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Private information in currency markets☆

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ABSTRACT

Using daily abnormal currency returns for the universe of countries with flexible exchange rates, we show local currency depreciations ahead of unscheduled, public sovereign debt downgrade announcements. Consistent with the private information hypothesis, the effect is stronger in lower institutional quality countries and holds after we control for concurrent public information and for publicly available rumors about the forthcoming downgrades. Our results persist when abnormal currency returns are adjusted for global carry and dollar risk factors, world equity and bond returns, as well as local stock market returns. Finally, the currency depreciations are permanent, providing evidence for a link between fundamentals and currency markets.

1. Introduction

In classic textbook treatments, asset prices respond to public news about fundamentals: as public news become common knowledge, asset prices move (for instance, Glosten and Milgrom, 1985). Traditional models of foreign exchange rate determination also posit an information environment where all relevant information is publicly known and immediately incorporated into prices. Empirical evidence supports the idea that high-frequency foreign exchange rate movements are linked to fundamentals that become common knowledge at pre-specified announcement dates (Andersen et al., 2003, 2007). Nevertheless, the link between foreign exchange rate changes and fundamental shocks remains controversial since the seminal work by Meese and Rogoff (1983).

We contribute to this debate by uncovering evidence for a link between sovereign risk and daily nominal foreign exchange (FX) rates through the behavior of bilateral FX rates against the US dollar before, during, and after unscheduled official sovereign debt rating announcements around the world.[1] We find that in lower institutional quality countries, where information leakage is more likely (Michaelides et al., 2015, henceforth MMNP), the

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[1] Our FX sample accounts for about 58% of the global FX volume (BIS, 2013).

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currency depreciates before the official sovereign debt downgrade announcement, the announcement becomes a nonevent (Bhattacharya et al., 2000), and the depreciation is permanent, suggesting that sovereign risk is a priced risk factor in currency markets.[2] Using both media-based and market-based measures that control for public information, we build the argument that information leakage reveals fundamental information ahead of the official announcement, thereby providing a direct test of private information in FX markets.

Private information in FX markets is not new in the literature. Lyons (2001) emphasizes the deficiencies of the public information approach and analyses the role of a "dispersed information approach" in FX rate determination (resulting from differential interpretation of public news), where market microstructure and order flow play a key role. This approach implies the presence of private information that gets incorporated into prices through order flow. Private information could be associated with dispersed information resulting from differential interpretation of public news or could be unique information acquired by a group of (large) investors or investors that are followed by other market participants. Evans and Lyons (2008) further emphasize that FX order flow causally affects FX rates and interpret order flow as a proxy for private information that gets transmitted incrementally to FX rates. Consistent with these findings, Ito et al. (1998) use the abolition of a trading restriction during lunchtime in the Tokyo market in 1994 to argue for the presence of private information in the FX market. The recent FX scandal involving the 4 p.m. fix[3] also illustrates that private information and intra-day trading can materially affect FX rate determination.

To further build the empirical argument for private information and the importance of sovereign risk in foreign exchange markets, we collect all sovereign debt rating announcements between 1988 and 2012 for the three largest credit rating agencies (Fitch, Moody's, and Standard and Poor's). An event is defined as a change in sovereign debt rating that is not preceded by other changes in sovereign debt ratings (and/or outlooks) changes in the previous 20 days, by any of the three credit rating agencies (CRAs). This becomes the baseline event and is called the first mover sample. We focus on downgrades based on the recent strong empirical finding (that we also replicate) that downgrades are more important than upgrades in affecting returns (whether in the corporate bond market (e.g., Holthausen and Leftwich, 1986; Hand et al., 1992; Berwart et al., 2018; Milidonis, 2013) or in the sovereign debt market (MMNP). We are left with 195 events in 55 countries that are categorized as flexible exchange rate regime countries.

Using daily FX returns we provide strong empirical evidence of negative abnormal (mean-adjusted) local FX returns ahead of public sovereign downgrade announcements. The effect before the public announcement is followed by a significant effect at the announcement. The overall negative effect of the sovereign rating downgrade appears to be permanent, suggesting that sovereign risk is a priced risk factor in currency markets. Moreover, using a conservative split across the median corruption perception index (a proxy for institutional quality) provided by Transparency International, we find differential responses across countries. In particular, we show a significant depreciation ahead of the public announcement and insignificant announcement effects in lower institutional quality countries, as opposed to a significant depreciation only at the announcement for higher institutional quality countries. These findings are consistent with information leakage being more likely in lower institutional quality countries and are consistent with the empirical results in Doidge et al. (2007) that institutional quality at the macroeconomic level can be of first order importance in determining corporate governance quality (for example, how market sensitive information is treated).[4]

Nevertheless, correlation does not necessarily imply causation; FX volatility and abnormal reactions to political and macroeconomic developments might be strong reasons to classify a country as low institutional quality (Khwaja and Mian, 2005). We therefore resort to instrumental variable techniques to make a causal case. We use four variables widely used in this literature: legal origin (La Porta et al., 1998), ethnicity and religion fractionalization (Alesina et al., 2003) and an indicator on whether a country is landlocked or not (Easterly and Levine, 2003). The instrumental variable regressions validate that local currency depreciation is more likely in lower institutional quality countries ahead of the official sovereign debt rating announcements, confirming the robustness of our findings to endogeneity concerns.[5]

Our interpretation that private information in currency markets generates the pre-event depreciation ahead of sovereign debt rating announcements relies on controlling for media-based sources of public information in three different ways. First, we capture public rumors by manually constructing a *SDN* (sovereign downgrade news) indicator: it takes a value of one if there is at least one news article referring to the forthcoming sovereign debt downgrade in the 15 calendar-day period before the announcement, and zero otherwise. Second, we capture all public news about the country of reference that could affect markets using the Thomson Reuters Marketpsych Indices (TRMI). TRMI captures two dimensions of news: the frequency and intensity (sentiment) of references in the news related to the referenced country through two variables called *TRMI buzz* and *TRMI sentiment*, respectively. Third, we use Google searches for the country name around the downgrade events as reported by Google Trends. Our

[2] This is consistent with Della Corte et al. (2016) who find that sovereign risk captured by credit default swap (CDS) spreads predicts FX movements at the monthly frequency.

[3] http://www.bbc.com/news/business-26526905. Payne (2003) is an early contribution illustrating informed trading in spot FX markets.

[4] The results also echo the findings of Bailey et al. (2006) who find that the information environment can be more important than information itself in generating differential responses to earnings announcements across countries.

[5] Our results are also robust to two alternative measures of institutional quality: the World Bank's development classification and the PRS Group (http://www.prsgroup.com/) Law and Order score.

results show that the preannouncement depreciation of the local currency exists both in the absence of rumors ($SDN = 0$) and in the absence of abnormal *TRMI buzz* and *sentiment*. We also show that there is no abnormal Google search activity ahead of the official downgrade announcements. These findings are consistent with the presence of private information in currency markets since the preannouncement depreciation of the local currency cannot be explained by media-based public information.

To further strengthen the private information interpretation, we also control for market-based sources of public information in different ways. First, exchange rate changes may arise from global shocks affecting each country differently, yet our results hold after controlling for global risk factors. Specifically, we control for global price shocks captured by dollar risk and carry risk factors (Lustig et al., 2011; Verdelhan, 2018), as well as world equity index and world bond index returns. Second, exchange rates may react to local stock market returns in light of the evidence in MMNP that local stock market returns fall ahead of official sovereign debt downgrade announcements. We therefore control for local stock market returns, and our results continue to hold. Third, an alternative pricing channel shown by Della Corte et al. (2016) is the predictability of FX rate changes at a monthly frequency using CDS traded instruments. To eliminate the possibility that our results are driven by public information from changes in CDS spreads, we repeat the analysis for countries with no CDS markets. We find that the local currency depreciates ahead of the public announcement, even in the absence of information from CDS markets, further strengthening the private information hypothesis. Finally, we find no statistically significant daily abnormal equity fund outflows before the official announcement, consistent with the private information hypothesis.

We also conduct several robustness checks that confirm the validity of our results. First, we address the potential impact of illiquidity in FX returns. To do this, we estimate the liquidity spread following Corwin-Schultz (2012) model and control for it both in event studies and cross-sectional regressions. In a second robustness check we control for each country's FX return volatility before running event studies and cross-sectional regressions. Other tests include controlling for the level of sovereign rating and whether a sovereign downgrade crosses into the noninvestment grade category. Finally, we extend our sample to include changes in outlooks and watchlist inclusions/exclusions.

Our empirical results are consistent with the findings in MMNP who find evidence for private information leakage ahead of official sovereign debt rating downgrades and a local stock market drop ahead of the official announcement that is larger for lower institutional quality countries. We interpret the FX results as evidence that sovereign debt rating changes reveal fundamental information that affects the FX market. Moreover, private information related to the forthcoming sovereign debt downgrade moves FX markets ahead of the public announcement. Nevertheless, our results are silent on where information leakage might be coming from. Private information might leak during the consultation process between the CRA and local government officials; it might be intentional or inadvertent and can also be of either legal or illegal nature as discussed in Bhattacharya (2014).

Our analysis contributes to the FX determination literature in four important ways. First, we establish that sovereign debt rating changes constitute fundamental information that affects FX markets in a permanent way. Second, we find local currency abnormal depreciations (i.e., after adjusting for global risk factors and local stock market returns) ahead of unscheduled sovereign debt rating downgrades. These depreciations are stronger in lower institutional quality countries; both findings are consistent with the leaked private/inside information hypothesis. Third, we use new measures of public news to show that the pre-announcement currency depreciations are not driven by public information. Fourth, our news measures, namely the *TRMI buzz* and *sentiment* variables, capture not only the quantity but also the content (tone) of news articles, an improvement over the traditional Reuters screen news items used in previous literature to test the dispersion of information hypothesis. Our approach allows us to test the dispersion of information approach in a unique setting, covering a long time period and multiple foreign currencies against the US dollar as an alternative solution to the significant constraint of order flow data availability in many emerging markets.[6]

The remainder of the paper is organized as follows. In Section 2 we derive our empirical hypotheses, and in Section 3 we discuss our data and provide descriptive statistics. The results are reported in Section 4, and Section 5 discusses the public versus private information interpretation. Section 6 considers policy implications, and Section 7 concludes.

