0.75		0.45	3.9	-2.6	-5.2	5.3	2.4
Δ-Dw by Moody's h=1	-0.84	-6.98	**	0.00	15.3	20.3	10.2	-25.7	-4.8
Δ-Dw by Moody's h=2	-0.65	-5.49	**	0.00	12.2	14.3	10.1	-20.2	-4.2
Δ-Dw by Moody's h=3	-0.61	-3.59	**	0.00	11.5	13.2	9.8	-19.0	-4.0
Δ-Dw by Moody's h=4	-0.36	-2.39	*	0.02	7.0	6.7	7.3	-11.1	-2.9
Δ-Dw by Moody's h=5	-0.31	-1.77		0.08	6.1	5.6	6.5	-9.6	-2.6
Δ-Dw by Moody's h=6	-0.44	-2.23	*	0.03	8.4	8.6	8.3	-13.6	-3.3
Observations	2,194				Total	8.1	41.5	45.2	5.2
Section II: Moody's as rating follower									
Δ-Up by Egan & Jones h=1	0.69	3.99	**	0.00	13.5	-13.5	-13.5	14.4	12.6
Δ-Up by Egan & Jones h=2	0.49	2.25	*	0.03	9.5	-10.5	-8.6	11.1	7.9
Δ-Up by Egan & Jones h=3	0.69	4.07	**	0.00	13.5	-13.4	-13.7	14.3	12.8
Δ-Up by Egan & Jones h=4	0.80	4.44	**	0.00	15.5	-14.7	-16.4	15.5	15.5
Δ-Up by Egan & Jones h=5	0.52	2.75	**	0.01	10.2	-11.1	-9.2	11.8	8.6
Δ-Up by Egan & Jones h=6	0.25	1.66		0.10	4.8	-6.1	-3.6	6.2	3.5
Δ-Dw by Egan & Jones h=1	-0.85	-6.58	**	0.00	14.4	28.8	-3.6	-19.5	-5.7
Δ-Dw by Egan & Jones h=2	-0.66	-5.27	**	0.00	10.9	21.8	-1.1	-15.8	-4.9
Δ-Dw by Egan & Jones h=3	-0.45	-3.07	**	0.00	7.4	14.0	0.8	-11.0	-3.7
Δ-Dw by Egan & Jones h=4	-0.46	-3.10	**	0.00	7.6	14.5	0.8	-11.4	-3.8
Δ-Dw by Egan & Jones h=5	-0.39	-2.63	**	0.01	6.5	12.0	1.1	-9.7	-3.4
Δ-Dw by Egan & Jones h=6	-0.52	-3.51	**	0.00	8.4	16.7	0.1	-12.7	-4.1
Observations	922				Total	18.8	47.0	28.6	5.7

Panel III:Lead lag relationships between Standard and Poor's an Egan-Jones

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects % <=-2	-1	1	>=2
Section I: Egan & Jones as rating follower									
Δ-Up by S&P h=1	0.85	6.09	**	0.00	15.0	-6.6	-23.5	13.7	16.4
Δ-Up by S&P h=2	0.42	2.27	*	0.02	8.1	-4.5	-11.7	10.0	6.2
Δ-Up by S&P h=3	0.67	5.07	**	0.00	12.4	-6.0	-18.8	13.2	11.6
Δ-Up by S&P h=4	0.22	1.31		0.19	4.4	-2.7	-6.1	6.0	2.8
Δ-Up by S&P h=5	0.42	2.13	*	0.03	8.2	-4.5	-11.9	10.2	6.3
Δ-Up by S&P h=6	0.24	1.06		0.29	4.7	-2.9	-6.5	6.3	3.1
Δ-Dw by S&P h=1	-0.88	-7.58	**	0.00	15.7	20.7	10.7	-26.6	-4.9
Δ-Dw by S&P h=2	-0.86	-7.24	**	0.00	15.5	20.4	10.6	-26.2	-4.8
Δ-Dw by S&P h=3	-0.61	-4.12	**	0.00	11.5	12.7	10.3	-19.0	-4.1
Δ-Dw by S&P h=4	-0.47	-2.98	**	0.00	9.1	9.2	9.1	-14.8	-3.5
Δ-Dw by S&P h=5	-0.11	-0.68		0.50	2.2	1.7	2.7	-3.3	-1.1
Δ-Dw by S&P h=6	-0.36	-2.12	*	0.03	7.1	6.6	7.6	-11.2	-2.9
Observations	2,281				Total	7.7	41.9	45.2	5.2
Section II: S&P as rating follower									
Δ-Up by Egan & Jones h=1	0.52	2.55	*	0.01	9.9	-10.5	-9.4	12.4	7.4
Δ-Up by Egan & Jones h=2	0.33	1.51		0.13	6.2	-7.2	-5.1	8.2	4.1
Δ-Up by Egan & Jones h=3	0.54	2.59	*	0.01	10.3	-10.7	-9.9	12.7	7.8
Δ-Up by Egan & Jones h=4	0.57	3.46	**	0.00	11.0	-11.2	-10.8	13.4	8.6
Δ-Up by Egan & Jones h=5	0.49	2.60	**	0.01	9.4	-10.0	-8.8	11.8	7.0
Δ-Up by Egan & Jones h=6	0.64	5.11	**	0.00	12.4	-12.1	-12.8	14.7	10.2
Δ-Dw by Egan & Jones h=1	-0.66	-5.97	**	0.00	10.5	20.9	-1.3	-15.5	-4.1
Δ-Dw by Egan & Jones h=2	-0.50	-4.06	**	0.00	7.8	15.6	0.1	-12.2	-3.4
Δ-Dw by Egan & Jones h=3	-0.52	-4.36	**	0.00	8.0	15.9	0.0	-12.5	-3.5
Δ-Dw by Egan & Jones h=4	-0.46	-3.75	**	0.00	7.3	14.0	0.6	-11.3	-3.3
Δ-Dw by Egan & Jones h=5	-0.22	-1.42		0.16	3.7	6.1	1.2	-5.5	-1.8
Δ-Dw by Egan & Jones h=6	-0.52	-3.50	**	0.00	8.1	16.2	-0.2	-12.6	-3.5
Observations	1,012				Total	17.7	50.8	26.9	4.7