2. Hypothesis development

We test two main hypotheses for the effects of private information regarding sovereign debt rating changes in the FX market. We hypothesize that a sovereign debt rating change constitutes fundamental information that is priced in FX markets. Moreover, private information regarding the forthcoming rating change may be associated with a significant pre-announcement FX market reaction.

Della Corte et al. (2016) show that sovereign risk matters for FX rates, even for developed countries when there is no actual default, extending the findings of previous studies that consider sovereign default risk as a natural source for currency crashes (see, e.g., Brunnermeier et al., 2008; Gourio et al., 2013; Chernov et al., 2018; Farhi and Gabaix, 2016; Farhi et al., 2009; Jurek, 2014; Jurek and Xu, 2014). Reinhart (2002) concentrates on defaults and currency crises in emerging markets and finds that even though sovereign default events are often followed by currency crises, credit ratings fail to predict currency crises. Gande and Parsley (2014) examine monthly mutual fund

[6] Our analysis is conducted using daily data as opposed to the intraday analysis conducted by most papers using order flow data. However, it is common in the empirical literature to assume that the impact of information on prices is long lived (permanent) as opposed to pricing errors, which are transient. Evans and Lyons (2008) highlight the importance of using daily data even in the order flow literature.

flows between 1996 and 2002 and find that sovereign downgrades are strongly associated with outflows of capital from the downgraded country, especially if the downgraded country belongs to a lower institutional quality category. If sovereign debt rating changes contain fundamental information about sovereign default risk, we expect that the rating change announcement will be associated with a significant abnormal return reaction in the FX market.

Hypothesis 1. Announcements of sovereign debt rating changes are associated with significant and permanent FX rate reactions.

A significant part of the FX rate determination literature examines the process through which information is incorporated into prices. Traditional models of FX rate determination posit an information environment where all relevant information is publicly known and immediately incorporated into prices. Lyons (2001) points out that there is general agreement that this public information approach is deficient and analyses the role of a "dispersed information approach" in FX rate determination, where market microstructure and order flow play a key role. This approach implies the presence of private information that gets incorporated into prices through order flow. Private information could be associated with dispersed information resulting from differential interpretation of public news or it could be unique information acquired by a group of investors.

Ito et al. (1998) argue that the FX market has an information structure that is skewed toward public information, and thus they claim there is no plausible analog to an inside information argument that is common in equity markets. Consequently, there is a long standing literature on public news and currency prices (see Evans and Lyons, 2008). Part of this literature examines the impact of order flow on prices as evidence of dispersed/private information in FX markets and uses the Reuters Money Market News Headlines (news items observed on screens by market participants) or scheduled announcement of macro variables as proxies of public news.

In previous work on sovereign debt downgrades, MMNP find evidence that local stock markets move before an official sovereign debt downgrade and argue that this arises from information leakage ahead of the public announcement of sovereign debt downgrades. A natural question arises as to whether such information leakage can also affect the FX market, an arguably larger and more diverse market than local stock markets. Information leakage can be associated with a consultation period between CRAs and local government officials after the rating decision is taken and before it is made public, a common practice internationally. Such information leakage may be either private (held by a few investors) or public (a rumor appearing in news channels). MMNP show an abnormal local stock market drop ahead of the official sovereign debt rating downgrade announcement, consistent with the theoretical analysis of information leakage on trading behavior in Brunnermeier (2005). The possibility that fundamental information can leak during a consultation process can be used to test the effects of private information in the FX market. If private information does affect FX determination, we should expect a negative abnormal reaction in FX prices ahead of the public announcement of sovereign debt downgrades.

Hypothesis 2. FX prices move ahead of the public announcement of a sovereign rating change.

Furthermore, MMNP find that information leakage is more prevalent in countries that can be categorized as lower institutional quality, a finding that is consistent with the information leakage hypothesis. Therefore, we expect that the negative reaction should be more prevalent in lower institutional quality countries, where evidence for information leakage is presented in MMNP. Moreover, if leakage is indeed driven by fundamental information, the effect on the FX market should be permanent, as opposed to transient.

Hypothesis 3. In lower institutional quality countries, where information leakage is more prevalent than higher institutional quality countries, FX prices move more ahead of the announcement dates.

An alternative explanation for negative abnormal reactions before official downgrade announcements would be the presence of concurrent bad news unrelated to the downgrade, such as macro, political, or any other market-moving bad news. If there is a large amount of unrelated news prior to the downgrade, the information leakage story is less likely to hold. In the process of testing the hypotheses outlined above, we therefore need to control for the effect of public news on FX determination. To achieve this, we use new measures of public information. First, we use *TRMI buzz* and *sentiment* variables that capture not only the quantity but also the content (tone) of all country-specific news articles, an improvement over the traditional Reuters screen news items used in previous literature. Second, we use various news sources to determine whether leakage is observed (that is, rumors make it to the newswires before the official announcement) or unobserved (does not make it to news outlets before the public announcement). Finally, we use Google searches as an alternative proxy of public information-related activity.

3. Data and descriptive statistics

In this section we provide details regarding the data, any filters used, and descriptive statistics of our final sample. Specifically, in Section 3.1 we describe the collection of sovereign debt rating changes; in Section 3.2 the FX rate regime of the rated countries, and in Section 3.3 the institutional quality proxies employed. In Section 3.4 we explain our approach to control for publicly available information prior to sovereign debt downgrades, and in Section 3.5 we describe the data needed to construct the dollar and carry risk factors. In Section 3.6 we describe our approach to control for information that may be available in CDS markets, and finally in Section 3.7 we provide descriptive statistics for our main variables.

3.1. Changes in sovereign debt ratings

We collect all changes in sovereign debt ratings from the three largest CRAs: Fitch, Moody's, and Standard and Poor's (S&P). Our sample starts in 1994 for Fitch, 1986 for Moody's, and 1983 for S&P. The sample ends in March 2012 because the 2013 regulatory reforms in the European Union (MMNP, 2015) may have affected the information transmission process between CRAs and local government officials during the consultation process before the public announcement of a rating change.

Ratings represent CRAs' opinions about the default risk of sovereigns on their debt (both local and foreign currency debt). In addition to changes in ratings, CRAs frequently publish changes in outlooks or make changes to watchlist inclusions.[7] The latter provide signals of improving (e.g., positive outlooks) or deteriorating (e.g., negative outlooks) credit quality, which are announced either before, or concurrently, with ratings changes. We transform ratings and outlooks changes into a numerical scale that we describe in Appendix A. To derive changes in credit quality for each country, we take differences in the numeric scale over time for each country.

The sample of ratings and outlook changes is composed of 2278 observations, which is the union of 647, 630, and 1001 observations from Fitch, Moody's, and S&P, respectively. The union of ratings changes then undergoes several filters before it reaches the final form used to test our hypotheses. The first filter relates to the removal of duplicates, that is, we remove any changes in ratings for the same country on the same day but from different CRAs. This reduces the sample to 2147.

The second filter relates to the construction of the "First mover" rating agency. Since we are interested in currency movements before the announcement of ratings and outlook changes, then the period preceding the change should be as clean as possible from other information that may affect the currency. To control for contamination in the period before each announcement we construct the First mover rating agency by removing any observations that are led by other changes in ratings and outlooks (by any of the CRAs) in the 20 trading days before the event. This filter reduces the sample to 1902.

3.2. Floating FX rate regimes

The next step in the data construction matches the sample of rating and outlook changes with available currency data for each country from Datastream. The following conditions need to be satisfied: (a) we use bilateral FX rates against the US dollar; (b) the country is not part of a currency union (for example, Euro Area countries are excluded); (c) the currency is not considered a "hard fixer" with respect to FX rate flexibility, a definition we elaborate further below.

To investigate the effects of any variable on FX markets, we first need to ensure a minimum amount of nominal FX rate variability existing for a particular currency against the U.S. dollar. We rely on Ilzetzki et al. (2008) for the categorization of different currency arrangements around the world before 2008. In general, we exclude countries with a currency (or specific periods from their currency) that have "no separate legal tender," a "de facto peg," a currency board arrangement, and any formally announced peg to a major currency or a horizontal band less than or equal to a positive or negative 2%. We adopt a conservative approach as we include countries that allow limited changes in their currency, such as pre-announced crawling bands (within ± 2%) or managed floating regimes among others. Our approach incorporates currencies whose fluctuations could limit the size of the FX returns, as they could be bounded by the constraints mentioned above.

We complement our sample after 2008 with information from Rose (2014) and additional manual searches if needed. In addition, all countries that are deemed "hard fixers" are dropped from our sample. If a country is categorized as a "hard fixer" over a specific time period, then only that period is removed from our sample. The idea is that in a country with a fixed FX rate, the stock market, for instance, might move ahead of an official announcement, but the central bank may maintain the peg and therefore the FX market will not be affected.

We also exclude countries that are part of a larger currency union, since events in the sovereign debt market of one of the smaller countries should not affect the currency of the union. We do acknowledge that to the extent that contagion exists, a currency like the euro might be affected even by events in a smaller component of the Euro Area, like Greece. We err on the side of caution, however, by not considering these countries as this would be testing the joint hypothesis of contagion and reaction in the FX market.

Finally, we remove the events with zero FX daily returns for the entire period (-270,-21), which do not allow the estimation of standard errors following Kolari and Pynnönen (2010). In addition, if there are less than 60 non-missing observation in the estimation period, the event is deleted. In summary, our final sample covers 58% of the global FX volume and comprises 953 changes in ratings and outlooks from the 71 countries listed in Appendix B.[8]

3.3. Institutional quality

To measure the institutional quality of each country, we use the commonly accepted measure of transparency from Transparency International.[9] We use the Transparency International (*TI*) score since its inception in 1995 until 2012 for the countries in our sample. For observations before 1995 we use the same score as year 1995. The *TI* score is produced every year based on surveys sent out globally to industry professionals who score their perception of corruption in different countries using a scale of one to ten. A lower score implies a lower level of transparency (higher corruption).

[7] Throughout the paper we use the word outlook changes to refer to both changes in outlooks but also changes in watchlist inclusion/exclusions by the three CRAs.

[8] According to the Bank of International Settlements Triennial Central Bank Survey in 2013: https://www.bis.org/publ/rpfx13fx.pdf.

[9] See www.transparency.org.

We use two additional measures for institutional quality for robustness purposes. First, we use the country's development classification by the World Bank: any country denoted as high income is classified as developed, while countries with low and medium income are classified as developing. Second, we use the PRS group's law and order score,[10] which ranges from 1 (lowest institutional quality) to 6. PRS describes its law and order score as capturing the "… strength and impartiality of the legal system …" and also "… assesses popular observance of the law …"

As part of our two-stage least squares (TSLS) analysis, we use several instruments for the institutional quality variables. First, we employ the type of legal origin of each country (La Porta et al., 1998), and we categorize each country into the variable *common law*. This variable gets the value of one if there is a common law origin, and zero otherwise. Next, we use the *religion* and *ethnic fractionalization* by Alesina et al. (2003). Additionally, we gather information on whether a country is *landlocked* or not (Easterly and Levine, 2003).

3.4. News analysis

To account for any publicly available information that may be present in the period before rating downgrades, we use three different approaches.

First, we conduct a manual news search in LexisNexis to identify news articles that explicitly refer to a forthcoming downgrade by the specific CRA that actually downgrades the country in the next days. The keywords we use are the name of the country of interest, the name (and variations) of the rating company, and the word "downgrade." After collecting and reading the identified news articles, we create an indicator variable called *SDN* (sovereign downgrade news) that takes the value of one if there is at least one news article in the previous 15 calendar days before a sovereign downgrade, and zero otherwise.

Our second step is broader in nature, as we want to control for all information related to the country of interest (which also captures any public information about the forthcoming downgrade). In this approach we employ the country-specific TRMI, which account for both the number of news articles present but also the sentiment inferred from these articles.[11] News typically used by TRMI includes, but is not limited to, macroeconomic and political events related to each country.

TRMI is produced through a collaboration between Thomson Reuters (which provides the news sources such as Factiva among other news sources[12]) and MarketPsych LLC (which provides the proprietary algorithm identifying and scoring news articles). We use the indices *TRMI buzz* and *TRMI sentiment. TRMI buzz* is a nonnegative number which captures the total number of references (words and phrases) to a specific country. *TRMI sentiment* ranges from -1 to $+1$ and measures the normalized sentiment inferred from news articles. Specifically, the expertise introduced by MarketPsych LLC allows TRMI to use the tone, type of words chosen, and the general context of the article to produce a score that moves in the same direction as overall market sentiment. Both indices are available since 1998 on a daily basis, and they have been firstly used in the academic literature by MMNP.

The third media-based measure of public information is the number of Google searches (i.e., as captured through Google Trends available from January 1, 2004), variants of which have recently been used to capture retail investor interest about particular US stocks (Niessner, 2015; Ben-Rephael et al., 2017). This approach captures active searches for potential rumors of sovereign rating changes.

3.5. Global risk factors

To isolate country-specific information channels, we control for global shocks in different ways. First, we gather daily spot and (one-month) forward foreign exchange rates (versus the US dollar) for all countries in our sample (from Datastream) to construct the dollar and carry risk factors in FX markets (Lustiget al., 2011; Verdelhan, 2018). Following Verdelhan (2018), we use the log spot FX rate ($s_{i,\,t}$), the log (one-month) forward FX rate ($f_{i,\,t}$), the country's nominal risk-free ($r_{i,\,t}$), and the respective US nominal risk-free rate (r_t). Under the assumption that covered interest rate parity holds, $(r_{i,t} - r_t)$ is approximated using $(f_{i,t} - \; s_{i,t})$. Second, we use the MSCI world stock market index to control for global stock returns ($r_{S,\,t}$) and the Citibank world bond index return (WGBI, available since January 1993) to control for world bond returns ($r_{B,\,t}$).

3.6. Credit default swaps (CDS)

To control for relevant pricing information that may be present in CDS markets, we collect all available CDS information from Markit (www.markit.com/). Specifically, we use the daily mid-market spread on the five-year CDS derivative for each country. We classify each event as having an available CDS traded instrument, if there are available data from the CDS market in the period (-270,-21) before the event.

3.7. Descriptive statistics

The final sample comprises 71 countries and 953 rating and outlook changes, separated into 503 changes in ratings (66 countries), and 450 outlook changes (66 countries). Table 1 provides descriptive statistics of the main variables of interest.

The average (median) rating level after the rating change is 9.68 (10), which is near the investment grade category (value of 10). With respect to the institutional quality of the sample, we observe an average (median) *TI* index score of 4.36 (3.7), which varies by a standard deviation of 2.09. Focusing on the other two institutional quality measures, we observe that 85% of the sample falls into the developing economy category according to the World

[10] http://www.prsgroup.com/.

[11] We would like to acknowledge the generous contribution of this dataset by Thomson Reuters. For more information about the *TRMI sentiment* construction, see https://www.marketpsych.com/.

[12] Based on the description of text sources used in TRMI indices, their sentiment coverage uses text sources from 40,000 primary global news sources.

Table 1
Descriptive statistics.
This table shows the descriptive statistics of the main variables used either as control or instrumental variables. Rating is the sovereign debt rating for the rated country after the rating action. *TI* index is the value of the corruption perception index from transparency international (www.transparency.org). Developing is an indicator variable identifying if the country is developing (1) or developed (0) based on the World Bank classification; i.e., nonhigh income and high income, respectively. *PRS law and order* captures institutional quality with respect to the country's law system and its perceived implementation. *Legal origin* is an indicator variable identifying if the country has English law, and zero otherwise (La Porta et al., 1998). The next two variables are from Alesina et al. (2003), and they capture how fractionalized ethnicity and religion are within a country (*Ethnicity fractionalization* and *Religion fractionalization* that takes value between zero and one). *Landlocked* is an indicator variable that takes the value of one if the country is surrounded by land, and zero otherwise. *SDN* is an indicator variable identifying events with at least one Sovereign Downgrade News (*SDN*) article relevant to the forthcoming downgrade in the 15 calendar days before the public sovereign downgrade announcement (this variable is constructed only for downgrades).

Variable	N	Min	Max	Mean	Median	Std dev
Rating	953	1.000	22.000	9.681	10.000	4.425
TI index	953	1.500	10.000	4.363	3.700	2.089
Developing	953	0.000	1.000	0.852	1.000	0.355
PRS law and order	942	1.000	6.000	3.781	4.000	1.339
Legal origin	953	0.000	1.000	0.248	0.000	0.432
Ethnicity fractionalization	953	0.002	0.930	0.381	0.344	0.223
Religion fractionalization	945	0.003	0.860	0.404	0.384	0.224
Landlocked	953	0.000	1.000	0.111	0.000	0.315
SDN	195	0.000	1.000	0.281	0.000	0.450

Bank, while the average (median) PRS law and order score is 3.78 (4).

In terms of the instruments we use, the following information describes our sample: about 25% of the country-year observations are classified as having English law as their legal origin; 11% are landlocked, while the average (country-year) ethnic and religion fractionalization scores are 0.38 and 0.40, respectively.

To get an idea of the presence of publicly available information specific to a forthcoming downgrade (note that this variable is only collected for downgrades), we observe that 28% of forthcoming downgrades have at least one news article in the pre-event window (*SDN* = 1), which serves as strong indication that information about the forthcoming downgrade is rumored in the press before the actual announcement.

To get a broader idea of the number of country-related news articles in the estimation period (−40, −21) across all events (after 1998, which is the starting year of TRMIs) we provide below (untabulated) some descriptive statistics for TRMIs for upgrades versus downgrades. The average (standard deviation) value for *Buzz* is 1752 (3578) for upgrades and 1,987 (4439) for downgrades. To limit the influence of extreme *Buzz* values in our analysis, especially for larger countries, we later work with the logarithmic transformation of *TRMI buzz*. Finally, the average (standard deviation) value of the *TRMI sentiment* in the estimation period is -0.112 (0.064) for upgrades and −0.162 (0.075) for downgrades.

4. FX reactions around the world ahead of sovereign debt rating announcements

4.1. Methodology

4.1.1. Baseline

We conduct event studies using daily data on bilateral FX returns (relative to the US dollar) around the first mover sample of official sovereign debt rating announcements. We use mean-adjusted abnormal FX returns as our baseline specification. In particular, we compute the mean FX return[13] in the estimation period (−270, −21) relative to event day 0.[14] We then calculate abnormal returns over the (−20, +20) event period as the difference between the actual FX return in the event window and the average FX return in the estimation period (abnormal = actual - expected).

4.1.2. Controlling for global risk factors

Since we focus on capturing country-specific information related to sovereign risk, we control for global factors that may affect bilateral currency returns in two ways. First, we use the global risk factor models from Lustig et al. (2011) and Verdelhan (2018) to construct abnormal FX returns. To keep our sample size comparable to the baseline (mean-adjusted FX returns) sample, we only focus on the dollar risk and carry risk factors in our analysis. Second, we also add world equity and bond index returns to this specification.[15]

To estimate the dollar risk factor for day *t*, we first estimate changes in each country's spot FX rate for the day ($\Delta s_{i,\,t}$). We then average across all countries for day *t*. To estimate the carry risk factor for each currency, we sort all currencies based on the forward discount (at a daily basis), categorize sorted currencies into six portfolios, and estimate the difference in the average changes in currency returns in the top portfolio (low interest rate countries) from the bottom portfolio (high interest rate countries). If the currency of interest is included in the top or bottom portfolio, then we exclude it from the calculations.

[13] We use the US dollar/local currency FX quotation to calculate FX returns so that a negative return reflects local currency depreciation.

[14] We do not have the exact timing of the rating announcement, that is, whether it happens before or after the close of business on the event date. We therefore use the window (0, 1) to capture the announcement effect.

[15] We thank an anonymous referee for these recommendations.