This table reports the results of ordered probit estimation of equations (7) and (8), using data from each pair of agencies. The sample period is 01 June 2002 to 21 December 2007. The dependent variables $\Delta R^{A}_{i,t}$ and $\Delta R^{B}_{i,t}$ represent rating changes by EJR andMoody's or EJR and S&P, respectively, for firm i at time t. **,* denotes if the coefficients are statistically significant at 1 and 5 percent, respectively.

Appendix D

Table D1 Lead-lag relationship between "issuer-paid" and "investor-paid" credit rating agencies including outlooks and watch list information: Pre sample

Panel I: Lead-lag relationships between Egan-Jones and Fitch including outlooks and watch list information

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects %							
						<=-4	-3	-2	-1	1	2	3	>=4
Section I: Egan & Jones as rating follower													
Δ-Up by Fitch h=1	-0.52	-0.70		0.48	5.0	9.0	5.1	6.0	-4.0	-6.8	-5.6	-1.6	-2.2
Δ-Up by Fitch h=2	0.20	0.26		0.79	1.9	-2.0	-1.7	-3.0	-0.8	2.2	2.7	0.9	1.6
Δ-Up by Fitch h=3	0.69	20.35	**	0.00	6.7	-4.6	-4.6	-9.8	-7.6	5.1	9.2	3.9	8.6
Δ-Up by Fitch h=4	-0.07	-2.38	*	0.02	0.6	0.9	0.7	1.0	0.0	-0.9	-0.9	-0.3	-0.5
Δ-Up by Fitch h=5	-0.07	-2.38	*	0.02	0.6	0.9	0.7	1.0	0.0	-0.9	-0.9	-0.3	-0.5
Δ-Up by Fitch h=6	-0.87	-24.73	**	0.00	8.4	18.4	8.1	7.1	-10.3	-10.8	-7.8	-2.0	-2.7
Δ-Dw by Fitch h=1	-0.64	-3.63	**	0.00	6.2	11.6	6.3	6.9	-5.5	-8.3	-6.6	-1.8	-2.5
Δ-Dw by Fitch h=2	-0.08	-0.36		0.72	0.7	1.0	0.7	1.1	0.0	-1.0	-1.0	-0.3	-0.5
Δ-Dw by Fitch h=3	-0.40	-1.97	*	0.05	3.8	6.2	3.9	5.0	-2.2	-5.2	-4.5	-1.3	-1.9
Δ-Dw by Fitch h=4	-0.61	-3.24	**	0.00	5.8	10.8	5.9	6.6	-5.0	-7.8	-6.3	-1.7	-2.4
Δ-Dw by Fitch h=5	0.44	1.59		0.11	4.2	-3.6	-3.3	-6.5	-3.4	4.2	6.0	2.3	4.4
Δ-Dw by Fitch h=6	-0.17	-0.53		0.59	1.5	2.2	1.6	2.3	-0.3	-2.1	-2.0	-0.6	-1.0
Observations	1,455				Total	5.8	6.7	18.2	36.6	17.1	10.1	2.4	3.0
Section II: Fitch as rating follower													
Δ-Up by Egan & Jones h=1	-0.03	-0.07		0.95	0.3	1.1	0.1	-0.4	-0.2	-0.1	-0.3	0.0	-0.2
Δ-Up by Egan & Jones h=2	0.19	0.29		0.77	1.7	-6.1	-0.6	1.4	1.5	0.5	2.0	0.1	1.2
Δ-Up by Egan & Jones h=3	1.15	2.35	*	0.02	10.2	-24.7	-4.3	-11.6	6.1	2.4	14.3	1.4	16.5
Δ-Up by Egan & Jones h=4	1.30	2.57	*	0.01	11.6	-25.6	-4.7	-16.2	5.8	2.5	15.6	1.6	21.0
Δ-Up by Egan & Jones h=5	-0.01	-0.02		0.99	0.1	0.3	0.0	-0.1	-0.1	0.0	-0.1	0.0	0.0
Δ-Up by Egan & Jones h=6	6.39	20.60	**	0.00	24.5	-30.1	-6.2	-47.8	-6.6	-1.6	-5.5	-0.3	98.1
Δ-Dw by Egan & Jones h=1	-0.85	-4.33	**	0.00	8.1	31.2	1.4	-15.3	-5.8	-1.6	-6.4	-0.4	-3.1
Δ-Dw by Egan & Jones h=2	-0.47	-2.36	*	0.02	4.5	16.8	1.0	-7.5	-3.4	-1.0	-3.9	-0.3	-1.8
Δ-Dw by Egan & Jones h=3	-0.53	-2.37	*	0.02	5.1	19.3	1.0	-9.2	-3.7	-1.1	-4.1	-0.3	-1.9
Δ-Dw by Egan & Jones h=4	-0.46	-2.18	*	0.03	4.5	16.9	0.9	-7.9	-3.3	-0.9	-3.7	-0.2	-1.7
Δ-Dw by Egan & Jones h=5	-0.56	-2.48	*	0.01	5.4	20.6	1.0	-9.9	-3.9	-1.1	-4.4	-0.3	-2.0
Δ-Dw by Egan & Jones h=6	-0.06	-0.23		0.82	0.6	2.1	0.2	-0.8	-0.5	-0.1	-0.6	0.0	-0.3
Observations	224				Total	28.1	6.1	48.5	7.0	1.8	6.0	0.3	2.2