Our first model requires estimation of the following equation over the estimation period $(-270,-21)$ of each event, where only the global carry risk and dollar risk factors are used:

$$\Delta s_{i,t} = a_i + \beta_i Carry_t + \gamma_i Dollar_t + \varepsilon_{i,t}. \tag{1}$$

Next, we estimate the expected FX return $(E(\Delta s_{i,\ t}))$ in the testing period $(-20,+20)$, using the estimated coefficients from Eq. (1):

$$E(\Delta s_{i,t}) = a_i + \hat{\beta}_i Carry_t + \hat{\gamma}_i Dollar_t. \tag{2}$$

Finally, we construct abnormal FX returns in the period $(-20,+20)$ as the actual FX return minus the expected FX return:

$$Abnormal(\Delta s_{i,t}) = \Delta s_{i,t} - \ E(\Delta s_{i,t}). \tag{3}$$

Our second model extends Eq. (1) in the estimation period to include global stock index returns $(r_{S,\ t})$ and global bond index returns $(r_{B,\ t})$:

$$\Delta s_{i,t} = a_i + \beta_i Carry_t + \gamma_i Dollar_t + \delta_i r_{S,t} + \tau_i r_{B,t} + \varepsilon_{i,t}. \tag{4}$$

Expected FX returns in the period $(-20,+20)$ are then computed as:

$$E(\Delta s_{i,t}) = a_i + \hat{\beta}_i Carry_t + \hat{\gamma}_i Dollar_t + \hat{\delta}_i r_{S,t} + \hat{\tau}_i r_{B,t}. \tag{5}$$

Abnormal FX returns are then estimated as the actual FX return minus the expected FX return from Eq. (5) in the period (−20, +20).

In all models we use the methodology of Kolari and Pynnönen (2010) to calculate standard errors, as it addresses both event-induced variance and cross-sectional correlation in abnormal returns.

4.2. Empirical results

4.2.1. Local currency responses to ratings announcements

Fig. 1 presents the response of FX returns before, during, and after sovereign debt rating announcements. Panel A (Panel B) presents the cumulative average raw (mean-adjusted abnormal) FX returns after both upgrades and downgrades for the baseline first mover sample. A fall in the figure signifies a depreciation of the currency relative to the US dollar. Fig. 1 illustrates that there is depreciation (appreciation) of the local currency relative to the USD ahead of official downgrades (upgrades). The response holds for both raw and abnormal cumulative returns and tends to be more substantial quantitatively for downgrades than for upgrades.

Table 2 shows the statistical significance of these findings for the baseline case (cumulative average abnormal (mean-adjusted) FX returns). Using the recent Kolari and Pynnönen (2010) method of computing standard errors, we find statistically significant abnormal returns both ahead of official sovereign debt rating announcements and at the announcement. In particular, for downgrades, we find statistically significant abnormal currency depreciation in the pre-event windows $(-20,-3)$, $(-10,-3)$ and $(-5,-3)$ at least at the 5% level. For downgrades, we also show further statistically significant abnormal currency depreciation on the event date window (0,+1) at the 5% level. Moreover, we observe a permanent depreciation of the local currency relative to the USD as is evidenced by the statistically significant overall abnormal return reaction in the windows $(-5,+5)$, $(-10,+10)$ and $(-20,+20)$ around the event and by the non-significant abnormal returns in all post-event windows $(+2,+5)$, $(+2,+10)$, $(+2,+20)$. Finally, the effects around sovereign rating downgrades are also economically significant (−0.746% and −0.954% over the $(-10,+10)$ and $(-20,+20)$ windows, respectively).

Given the lack of predictability in FX markets, especially at the daily frequency, these results are interesting because they illustrate that sovereign risk is priced, there can be an effect both before and at the announcement, and these effects can be permanent. These empirical results are consistent with Hypothesis 1 that sovereign debt rating announcements affect FX returns. They are also consistent with Hypothesis 2 that FX markets move before the official announcement of the sovereign debt rating change.

Consistent with previous literature on the effects of rating announcements, the response of local currencies is stronger before official downgrade announcements rather than official upgrade announcements. Nevertheless, in this instance, even upgrades can have statistically important effects that are consistent with (even though weaker than) the effects of downgrades. Specifically, local currencies appreciate ahead of the upgrade; for example, window $(-10,-3)$, further appreciate on the event date, window $(0,+1)$, and the appreciation is permanent $(-20,+20)$. Nevertheless, the economic magnitudes are typically two to five times lower than the economic magnitudes from downgrades, and we therefore focus on downgrades in what follows.

4.2.2. Does local institutional quality affect the FX response?

4.2.2.1. Baseline case. Given the investigations on FX and London interbank offered rate (LIBOR)manipulation in advanced economies and the results in Gande and Parsley (2014) and MMNP, we proceed to break the sample in two subsamples based on institutional quality. Using the median values in Transparency International's (*TI*) index score to separate countries (events) into a higher and a lower institutional quality category, we can investigate whether the response varies depending on this measure.

Fig. 2 shows that the FX depreciation (mean adjusted, cumulative average abnormal FX returns) ahead of downgrades is larger in lower institutional quality countries. In contrast, at the event, there does not seem to exist a significant reaction for low institutional quality countries, but there seems to exist a sizable effect for high institutional quality countries. The significance of FX reactions in both samples (including the apparent permanence of the effects) is examined in detail in Table 3.

Table 3, Panel A, reports the statistical significance of the results shown in Fig. 2. The cumulative average abnormal (mean-adjusted) FX returns for different windows before the official rating announcement are shown to be statistically significant for downgrades in the high corruption (low *TI* index) countries (at least at the 5% level). Moreover, there is no abnormal reaction at the event windows $(0,+1)$ and $(-1,+1)$. On the other hand, for the low corruption (high *TI* index) countries, there is no statistically

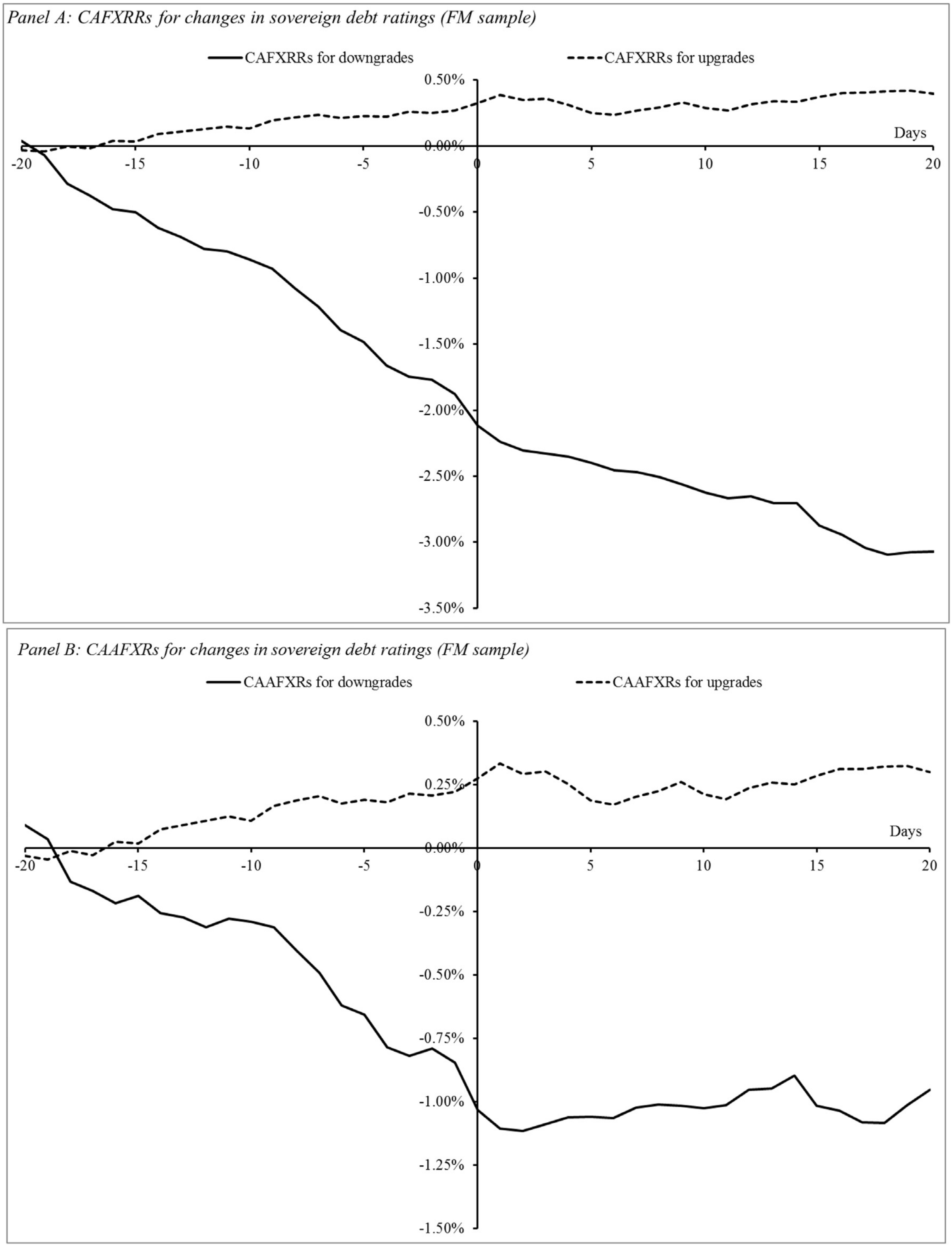

Fig. 1. Foreign Exchange (FX) returns around changes in sovereign ratings.
Panel A shows the cumulative average foreign exchange raw returns (*CAFXRRs*) for FM ratings changes. FM rating changes comprise the union of all rating changes by the three CRAs (Fitch, Moody's, and S&P), which are not led (contaminated) by rating actions (rating and outlook changes) by any of the three CRAs, in the 20 trading days prior to their official announcements. FX returns used are given relative to the US dollar and are winsorized at the 1st and 99th percentiles). Panel B shows cumulative average abnormal foreign exchange (FX) returns (*CAAFXRs*) for the FM rating agency. Abnormal FX returns are estimated as actual returns over the testing period (−20, +20) minus the average FX returns over the estimation period (−270, −21). The final sample comprises 307 upgrades and 195 downgrades.

Table 2
Event study of changes in sovereign debt ratings on the country's currency.
This table presents the event study results of how changes in sovereign debt ratings affect sovereign daily foreign exchange returns (relative to the USD; winsorized at the 1st and 99th percentiles). Results are reported separately for upgrades and downgrades for the FM rating agency. FM rating changes comprise the union of all rating changes by the three ratings agencies (Fitch, Moody's, and S&), which are not led (contaminated) by rating actions (rating and outlook changes) by any of the three rating agencies, in the 20 trading days prior to their official announcements. We report cumulative average abnormal (mean-adjusted) foreign exchange return (*CAAFXR(i,t,)*) for several event windows relative to the event day (day 0). Sample size is denoted by *n*. *p*-values are based on Kolari and Pynnönen (2010) approach. ***, **, and * denote statistical significance (SS) at the 1%, 5%, and 10% level, respectively.

Mean-adjusted CAAFXRs						
	Upgrades ($n=307$)			Downgrades ($n=195$)		
Event window	*CAAFXR* (%)	*p*-value	SS	*CAAFXR* (%)	*p*-value	SS
Pre-event						
(−20,−3)	0.214	0.013	**	−0.819	0.001	***
(−10,−3)	0.090	0.064	*	−0.540	0.002	***
(−5,−3)	0.039	0.098	*	−0.199	0.012	**
At the event						
(0,+1)	0.111	0.005	***	−0.258	0.023	**
(−1,+1)	0.127	0.008	***	−0.315	0.027	**
After the event						
(+2,+5)	−0.146	0.046	**	0.046	0.449	
(+2,+10)	−0.122	0.602		0.080	0.330	
(+2,+20)	−0.034	0.343		0.151	0.191	
Around the event						
(−5,+5)	0.011	0.251		−0.439	0.021	**
(−10,+10)	0.086	0.094	*	−0.746	0.019	**
(−20,+20)	0.299	0.007	***	−0.954	0.035	**

significant pre-event reaction, and the only significant abnormal return reaction occurs at the event.

4.2.2.2. Controlling for global risk factors. In Table 3, Panels B and C, we report results for abnormal returns controlling for global risk factors. In Eq. (2) we control for the dollar risk and carry risk factors and report results in Panel B. Our conclusions do not change: the pre-event window results continue to hold for both samples. Specifically, the pre-event currency depreciation, windows $(-20,-3)$ and $(-10,-3)$, is statistically significant for low transparency countries but not for the high transparency ones. Moreover, there is stronger evidence of a significant depreciation at the event for the high transparency countries relative to the low transparency countries, as the event window $(0,+1)$ is insignificant for the low transparency countries (similar to Panel A). Furthermore, there continues to be a permanent depreciation of the local currency around the event for the low transparency countries, as shown in event windows $(-5,+5)$ and $(-10,+10)$. When the full global risk factor model, (Eq. (4)), is used, the results (Panel C) in the pre-event period and at the event continue to hold, and the coefficient estimates remain qualitatively the same.[16]

These findings strongly support Hypothesis 3, indicating that the pre-event reaction is predominantly concentrated in low institutional quality countries, where information leakage on the forthcoming downgrade is more likely. Furthermore, we observe that the public announcement of the downgrade is rendered a nonevent in low institutional countries, while a significant coefficient is obtained for high institutional countries. This result is consistent with Bhattacharya et al. (2000) who examine shares trading in Mexico and find evidence suggesting that unrestricted insider trading causes prices to fully incorporate information in corporate news announcements before their public release, thus rendering the announcements a nonevent.

4.3. Endogeneity

4.3.1. Baseline

The results reported in earlier sections provide evidence consistent with the hypothesis that there is correlation of FX movements ahead of an official rating announcement and the institutional quality in a particular country. We test for the relationship between institutional quality and FX movements on our entire sample using a series of regression models.

In Table 4, Panel A, we conduct cross-sectional ordinary least squares (OLS) regressions of the cumulative (mean-adjusted) abnormal FX returns, $(-10,-3)$ and $(-20,-3)$

[16] One robustness check is to include changes in outlooks in addition to rating changes in our sample (results available from the authors upon request). This step increases our sample size to 380 negative changes (downgrades and negative outlook changes). Since sample size increases, the statistical power of our analysis improves, and we typically observe higher statistical significance throughout our results. On the other hand, since outlook changes constitute a weaker signal than rating changes (MMNP), the economic significance of results is sometimes slightly lower.

Table 3
Abnormal FX returns conditional on Transparency International (*TI*) Index.
This table presents event studies of how sovereign debt downgrades affect foreign exchange (FX) returns (relative to the USD; winsorized at the 1st and 99th percentiles) conditional on the Corruption Perception Transparency Index (*TI*) score for the FM rating agency. FM rating changes comprise the union of all rating changes by the three rating agencies (Fitch, Moody's, and S&P), which are not led by rating actions (rating and outlook changes) by any of the three rating agencies, in the 20 trading days prior to their official announcements. We report cumulative average abnormal foreign exchange return (*CAAFXR*) using three different regression models. Panel A shows mean-adjusted *CAAFXRs*. Panel B also adjusts for dollar risk and carry risk. Panel C further adjusts for the global equity index return and the global bond index return. Sample size is denoted by *n*. *p*-values are based on the Kolari and Pynnönen (2010) approach. ***, **, and * denote statistical significance (SS) at the 1%, 5%, and 10% level, respectively. *CAAFXRs* are winsorized at the 1st and 99th percentiles.

Panel A: Mean-adjusted CAAFXRs conditional on TI score

Event window	High *TI* (low corruption; $n = 92$)			Low *TI* (high corruption; $n = 103$)		
	CAAFXR (%)	*p*-value	SS	*CAAFXR* (%)	*p*-value	SS
Pre-event						
(−20,−3)	−0.298	0.343		−1.284	0.002	***
(−10,−3)	−0.209	0.384		−0.835	0.002	***
(−5,−3)	−0.078	0.224		−0.306	0.029	**
At the event						
(0,+1)	−0.459	0.037	**	−0.078	0.299	
(−1,+1)	−0.638	0.050	*	−0.025	0.210	
After the event						
(+2,+5)	0.192	0.462		−0.085	0.804	
(+2,+10)	−0.040	0.273		0.188	0.866	
(+2,+20)	0.738	0.501		−0.373	0.265	
Around the event						
(−5,+5)	−0.524	0.125		−0.363	0.013	**
(−10,+10)	−0.888	0.139		−0.619	0.006	***
(−20,+20)	−0.199	0.249		−1.628	0.082	*

Panel B: CAAFXRs (adjusted for dollar risk and carry risk) conditional on TI score

Event window	High TI (low corruption; n = 56)			Low TI (high corruption; n = 57)		
	CAAFXR (%)	*p*-value	SS	*CAAFXR* (%)	*p*-value	SS
Pre-event						
(−20,−3)	0.188	0.807		−1.196	0.003	***
(−10,−3)	0.022	0.950		−0.867	0.002	***
(−5,−3)	0.011	0.489		−0.291	0.030	**
At the event						
(0,+1)	−0.521	0.027	**	−0.103	0.232	
(−1,+1)	−0.673	0.053	*	−0.056	0.152	
After the event						
(+2,+5)	0.070	0.433		−0.074	0.791	
(+2,+10)	−0.346	0.228		0.214	0.804	
(+2,+20)	0.281	0.389		−0.316	0.273	
Around the event						
(−5,+5)	−0.598	0.142		−0.350	0.013	**
(−10,+10)	−1.002	0.152		−0.638	0.007	***
(−20,+20)	−0.210	0.281		−1.498	0.091	*

Panel C: CAAFXRs (adjusted for mean, dollar risk, carry risk, global equity index return, global bond index return)

Event window	High TI (low corruption; n = 77)			Low TI (high corruption; n = 97)		
	CAAFXR (%)	*p*-value	SS	*CAAFXR* (%)	*p*-value	SS
Pre-event						
(−20,−3)	0.209	0.896		−1.289	0.003	***
(−10,−3)	0.055	0.818		−0.886	0.004	***
(−5,−3)	0.004	0.419		−0.313	0.027	**
At the event						
(0,+1)	−0.497	0.025	**	−0.059	0.405	
(−1,+1)	−0.644	0.053	*	0.000	0.320	
After the event						
(+2,+5)	0.004	0.338		−0.052	0.903	
(+2,+10)	−0.332	0.200		0.264	0.493	
(+2,+20)	0.413	0.376		−0.220	0.298	
Around the event						
(−5,+5)	−0.608	0.120		−0.270	0.040	**
(−10,+10)	−0.893	0.136		−0.527	0.027	**
(−20,+20)	0.006	0.271		−1.414	0.110	