Panel II: Lead-lag relationships between Egan-Jones and Moody's including outlooks and watch list information

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects %							
						<=-4	-3	-2	-1	1	2	3	>=4
Section I: Egan & Jones as rating follower													
Δ-Up by Moody's h=1	0.58	1.37		0.17	5.7	-6.0	-3.7	-8.3	-4.6	4.8	6.9	3.4	7.5
Δ-Up by Moody's h=2	-0.10	-0.25		0.81	0.9	1.7	0.8	1.3	-0.2	-1.2	-1.1	-0.4	-0.7
Δ-Up by Moody's h=3	0.81	2.44	*	0.02	7.8	-7.1	-4.7	-11.1	-8.5	5.1	9.1	4.9	12.3
Δ-Up by Moody's h=4	-0.09	-0.23		0.82	0.8	1.4	0.7	1.1	-0.2	-1.1	-1.0	-0.4	-0.6
Δ-Up by Moody's h=5	-0.06	-0.23		0.82	0.6	1.0	0.5	0.8	-0.1	-0.7	-0.7	-0.3	-0.5
Δ-Up by Moody's h=6	0.70	2.21	*	0.03	6.8	-6.7	-4.3	-9.8	-6.5	5.1	8.1	4.2	9.9
Δ-Dw by Moody's h=1	-0.56	-5.24	**	0.00	5.5	12.0	4.4	5.5	-4.6	-7.1	-5.4	-1.9	-2.8
Δ-Dw by Moody's h=2	-0.47	-3.04	**	0.00	4.5	9.7	3.7	4.8	-3.4	-5.9	-4.6	-1.7	-2.5
Δ-Dw by Moody's h=3	-0.15	-0.97		0.33	1.4	2.5	1.2	1.8	-0.4	-1.8	-1.6	-0.6	-1.0
Δ-Dw by Moody's h=4	-0.38	-2.91	**	0.00	3.7	7.5	3.0	4.2	-2.4	-4.8	-3.9	-1.4	-2.1
Δ-Dw by Moody's h=5	0.00	-0.01		0.99	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0	0.0
Δ-Dw by Moody's h=6	-0.01	-0.09		0.93	0.1	0.2	0.1	0.2	0.0	-0.2	-0.1	-0.1	-0.1
Observations	1,965				Total	8.5	6.4	18.8	34.2	16.6	9.2	2.8	3.5
Section II: Moody's as rating follower													
Δ-Up by Egan & Jones h=1	0.72	3.41	**	0.00	6.8	-14.0	-3.7	-9.6	1.0	6.9	9.4	1.4	8.6
Δ-Up by Egan & Jones h=2	0.04	0.15		0.89	0.4	-1.2	-0.2	-0.3	0.4	0.5	0.5	0.1	0.3
Δ-Up by Egan & Jones h=3	1.13	3.49	**	0.00	10.6	-17.4	-5.2	-16.2	-3.4	7.3	14.0	2.4	18.5
Δ-Up by Egan & Jones h=4	0.63	1.80		0.07	6.0	-12.6	-3.3	-8.1	1.3	6.2	8.2	1.2	7.1
Δ-Up by Egan & Jones h=5	1.08	2.67	**	0.01	10.1	-17.0	-5.1	-15.4	-2.9	7.4	13.5	2.3	17.1
Δ-Up by Egan & Jones h=6	0.76	2.07	*	0.04	7.1	-14.3	-3.9	-10.3	0.4	7.0	10.0	1.5	9.6
Δ-Dw by Egan & Jones h=1	-0.70	-6.25	**	0.00	6.5	22.8	2.6	0.7	-8.5	-7.8	-6.2	-0.7	-2.9
Δ-Dw by Egan & Jones h=2	-0.54	-4.75	**	0.00	5.2	17.3	2.2	1.2	-6.4	-6.2	-5.0	-0.6	-2.4
Δ-Dw by Egan & Jones h=3	-0.63	-4.85	**	0.00	5.9	20.2	2.4	0.8	-7.6	-7.0	-5.6	-0.6	-2.7
Δ-Dw by Egan & Jones h=4	-0.37	-2.92	**	0.00	3.6	11.4	1.6	1.4	-4.2	-4.3	-3.6	-0.4	-1.8
Δ-Dw by Egan & Jones h=5	-0.21	-1.49		0.14	2.0	6.0	1.0	1.1	-2.1	-2.4	-2.1	-0.3	-1.1
Δ-Dw by Egan & Jones h=6	-0.53	-4.16	**	0.00	5.0	16.8	2.1	1.0	-6.3	-6.0	-4.8	-0.5	-2.3
Observations	736				Total	19.2	6.7	27.1	23.2	12.6	7.6	0.7	2.8

This table reports the results of ordered probit estimation of equations (7) and (8), using data from each pair of agencies. The sample period is 17 July 1997 to 31 May 2002. The dependent variables $\Delta R^A_{i,t}$ and $\Delta R^B_{i,t}$ represent comprehensive rating changes (rating changes plus watchlist or outlooks) by EJR and Fitch or EJR and Moody's, respectively, for firm i at time t. **,* denotes if the coefficients are statistically significant at 1 and 5 percent, respectively.