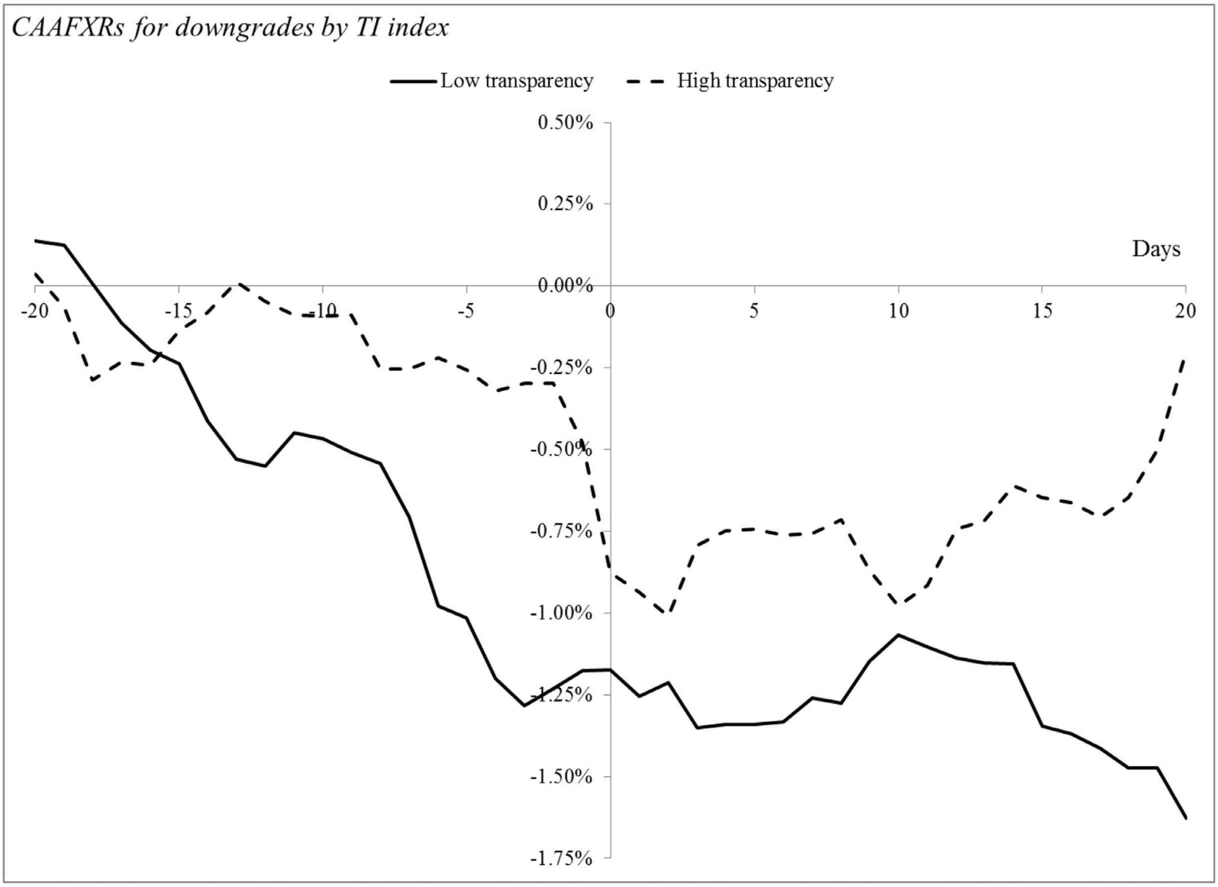

Fig. 2. Cumulative average abnormal FX returns (downgrades) by institutional quality.
The graph shows the cumulative average abnormal FX returns (*CAAFXRs*) for downgrades in sovereign ratings, according to the Corruption Perception Transparency International (*TI*) index score (FM ratings). FM rating changes comprise the union of all rating changes by the three rating agencies (Fitch, Moody's, and S&P), which are not led (contaminated) by rating actions (rating and outlook changes) by any of the three CRAs, in the 20 trading days prior to their official announcements. There are a total of 103 downgrades with low *TI* (high corruption) score and 92 downgrades with high *TI* (low corruption) score. The separation of each category is made at the median value of the ratings FM sample. FX returns used are given relative to the US dollar. Abnormal FX returns are estimated as actual returns over the testing period $(-20,+20)$ minus the average FX returns over the estimation period $(-270,-21)$.

event windows, on the actual value of the *TI* score. We find a statistically significant positive coefficient at the 5% level for the $(-10,-3)$ window. Correlation does not necessarily imply causation, however. FX volatility and abnormal reactions to political and macroeconomic developments might be strong reasons to classify a country as low institutional quality (see the discussion in Khwaja and Mian, 2005, for instance). With this interpretation, institutional quality could be the result of high volatility of asset prices and exchange rates, instead of the other way around.

To make a causal case we therefore proceed and use four instrumental variables widely used in the literature on institutional quality. The first variable is legal origin that makes the distinction between common and civil law (La Porta et al., 1998). The second set of variables involves ethnicity and religion fractionalization (Alesina et al., 2003). Finally, we use an indicator on whether a country is landlocked or not (Easterly and Levine, 2003). These four variables are arguably exogenous given that they were determined earlier than the sovereign debt rating announcements that we study. Moreover, they can be correlated with the measures of institutional quality without being directly related to the cumulative abnormal FX returns that is the dependent variable of interest.

We follow the procedure in Baum et al. (2011) to identify the best instruments for the *TI* score. We first assume that all four instrumental variables (IVs) are valid. The first stage of the TSLS regressions (Table 4, Panel B) shows that three of the four candidate instruments are valid with p-values less than 1%. The legal origin (common versus civil law) indicator is the fourth instrument that is statistically significant at the 5% level, and we therefore report results with the legal origin variable excluded, even though the TSLS results are not affected by either choice.

Table 4, Panel C reports the second stage of the TSLS results for two windows in the pre-event period. The dependent variables are cumulative abnormal (mean-adjusted) FX returns in the $(-10,-3)$ and $(-20,-3)$ periods. The coefficients are strongly statistically and economically significant. Focusing on the economic significance of Table 4, we observe that a standard deviation (2.09) decrease in the *TI* index leads to an abnormal depreciation of 1.271% over

Table 4
Regressions of cumulative abnormal FX returns (*CAFXRs*) on institutional quality.
This table reports OLS and TSLS regressions of cumulative abnormal (mean-adjusted) FX returns (*CAFXR*) on the institutional quality proxy (*TI index*), for FM downgrades (raw FX returns relative to the USD are winsorized at the 1st and 99th percentiles). FM rating changes comprise the union of all rating changes by the three rating agencies (Fitch, Moody's and Standard & Poor's), which are not led (contaminated) by rating actions (rating and outlook changes) by any of the three rating agencies, in the 20 trading days prior to their official announcements. Panel A shows OLS slope coefficient estimate. Panel B shows the first stage of the TSLS regression to address potential endogeneity concerns of the institutional quality proxy (*TI index*). The cumulative abnormal FX return (*CAFXR*) over (*−10, −3)* is used as the dependent variable. Similar results obtain with dependent variable being *CAFXR (−20, −3)*. Four IVs are used. The first one is a *landlocked* indicator variable that takes the value of one if the country is surrounded by land, and zero otherwise. The second IV is the country's *legal origin* that takes the value of one if the legal system is based on English (common) law, and zero otherwise (La Porta et al., 1998). The last two IVs are from Alesina et al., (2003), and they capture how fractionalized ethnicity and religion are within a country (*Ethnicity fractionalization* and *Religion fractionalization* that takes values between zero and one). In Panel B, we test the null hypothesis of whether each instrument is redundant, which we evaluate using the Baum, Schaffer, and Stillman (2011) test statistics (TSLS 1st stage). There are 195 observations in each regression. The initial regression model includes all IVs (Round 1) and after redundant IVs are dropped the regression is conducted again (Round 2). The final list of IVs is determined in Round 2. Panel C shows the 2nd stage of the TSLS regression. Dependent variables are *CAFXR (−20, −3)* and *CAFXR (−10, −3)* from sovereign debt downgrades of the FM rating agency. The final list of IVs used is *landlocked, ethnicity fractionalization,* and *religion fractionalization*. The following outputs are reported: the regressions "Coefficient" and the robust *z*-value and *p*-value of the regression coefficient (regression's constant is not shown). UID stands for under-identification test, OID stands for the over-identification test, and WID stands for the weak-identification test. Furthermore, the Stock-Yogo weak-identification test critical values that correspond to the relative bias and size of the IVs (10% maximal values). Statistical significance (SS) is shown at the 1%, 5%, and 10% levels, using ***, **, and *, respectively.

Panel A: OLS regression of pre-event FX market reaction on institutional quality

$n = 195$	*CAFXR[−10, −3]*				*CAFXR[−20, −3]*			
Institutional quality	Coefficient	*t*-value	*p*-value	SS	Coefficient	*t*-value	*p*-value	SS
TI index	0.0015	1.970	0.050	**	0.0019	1.430	0.153	

Panel B: TSLS (first stage) instrument selection for TI index

$n = 195$	Round 1			Round 2		
Instrument	Test statistic	*p*-value	SS	Test statistic	*p*-value	SS
Legal origin	5.068	0.024	**			
Ethnicity fractionalization	38.243	0.000	***	35.672	0.000	***
Religion fractionalization	7.769	0.005	***	12.702	0.000	***
Landlocked	7.218	0.007	***	10.116	0.002	***

Panel C: TSLS (second stage)regression of pre-event FX market reaction on institutional quality

$n = 195$	*CAFXR[−10, −3]*				*CAFXR[−20, −3]*			
Institutional quality	Coefficient	*z*-value	*p*-value	SS	Coefficient	*z*-value	*p*-value	SS
TI index	0.0043	2.480	0.013	**	0.0061	2.240	0.025	**
Test statistics								
UID (Kleibergen-Paap rk *LM* statistic)		38.920	0.000	***		38.920	0.000	***
OID (Hansen *J*-statistic)		3.057	0.217			0.014	0.993	
WID (Kleibergen-Paap rk Wald *F*-statistic)		37.442				37.442		
WID (Cragg-Donald Wald *F*-statistic)		24.079				24.079		
Stock-Yogo WID 10% relative bias		9.080				9.080		
Stock-Yogo WID 10% size		22.300				22.300		

the $(-20, -3)$ window and 0.889% over the $(-10, -3)$ window. It should be noted that these effects can be amplified substantially by leverage, since many FX trading companies provide their clients with large leverage ratios to enhance returns.[17]

The statistical tests reported in Table 4, Panel C also strongly reject the weak instrumental variable hypothesis. Both the Cragg-Donald (for independent and identically distributed (i.i.d.) errors) and the Kleibergen-Paap (for non-i.i.d. errors) test statistics exceed Staiger and Stock (1997) thresholds and reject the hypothesis of weak IVs. These tests also pass Stock and Yogo (2005) relative bias and relative size tests, providing comfort that identification does not suffer from poor instrumental variable choice.

A number of other robustness tests further support our conclusions (unreported results).[18] Controlling for each currency's FX return volatility by dividing FX returns by their standard deviation in the estimation period, window $(-270, -21)$, thereby transforming the dependent variable into an information ratio, keeps results similar to Table 4. Moreover, using alternative measures of institutional quality (such as the development indicator based on the World Bank classification and the PRS law and order score) do not change our results. Next, adding the entire scale of ratings (as described in Appendix A) as an explanatory variable

[17] Large leveraged trades in FX markets are widespread and one possible cause of concern among regulators, as discussed in https://www.ft.com/content/e03cc41a-93f3-11e4-92dd-00144feabdc0?mhq5j=e6.

[18] For brevity considerations, "unreported" means that the results are available from the authors upon request.