Table D1 (continued)

Panel III: Lead-lag relationships between Egan-Jones and Standard & Poor's including outlooks and watch list information

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects % <=-4	-3	-2	-1	1	2	3	>=4
Section I: Egan & Jones as rating follower													
Δ-Up by S&P h=1	0.85	2.35	*	0.02	8.2	-7.2	-4.8	-11.6	-9.3	4.8	9.1	5.2	13.8
Δ-Up by S&P h=2	-0.19	-0.40		0.69	1.8	3.4	1.5	2.4	-0.7	-2.4	-2.1	-0.8	-1.3
Δ-Up by S&P h=3	0.55	0.99		0.32	5.3	-5.8	-3.6	-7.8	-4.1	4.6	6.4	3.2	7.1
Δ-Up by S&P h=4	0.22	0.28		0.78	2.1	-3.0	-1.6	-3.2	-0.6	2.4	2.6	1.2	2.2
Δ-Up by S&P h=5	0.11	0.26		0.79	1.0	-1.6	-0.8	-1.5	-0.1	1.2	1.3	0.5	1.0
Δ-Up by S&P h=6	0.57	1.46		0.15	5.5	-5.9	-3.7	-8.0	-4.4	4.7	6.5	3.3	7.4
Δ-Dw by S&P h=1	-0.63	-5.73	**	0.00	6.1	13.9	4.8	5.8	-5.6	-7.8	-5.8	-2.1	-3.1
Δ-Dw by S&P h=2	-0.32	-2.29	*	0.02	3.1	6.1	2.6	3.7	-1.7	-4.1	-3.3	-1.3	-2.0
Δ-Dw by S&P h=3	-0.25	-1.65		0.10	2.3	4.5	2.0	3.0	-1.1	-3.1	-2.6	-1.0	-1.6
Δ-Dw by S&P h=4	-0.40	-3.18	**	0.00	3.9	8.0	3.2	4.4	-2.6	-5.1	-4.0	-1.5	-2.3
Δ-Dw by S&P h=5	-0.20	-1.39		0.16	1.9	3.4	1.6	2.4	-0.7	-2.4	-2.1	-0.8	-1.3
Δ-Dw by S&P h=6	-0.35	-2.51	*	0.01	3.3	6.6	2.7	3.9	-2.0	-4.3	-3.5	-1.4	-2.1
Nº Obs	1,997				Total	8.5	6.5	18.9	34.0	16.5	9.1	2.9	3.7
Section II: S&P as rating follower													
Δ-Up by Egan & Jones h=1	0.33	1.10		0.27	3.1	-9.7	-1.1	-1.7	2.5	1.9	3.9	0.4	3.8
Δ-Up by Egan & Jones h=2	0.52	1.35		0.18	5.0	-14.1	-1.8	-4.2	3.3	2.8	6.3	0.6	7.1
Δ-Up by Egan & Jones h=3	0.72	2.52	*	0.01	6.9	-17.9	-2.6	-7.4	3.6	3.5	8.7	0.9	11.1
Δ-Up by Egan & Jones h=4	0.23	0.64		0.52	2.2	-7.1	-0.8	-0.9	1.9	1.3	2.7	0.2	2.5
Δ-Up by Egan & Jones h=5	1.61	4.12	**	0.00	14.4	-26.0	-4.8	-24.8	-2.0	2.6	12.9	1.8	40.4
Δ-Up by Egan & Jones h=6	0.85	2.51	*	0.01	8.2	-19.8	-3.0	-10.1	3.4	3.8	10.1	1.1	14.6
Δ-Dw by Egan & Jones h=1	-0.71	-5.29	**	0.00	6.8	26.1	1.2	-5.3	-7.0	-3.8	-6.3	-0.5	-4.3
Δ-Dw by Egan & Jones h=2	-0.53	-3.77	**	0.00	5.0	18.9	1.1	-3.2	-5.2	-2.9	-4.9	-0.4	-3.4
Δ-Dw by Egan & Jones h=3	-0.38	-2.58	*	0.01	3.5	13.3	0.9	-1.7	-3.7	-2.1	-3.7	-0.3	-2.6
Δ-Dw by Egan & Jones h=4	-0.38	-2.77	**	0.01	3.6	13.6	0.9	-1.8	-3.8	-2.2	-3.8	-0.3	-2.7
Δ-Dw by Egan & Jones h=5	-0.66	-3.92	**	0.00	6.3	24.2	1.0	-5.3	-6.5	-3.5	-5.7	-0.4	-3.7
Δ-Dw by Egan & Jones h=6	-0.39	-2.53	*	0.01	3.6	13.7	0.9	-1.9	-3.8	-2.2	-3.7	-0.3	-2.6
Observations	583				Total	26.5	5.4	35.8	13.7	5.8	8.0	0.6	4.2

This table reports the results of ordered probit estimation of equations (7) and (8), using data from each pair of agencies. The sample period is 17 July 1997 to 31 May 2002. The dependent variables $\Delta R^{A}_{i,t}$ and $\Delta R^{B}_{i,t}$ represent comprehensive rating changes (rating changes plus watchlist or outlooks) by EJR and S&P, respectively, for firm i at time t. **,* denotes if the coefficients are statistically significant at 1 and 5 percent, respectively.

Table D2 Lead-lag relationship between "Issuer-paid" and "Investor-paid" credit rating agencies including outlooks and watch list information: Post sample

Panel I: Lead-lag relationships between Egan-Jones and Fitch including outlooks and watch list information

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects % <=-4	-3	-2	-1	1	2	3	>=4
Section I: Egan & Jones as rating follower													
Δ-Up by Fitch h=1	0.18	0.76		0.45	1.8	-1.5	-1.1	-2.8	-1.7	1.1	2.9	1.3	1.7
Δ-Up by Fitch h=2	0.27	1.35		0.18	2.6	-2.1	-1.6	-4.2	-2.7	1.3	4.4	2.0	2.8
Δ-Up by Fitch h=3	-0.21	-0.97		0.33	2.0	2.3	1.5	3.2	1.1	-2.1	-3.4	-1.2	-1.4
Δ-Up by Fitch h=4	0.20	1.09		0.28	2.0	-1.7	-1.2	-3.2	-1.9	1.2	3.4	1.5	2.0
Δ-Up by Fitch h=5	0.35	1.72		0.09	3.4	-2.5	-2.0	-5.4	-3.7	1.4	5.6	2.7	3.9
Δ-Up by Fitch h=6	0.21	1.33		0.18	2.1	-1.7	-1.3	-3.4	-2.1	1.2	3.5	1.6	2.1
Δ-Dw by Fitch h=1	-0.61	-4.33	**	0.00	5.8	9.3	4.7	8.4	0.7	-8.1	-9.2	-3.0	-3.0
Δ-Dw by Fitch h=2	-0.82	-6.25	**	0.00	7.7	14.5	6.4	10.0	-0.8	-11.5	-11.5	-3.5	-3.5
Δ-Dw by Fitch h=3	-0.56	-3.53	**	0.00	5.4	8.4	4.4	7.9	0.9	-7.3	-8.6	-2.8	-2.9
Δ-Dw by Fitch h=4	-0.79	-5.80	**	0.00	7.4	13.5	6.1	9.8	-0.6	-10.9	-11.1	-3.4	-3.4
Δ-Dw by Fitch h=5	-0.01	-0.05		0.96	0.1	0.1	0.1	0.2	0.1	-0.1	-0.2	-0.1	-0.1
Δ-Dw by Fitch h=6	-0.35	-2.11	*	0.04	3.4	4.4	2.6	5.3	1.4	-4.0	-5.5	-1.9	-2.1
Observations	3,056				Total	4.6	4.4	16.6	23.1	26.6	16.7	4.3	3.8
Section II: Fitch as rating follower													
Δ-Up by Egan & Jones h=1	0.47	2.96	**	0.00	4.6	-8.6	-2.4	-6.7	-0.8	1.3	10.0	0.5	6.7
Δ-Up by Egan & Jones h=2	0.75	3.60	**	0.00	7.3	-11.5	-3.6	-11.4	-2.7	0.9	14.4	0.9	12.9
Δ-Up by Egan & Jones h=3	0.33	1.14		0.25	3.3	-6.4	-1.7	-4.6	-0.3	1.1	7.2	0.4	4.4
Δ-Up by Egan & Jones h=4	0.35	2.15	*	0.03	3.4	-6.7	-1.8	-4.8	-0.3	1.2	7.5	0.4	4.5
Δ-Up by Egan & Jones h=5	0.44	2.41	*	0.02	4.4	-8.1	-2.3	-6.3	-0.7	1.3	9.4	0.5	6.2
Δ-Up by Egan & Jones h=6	0.60	3.37	**	0.00	5.9	-10.1	-3.0	-8.9	-1.6	1.2	12.3	0.7	9.4
Δ-Dw by Egan & Jones h=1	-0.73	-5.64	**	0.00	7.0	21.7	3.1	3.2	-3.9	-4.8	-14.2	-0.5	-4.6
Δ-Dw by Egan & Jones h=2	-0.54	-4.26	**	0.00	5.2	15.1	2.5	3.3	-2.5	-3.4	-10.9	-0.4	-3.8
Δ-Dw by Egan & Jones h=3	-0.31	-2.10	*	0.04	3.1	8.2	1.6	2.6	-1.1	-1.9	-6.6	-0.3	-2.5
Δ-Dw by Egan & Jones h=4	-0.38	-2.48	*	0.01	3.8	10.4	1.9	2.9	-1.5	-2.4	-8.0	-0.3	-3.0
Δ-Dw by Egan & Jones h=5	-0.42	-3.07	**	0.00	4.2	11.5	2.1	3.1	-1.7	-2.6	-8.8	-0.3	-3.2
Δ-Dw by Egan & Jones h=6	-0.47	-2.91	**	0.00	4.6	13.0	2.2	3.1	-2.1	-3.0	-9.6	-0.3	-3.4
Observations	843				Total	14.7	5.6	25.5	15.9	11.1	21.6	0.6	4.9

This table reports the results of ordered probit estimations of Eqs. (7) and (8), using data from each pair of agencies. The sample period is 01 June 2002 to 21 December 2007. The dependent variables $\Delta R^{A}_{i,t}$ and $\Delta R^{B}_{i,t}$ represent comprehensive rating changes (rating changes plus watchlist or outlooks) by EJR and Fitch, respectively, for firm i at time t. **,* denotes if the coefficients are statistically significant at 1 and 5 percent, respectively.