Table 5
Regressions of cumulative abnormal foreign exchange returns (*CAFXRs*) adjusted for global risk factors on institutional quality.
This table reports TSLS regressions of cumulative abnormal FX (*CAFXR*) returns on the institutional quality proxy (*TI Index*) for FM downgrades. (Raw FX returns relative to the USD are winsorized at the 1st and 99th percentiles). In Panel A, *CAFXRs* are adjusted for dollar risk and carry risk. In Panel B, *CAFXRs* are adjusted for dollar risk, carry risk, the global equity index return, and the global bond index return. Dependent variables in both panels are *CAFXR* (−20, −3) and *CAFXR* (−10, −3). FM rating changes comprise the union of all rating changes by the three rating agencies (Fitch, Moody's, and S&P), which are not led (contaminated) by rating actions (rating and outlook changes) by any of the three rating agencies, in the 20 trading days prior to their official announcements. Three instrumental variables (IVs) are used (see Table 4, Panel C). The first one is a *landlocked* indicator variable that takes the value of one if the country is surrounded by land and zero otherwise. The second and third IVs are from Alesina et al., (2003), and they capture how fractionalized ethnicity and religion are within a country (*Ethnicity fractionalization* and *Religion fractionalization* that takes values between zero and one). The following outputs are reported: the regressions "Coefficient", and the robust *z*-value and *p*-value of the regression coefficient (regression's constant is not shown). UID stands for under-identification test, OID stands for the over-identification test, and WID stands for the weak-identification test. Furthermore, the Stock-Yogo weak-identification test critical values that correspond to the relative bias and size of the IVs (10% maximal values). Statistical significance (SS) is shown at the 1%, 5%, and 10% levels, using ***, **, and *, respectively.

Panel A: TSLS (2nd stage) regression of pre-event abnormal FX returns (adjusted for carry risk and dollar risk), on institutional quality

n = 186	*CAFXR[−10, −3]*				*CAFXR[−20, −3]*			
Institutional quality	Coefficient	*z*-value	*p*-value	SS	Coefficient	*z*-value	*p*-value	SS
TI index	0.0048	2.34	0.019	**	0.0071	2.30	0.022	**
UID (Kleibergen-Paap rk *LM* statistic)		36.759	0	***		36.759	0	***
OID (Hansen *J*-statistic)		1.669	0.4342			0.087	0.9572	
WID (Kleibergen-Paap rk Wald *F*-statistic)		32.006				32.006		
WID (Cragg-Donald Wald *F*-statistic)		18.486				18.486		
Stock-Yogo WID 10% relative bias		9.08				9.08		
Stock-Yogo WID 10% size	22.30					22.30		

Panel B: TSLS (2nd stage) regression of pre-event abnormal FX returns (adjusted for carry risk, dollar risk, global equity index return, and global bond index return), on institutional quality.

n = 174	*CAFXR[−10, −3]*				*CAFXR[−20, −3]*			
Institutional quality	Coefficient	*z*-value	*p*-value	SS	Coefficient	*z*-value	*p*-value	SS
TI index	0.0048	2.06	0.039	**	0.0071	2.09	0.037	**
UID (Kleibergen-Paap rk *LM* statistic)		32.721	0	***		32.721	0	***
OID (Hansen *J*-statistic)		1.267	0.5308			0.007	0.9965	
WID (Kleibergen-Paap rk Wald *F*-statistic)		30.392				30.392		
WID (Cragg-Donald Wald *F*-statistic)		18.087				18.087		
Stock-Yogo WID 10% relative bias		9.08				9.08		
Stock-Yogo WID 10% size		22.3				22.3		

in the TSLS regression or using broad rating groups, such as immediately above and below the investment grade, do not change the statistical and economic significance of the *TI* index in Table 4.

To address potential concerns about the impact of illiquidity in FX returns (Karnaukh et al., 2015) we estimate the Corwin and Schultz (2012) liquidity measure on a daily basis and calculate average liquidity over the period $(-40, -21)$ for each event. Either re-running the event study conditional on high and low liquidity (split at the median value across events) or using the liquidity spread as a control variable in the TSLS regressions does not affect the results of Table 4, Panel C.

4.3.2. Controlling for global risk factors

We also run the TSLS regressions (in Table 4, Panel C) using cumulative abnormal returns adjusted for the global risk factors incorporated in Eqs. (2) and (4). The results in Table 5, Panel A show the results using Eq. (2). We observe that the coefficient on the *TI* index is 0.0048 with a *p*-value of 0.019 for the *CAFXR*$[-10,-3]$ case, and for the *CAFXR* $[-20,-3]$ case the coefficient is 0.0071 with a p-value of 0.022. Both coefficients remain approximately the same as in the baseline case and remain statistically significant. Next, we show the results using the full global risk model (Eq. (4)) in Table 5, Panel B. We observe that the coefficient on the *TI* index now being 0.0048 with a *p*-value of 0.039 for the *CAFXR*$[-10,-3]$ case, and for the *CAFXR*$[-20,-3]$ case the coefficient is 0.0071 with a *p*-value of 0.037. Both coefficients remain approximately the same as in the baseline case and remain statistically significant.

4.4. Summary

We conclude from the results in this section that sovereign risk is priced, and therefore local currency returns can be affected by sovereign debt rating downgrades, providing evidence in support of Hypothesis 1 that announcements of sovereign debt rating changes are associated with significant FX rate responses. The results are also consistent with Hypothesis 2, namely that FX prices move ahead of the public announcement of a sovereign rating change and with Hypothesis 3, which is that this is more likely to happen in countries associated with lower institutional quality.

5. Public or private information?

There are two main interpretations of our results. The first interpretation is that public information in the pre-event period can explain the negative market reaction, and this includes the possibility that analysts/investors widely anticipate the specific date and content of the forthcoming sovereign debt downgrade announcement. The second interpretation is that during the consultation period between the sovereign entity and a CRA, there is leakage of information on the specific downgrade that either makes it to the public domain in the form of a rumor or remains private information to some analysts/investors. As a result, trading takes place ahead of the official downgrade announcement, and the currency depreciates.

We control for two broad sources of public information to test whether our strong evidence of negative abnormal FX return reaction ahead of the sovereign rating announcements in low institutional quality countries can be explained away by public information. Specifically, we control for both media-based and market-based public information in what follows.

5.1. Controlling for media-based public information

The first way we control for media-based public information is to manually construct the *SDN* variable, that is, identify rumors in the press related to the specific forthcoming downgrade. The second measure we use is the TRMI dataset (i.e., the *Buzz* and *Sentiment* variables), which captures all local country or US public news that could contain fundamental information consistent with the significant currency depreciation (relative to the US dollar) in our event study results. The third media-based measure of public information is the number of Google searches as captured through Google Trends.

5.1.1. Manual news search (SDN variable)

Fig. 3 plots the cumulative average abnormal (mean-adjusted) FX returns 20 days before and 20 days after the event for all low institutional quality countries conditional on the presence of public rumors on the forthcoming downgrade. Specifically, it reports the cumulative abnormal FX return reaction around the downgrade announcement for the events where rumors on a forthcoming downgrade exist ($SDN=1$) and for the events where rumors do not appear in the public domain ($SDN=0$). The FX depreciation before the public announcement is evident in both subsamples, even though it appears that the magnitude of the effect is bigger in the $SDN=1$ subsample.

Table 6 reports the statistical significance of these results using Kolari and Pynnönen (2010) standard errors. Interestingly, the preannouncement depreciation is statistically significant in low institutional quality countries, regardless of the presence of rumors on the forthcoming sovereign debt downgrade or not. In fact, statistical significance appears to be stronger for the subsample without public rumors, perhaps because of the small size of the subsample with rumors ($n=28$).[19] The statistically significant pre-event abnormal reaction for the $SDN=0$ subsample provides a strong indication that public rumors on the forthcoming downgrade are not sufficient to explain the preannouncement FX depreciation.

5.1.2. News analytics search (TRMI variables)

Is there other concurrent bad news besides the downgrade rumors that might explain our findings? The preannouncement depreciation might be reflecting other concurrent bad public news. To control for all public news that might be affecting the local currency, we test for the existence of abnormal *TRMI buzz* and *TRMI sentiment* conditional on the presence of rumors in our event window. To measure abnormal changes in the news variables (*TRMI buzz* and *TRMI sentiment*), we estimate over the window $(-40, -21)$ and test over the window $(-20, +20)$. To estimate abnormal (log) *Buzz*, we subtract the average (log) *Buzz* over the estimation period from the daily (log) *Buzz* values in the testing period. To estimate abnormal *Sentiment*, we subtract the average sentiment in the estimation period from actual sentiment.

Table 7 presents the results of this analysis. Table 7, Panel A, presents the cumulative abnormal TRMIs for the local country, while Panel B presents the corresponding values for the United States, since all currency returns are measured relative to the US dollar. The findings for the $SDN=0$ sample indicate the absence of significant positive abnormal number of references in the news (i.e., positive abnormal *TRMI buzz*)[20] and the absence of significant abnormal negative sentiment in the news (i.e., negative abnormal *TRMI sentiment*) in all pre-event windows in either the local country or the United States. These results provide strong evidence that neither rumors on the forthcoming downgrade, nor any other public news, can explain away the pre-event currency depreciation. Moreover, the findings in Panel A reveal that both abnormal buzz and sentiment variables are statistically significant at the announcement window $(0, +1)$, confirming the validity of the measures in capturing public news as they both capture the public announcement of the rating downgrade.

5.1.3. Google searches (Google Trends variable)

A third media-based measure of public information is the volume of Google searches as captured through Google Trends (GT), with earliest available data starting on January

[19] Because the sample size for $SDN=1$ is small (28 observations), we also add changes in outlooks (unreported) that increases the sample size to 43. The statistical significance in all pre-event windows improves. In addition, the overall effect around the event for windows $(-10, +10)$ and $(-20, +20)$ becomes significant at the 5% and 10% level, respectively.

[20] The $(-20, -3)$ event window has a statistically significant *TRMI buzz*; however, the coefficient is negative, indicating a lower than average number of references in the news related to the rated entities. In robustness checks where the sample is expanded to include outlook changes as well, no statistically significant results are obtained for the sample with no rumors ($SDN=0$), indicating the absence of other information related to the country (unreported).