Table D2 (continued)

Panel II: Lead-lag relationships between Egan-Jones and Moody's including outlooks and watch list information

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects %							
						<=-4	-3	-2	-1	1	2	3	>=4
Section I: Egan & Jones as rating follower													
Δ-Up by Moody's h=1	0.43	3.00	**	0.00	4.1	-3.1	-2.4	-6.3	-4.6	1.1	6.7	3.5	5.2
Δ-Up by Moody's h=2	0.43	3.24	**	0.00	4.1	-3.1	-2.4	-6.4	-4.7	1.1	6.7	3.5	5.2
Δ-Up by Moody's h=3	0.07	0.41		0.68	0.7	-0.7	-0.5	-1.0	-0.6	0.5	1.1	0.5	0.6
Δ-Up by Moody's h=4	0.19	1.29		0.20	1.8	-1.6	-1.2	-2.9	-1.7	1.1	3.1	1.4	1.9
Δ-Up by Moody's h=5	0.19	1.12		0.26	1.9	-1.7	-1.2	-3.0	-1.8	1.1	3.2	1.5	1.9
Δ-Up by Moody's h=6	0.19	1.03		0.31	1.9	-1.7	-1.2	-2.9	-1.8	1.1	3.1	1.5	1.9
Δ-Dw by Moody's h=1	-0.76	-7.71	**	0.00	7.1	13.3	6.0	9.1	-0.2	-10.3	-10.9	-3.6	-3.4
Δ-Dw by Moody's h=2	-0.71	-6.74	**	0.00	6.6	12.1	5.6	8.7	0.1	-9.5	-10.3	-3.4	-3.3
Δ-Dw by Moody's h=3	-0.50	-3.97	**	0.00	4.8	7.3	3.9	6.8	1.2	-6.1	-7.7	-2.7	-2.7
Δ-Dw by Moody's h=4	-0.40	-3.31	**	0.00	3.9	5.5	3.1	5.7	1.4	-4.7	-6.3	-2.3	-2.4
Δ-Dw by Moody's h=5	-0.39	-3.40	**	0.00	3.8	5.3	3.0	5.6	1.4	-4.5	-6.2	-2.3	-2.4
Δ-Dw by Moody's h=6	-0.22	-1.69		0.09	2.2	2.7	1.7	3.3	1.2	-2.3	-3.6	-1.4	-1.6
Observations	4,268				Total	4.9	4.7	16.2	22.2	26.7	16.8	4.5	3.9
Section II: Moody's as rating follower													
Δ-Up by Egan & Jones h=1	0.81	5.43	**	0.00	7.8	-9.4	-5.9	-9.7	-6.3	3.6	10.7	5.1	12.0
Δ-Up by Egan & Jones h=2	0.41	2.68	**	0.01	3.9	-6.0	-3.4	-4.8	-1.6	3.1	5.9	2.3	4.5
Δ-Up by Egan & Jones h=3	0.46	3.03	**	0.00	4.5	-6.6	-3.8	-5.5	-2.2	3.3	6.7	2.7	5.4
Δ-Up by Egan & Jones h=4	0.70	5.01	**	0.00	6.8	-8.7	-5.4	-8.5	-4.9	3.7	9.7	4.3	9.7
Δ-Up by Egan & Jones h=5	0.42	2.59	*	0.01	4.1	-6.3	-3.5	-5.0	-1.8	3.2	6.2	2.4	4.7
Δ-Up by Egan & Jones h=6	0.23	1.54		0.12	2.3	-3.9	-2.0	-2.7	-0.5	2.1	3.4	1.3	2.3
Δ-Dw by Egan & Jones h=1	-0.66	-6.83	**	0.00	6.5	16.8	5.2	3.9	-4.6	-7.5	-8.1	-2.4	-3.3
Δ-Dw by Egan & Jones h=2	-0.61	-5.63	**	0.00	6.0	15.3	4.9	3.7	-4.2	-7.0	-7.5	-2.2	-3.1
Δ-Dw by Egan & Jones h=3	-0.21	-1.75		0.08	2.1	4.5	1.9	1.9	-0.7	-2.3	-3.0	-0.9	-1.4
Δ-Dw by Egan & Jones h=4	-0.51	-4.20	**	0.00	5.1	12.5	4.2	3.5	-3.2	-5.9	-6.5	-1.9	-2.7
Δ-Dw by Egan & Jones h=5	-0.64	-5.41	**	0.00	6.3	16.4	5.0	3.7	-4.7	-7.3	-7.8	-2.2	-3.1
Δ-Dw by Egan & Jones h=6	-0.37	-3.32	**	0.00	3.6	8.4	3.1	2.9	-1.8	-4.1	-4.9	-1.5	-2.2
Observations	1,371				Total	11.1	8.5	18.0	26.9	16.5	12.4	3.0	3.6

Panel III: Lead-lag relationships between Egan-Jones and Standard & Poor's including outlooks and watch list information

	Coefficients	z-statistic		p-value	Average |Change	Marginal effects %							
						<=-4	-3	-2	-1	1	2	3	>=4
Section I: Egan & Jones as rating follower													
Δ-Up by S&P h=1	0.54	4.34	**	0.00	5.1	-3.5	-2.9	-7.8	-6.2	0.4	8.0	4.5	7.5
Δ-Up by S&P h=2	0.40	2.97	**	0.00	3.8	-2.9	-2.3	-5.9	-4.2	1.1	6.2	3.2	4.9
Δ-Up by S&P h=3	0.47	3.49	**	0.00	4.5	-3.2	-2.6	-6.9	-5.2	0.8	7.1	3.8	6.1
Δ-Up by S&P h=4	0.18	0.93		0.35	1.8	-1.6	-1.1	-2.7	-1.6	1.0	2.9	1.3	1.9
Δ-Up by S&P h=5	0.17	1.14		0.25	1.7	-1.5	-1.1	-2.6	-1.5	0.9	2.7	1.3	1.7
Δ-Up by S&P h=6	-0.06	-0.35		0.73	0.6	0.6	0.4	0.9	0.4	-0.5	-1.0	-0.4	-0.5
Δ-Dw by S&P h=1	-0.77	-8.23	**	0.00	7.2	13.5	6.1	9.2	-0.2	-10.3	-11.0	-3.6	-3.7
Δ-Dw by S&P h=2	-0.66	-6.05	**	0.00	6.2	10.7	5.2	8.4	0.5	-8.5	-9.7	-3.3	-3.4
Δ-Dw by S&P h=3	-0.43	-3.87	**	0.00	4.2	5.9	3.3	6.0	1.4	-4.9	-6.7	-2.4	-2.7
Δ-Dw by S&P h=4	-0.46	-3.38	**	0.00	4.5	6.6	3.6	6.5	1.3	-5.5	-7.2	-2.6	-2.8
Δ-Dw by S&P h=5	-0.23	-1.83		0.07	2.2	2.7	1.7	3.4	1.2	-2.2	-3.7	-1.4	-1.7
Δ-Dw by S&P h=6	-0.14	-1.12		0.26	1.4	1.6	1.0	2.1	0.9	-1.3	-2.3	-0.9	-1.1
Observations	4,351				Total	4.8	4.7	16.2	22.1	26.5	17.0	4.6	4.2
Section II: S&P as rating follower													
Δ-Up by Egan & Jones h=1	0.51	3.12	**	0.00	4.9	-7.7	-4.4	-6.7	-1.0	2.5	8.8	2.4	6.1
Δ-Up by Egan & Jones h=2	0.41	2.69	**	0.01	3.9	-6.5	-3.5	-5.2	-0.4	2.2	7.0	1.8	4.5
Δ-Up by Egan & Jones h=3	0.48	2.94	**	0.00	4.7	-7.4	-4.1	-6.4	-0.8	2.5	8.3	2.2	5.7
Δ-Up by Egan & Jones h=4	0.77	5.15	**	0.00	7.5	-10.0	-6.1	-10.7	-3.3	2.6	12.4	3.8	11.3
Δ-Up by Egan & Jones h=5	0.30	1.82		0.07	2.8	-5.1	-2.6	-3.7	0.0	1.8	5.2	1.3	3.0
Δ-Up by Egan & Jones h=6	0.43	2.47	*	0.01	4.1	-6.8	-3.7	-5.5	-0.5	2.3	7.4	2.0	4.8
Δ-Dw by Egan & Jones h=1	-0.58	-6.18	**	0.00	5.7	15.1	4.7	3.0	-4.3	-4.9	-8.9	-1.7	-3.1
Δ-Dw by Egan & Jones h=2	-0.43	-3.63	**	0.00	4.2	10.6	3.6	2.7	-2.8	-3.6	-6.8	-1.3	-2.4
Δ-Dw by Egan & Jones h=3	-0.32	-2.97	**	0.00	3.2	7.5	2.8	2.4	-1.8	-2.6	-5.2	-1.0	-2.0
Δ-Dw by Egan & Jones h=4	-0.50	-4.79	**	0.00	4.9	12.7	4.1	2.9	-3.5	-4.2	-7.8	-1.5	-2.7
Δ-Dw by Egan & Jones h=5	-0.33	-2.96	**	0.00	3.2	7.7	2.8	2.4	-1.9	-2.7	-5.3	-1.1	-2.0
Δ-Dw by Egan & Jones h=6	-0.40	-3.34	**	0.00	4.0	9.8	3.4	2.6	-2.6	-3.3	-6.3	-1.2	-2.3
Observations	1,386				Total	12.2	9.3	23.0	21.9	12.3	15.3	2.3	3.6

This table reports the results of ordered probit estimations of Eqs. (7) and (8), using data from each pair of agencies. The sample period is 01 June 2002 to 21 December 2007. The dependent variables $\Delta R^A_{i,t}$ and $\Delta R^B_{i,t}$ represent comprehensive rating changes (rating changes plus watchlist or outlooks) by EJR and Moody's or EJR and S&P, respectively, for firm i at time t. **,* denotes if the coefficients are statistically significant at 1 and 5 percent, respectively.

Appendix E

Table E1 Comparison between abnormal stock returns of conditional and unconditional rating announcements including outlooks and watch list data
Panel one comprises a sample period from 17 July 1997 to 31 May 2002, while panel 2 includes observations from 01 Jun 2002 to 21 December 2007

Panel I	Type	Credit rating agency pair	Abnormal Returns			Variance of abnormal returns	N	mean difference T-test	p-value
Downgrades	Conditional	EJR - Moody's	-7.43%	***		0.017	82	-2.18	0.04
	Unconditional		-4.15%	***	(ΔΔΔ)	0.011	821		
	Conditional	Moody's - EJR	-4.18%	**		0.033	91	-1.28	0.18
	Unconditional		-1.66%	***	(ΔΔΔ)	0.007	275		
	Conditional	EJR - S&P	-7.67%	***		0.022	75	-1.91	0.07
	Unconditional		-4.34%	***	(ΔΔ)	0.012	842		
	Conditional	S&P - EJR	-5.06%	**		0.031	71	-1.05	0.23
	Unconditional		-2.77%	***	(ΔΔ)	0.010	271		
	Conditional	EJR - Fitch	-10.89%	***		0.030	36	-2.35	0.03
	Unconditional		-4.01%	***	(ΔΔ)	0.011	602		
	Conditional	Fitch - EJR	-4.94%			0.049	49	-0.83	0.28
	Unconditional		-2.25%	***	(ΔΔ)	0.006	112		
Upgrades	Conditional	EJR - Moody's	0.73%			0.001	7	-1.92	0.07
	Unconditional		3.45%	***	(ΔΔΔ)	0.006	382		
	Conditional	Moody's - EJR	0.41%			0.000	6	-0.32	0.37
	Unconditional		0.65%		(ΔΔΔ)	0.002	60		
	Conditional	EJR - S&P	2.18%			0.002	2	-0.41	0.27
	Unconditional		3.37%	***	(ΔΔΔ)	0.007	397		
	Conditional	S&P - EJR	4.89%	**		0.001	3	2.41	0.05
	Unconditional		0.42%		(ΔΔΔ)	0.001	57		
	Conditional	EJR - Fitch	2.80%			0.000	1	-1.20	-
	Unconditional		3.40%	***	(ΔΔΔ)	0.007	275		
	Conditional	Fitch - EJR	6.06%			0.002	2	2.03	-
	Unconditional		-0.63%		(ΔΔΔ)	0.001	16		

Panel II	Type	Credit rating agency pair	Abnormal Returns			Variance of abnormal returns	N	mean difference T-test	p-value
Downgrades	Conditional	EJR - Moody's	-7.24%	***	(ΔΔΔ)	0.019	76	-2.32	0.03
	Unconditional		-3.46%	***	(ΔΔΔ)	0.010	1,220		
	Conditional	Moody's - EJR	-1.78%	*	(ΔΔΔ)	0.008	97	-0.48	0.35
	Unconditional		-1.31%	***	(ΔΔΔ)	0.006	452		
	Conditional	EJR - S&P	-5.39%	***	(Δ)	0.018	85	-1.24	0.18
	Unconditional		-3.58%	***	(ΔΔΔ)	0.011	1,230		
	Conditional	S&P - EJR	-1.64%		(Δ)	0.015	103	0.44	0.36
	Unconditional		-2.21%	***	(ΔΔΔ)	0.009	527		
	Conditional	EJR - Fitch	-8.71%	***		0.028	42	-2.10	0.05
	Unconditional		-3.19%	***	(ΔΔΔ)	0.009	891		
	Conditional	Fitch - EJR	-3.74%	*		0.025	66	-1.31	0.17
	Unconditional		-1.06%	*	(ΔΔΔ)	0.008	244		
Upgrades	Conditional	EJR - Moody's	1.17%	**	(ΔΔ)	0.001	22	-2.03	0.05
	Unconditional		2.34%	***	(ΔΔΔ)	0.004	1,344		
	Conditional	Moody's - EJR	-0.45%	*	(ΔΔ)	0.000	28	-2.15	0.04
	Unconditional		0.15%		(ΔΔΔ)	0.000	284		
	Conditional	EJR - S&P	0.72%			0.003	25	-1.59	0.11
	Unconditional		2.36%	***	(ΔΔΔ)	0.004	1,374		
	Conditional	S&P - EJR	0.16%			0.000	23	0.08	0.39
	Unconditional		0.13%		(ΔΔΔ)	0.000	282		
	Conditional	EJR - Fitch	-0.93%			0.002	9	-1.83	0.08
	Unconditional		1.96%	***	(ΔΔΔ)	0.003	945		
	Conditional	Fitch - EJR	0.70%	*		0.000	28	1.46	0.14
	Unconditional		0.07%		(ΔΔΔ)	0.000	160		

This table reports the abnormal stock returns of the market to credit rating announcements. The event window is set at the announcement day. The benchmark is the CRSP equally weighted index. ***, **, * denote p-values from a two-sided test of zero mean abnormal returns being statistically significant different from zero at 1, 5 and 10 percent, respectively. ΔΔΔ, Δ, Δ denote p-values from a mean difference t-test for abnormal returns, between agencies in a pair, being statistically significant different from zero at 1, 5 and 10 percent, respectively. Mean difference t-test evaluates whether the difference in mean abnormal returns between conditional and uncoditional event studies for the same pair of credit rating agencies is statistically significant.