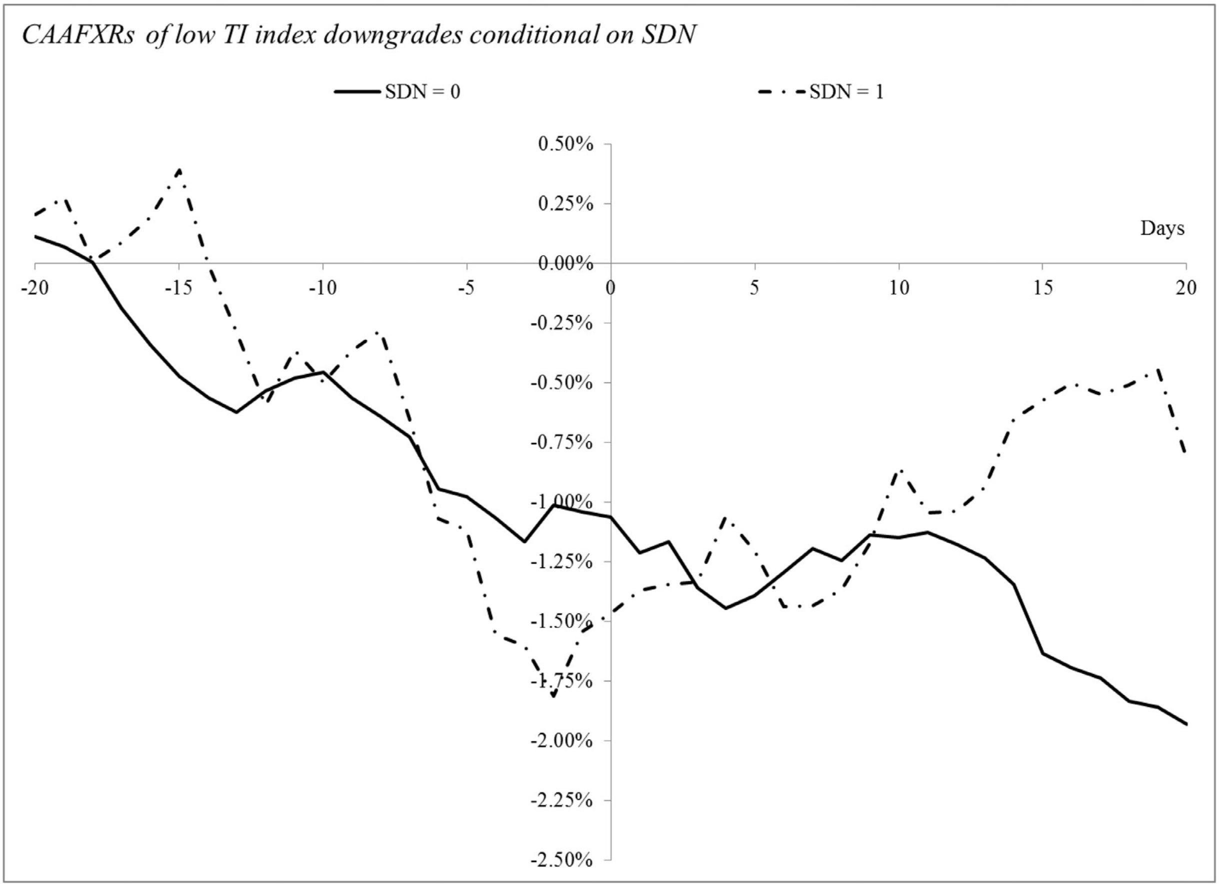

Fig. 3. Cumulative average abnormal FX returns (*CAAFXRs*) for low *TI* downgrades, conditional on rumors.
The graph shows the cumulative average abnormal FX returns for downgrades in sovereign ratings, for low institutional quality, conditional on the presence of Sovereign Downgrade News (*SDN*), 20 days before the announcement. We proxy institutional quality using the Corruption Perception Transparency International (*TI*) index score (FM ratings). FM rating changes comprise the union of all rating changes by the three rating agencies (Fitch, Moody's, and S&P), which are not led (contaminated) by rating actions (rating and outlook changes) by any of the three CRAs, in the 20 trading days prior to their official announcements. FX returns used are given relative to the US dollar. Abnormal FX returns are estimated as actual returns over the testing period $(-20,+20)$ minus the average FX returns over the estimation period $(-270,-21)$. *SDN* is assigned the value of one when there is at least one news item relevant to the downgrade before the announcement, and zero otherwise. *CAAFXRs* for the FM ratings sample (low *TI* downgrades) are shown separately for $SDN=0$ ($n=75$) and $SDN=1$ ($n=28$; mostly for completeness).

1, 2004. Our sample of events in low institutional quality countries comprises 104 downgrades over the entire time period (1988 to 2012), reduced to 42 after January 1, 2004. We follow the same approach used for TRMIs, where we have an estimation period of $(-40,-21)$ and a testing period of $(-20,+20)$, both in trading days. Since GT is given in calendar days, the estimation period in calendar days is $(-60,-31)$ and the testing period is $(-30,+30)$. Because we need the estimation period to start two months before the event, we lose two additional observations from our sample, thus reducing our sample to 40 observations (20 unique countries).

We search using the name of the country (e.g., "Argentina").[21] This search gives results for all 40 observations. To estimate abnormal GT scores in the testing period, we first estimate the average GT score in the estimation period, and we then subtract it from the actual scores in the testing period. Table 8, Panel A reports abnormal GT scores for the full sample of downgrades in low institutional quality countries and Panel B the corresponding values for the subsample with no observed rumors in the media ($SDN=0$). The results in both panels reveal no significant abnormal GT in all pre-event windows, indicating no evidence of abnormal news. Moreover, for both samples we observe a statistically significant increase in the post-event windows $(0,+20)$, $(0,+30)$ at least at the 10% level, which adds confidence that our Google searches measure does respond to the downgrade announcements. We conclude that the statistically significant negative FX abnormal returns in the pre-event period for the $SDN=0$ subsample

[21] We originally searched for "[Country name] downgrade" (e.g., "Argentina downgrade"). This search only gives us 11 (out of 40) nonmissing time series. We interpret the absence of data as an indication that there was not enough "active" searching of information regarding a forthcoming downgrade. Hence, we proceed with the more general search on the country name.

Table 6
Abnormal FX returns (low *TI* downgrades) conditional on Sovereign Downgrade News (*SDN*; rumors).
This table presents cumulative average abnormal FX returns (*CAAFXR*) for the event windows specified, before, at, after and around the announcement of sovereign debt rating downgrades (low *TI*, ratings FM sample) conditional on *SDN*. *CAAFXRs* are mean-adjusted FX returns (raw FX returns relative to the USD are winsorized at the 1st and 99th percentiles). FM rating changes comprise the union of all rating changes by the three rating agencies (Fitch, Moody's, and S&P), which are not led (contaminated) by rating actions (rating and outlook changes) by any of the three rating agencies, in the 20 trading days prior to their official announcements. The *SDN* (Sovereign Downgrade News) indicator variable is assigned the value of one when there is at least one news item relevant to the downgrade before the announcement, and zero otherwise. Results are shown separately for $SDN=0$ ($n=75$) and $SDN=1$ ($n=28$). *p*-values are based on the Kolari and Pynnönen (2010) approach. ***, **, and * denote statistical significance (SS) at the 1%, 5%, and 10% level, respectively.

Cumulative average abnormal FX returns (CAAFXRs)						
Event Window	$SDN=0$ (No rumors; $n=75$)			$SDN=1$ (Rumors; $n=28$)		
	CAAFXR	*p*-value	SS	*CAAFXR*	*p*-value	SS
Pre-event						
(−20,−3)	−1.166%	0.009	***	−1.600%	0.084	*
(−10,−3)	−0.685%	0.011	**	−1.236%	0.072	*
(−5,−3)	−0.222%	0.090	*	−0.532%	0.113	
At the event						
(0,+1)	−0.171%	0.286		0.171%	0.746	
(−1,+1)	−0.200%	0.189		0.442%	0.722	
After the event						
(+2,+5)	−0.179%	0.238		0.166%	0.237	
(+2,+10)	0.064%	0.799		0.519%	0.426	
(+2,+20)	−0.718%	0.254		0.553%	0.853	
Around the event						
(−5,+5)	−0.448%	0.022	**	−0.137%	0.294	
(−10,+10)	−0.667%	0.016	**	−0.488%	0.167	
(−20,+20)	−1.931%	0.109		−0.819%	0.197	

cannot be explained by either rumors in the news nor any other information that could possibly generate abnormal Google searches in the pre-event period.

In summary, our analysis of media-based public information comprises an extensive manual news search in LexisNexis, an automated news analysis by TRMI in over 40,000 text sources (including Factiva) and a more proactive search of information through Google Trends that minimizes the likelihood of missing other significant information in the public domain. Taken together, we interpret the evidence in Section 5.1 as being consistent with the presence of private information in FX markets since we show that publicly available information cannot explain the FX depreciation ahead of the forthcoming sovereign debt downgrade.

5.2. Controlling for market-based public information

There could be additional sources of public information not captured by our news measures. It is perhaps first important to note that the global risk factor models we previously used serve not only the purpose of controlling for sources of global risks, but they can also serve the dual purpose of controlling for additional sources of market-based public information. We examine three additional market-based sources of public information that could potentially explain the preannouncement FX depreciation: information through local stock returns, CDS markets, and equity fund flows.

5.2.1. Information through local stock market returns

We use the local stock market return as an additional possible source of public information that could influence currency returns, given the evidence in MMNP that local stock markets fall ahead of official sovereign debt downgrade announcements.

We first add the local stock market returns ($r_{LS,\,t}$) in Eq. (1), that is, in addition to the carry and dollar risk factors as shown below:

$$\Delta s_{i,t} = a_i + \beta_i Carry_t + \gamma_i Dollar_t + \gamma_i r_{LS,t} + \varepsilon_{i,t}. \tag{6}$$

We run Eq. (6) in the estimation period (−270, −21) and then estimate abnormal FX returns using the predicted values of Eq. (6) as expected returns in the testing period (−20, +20), as in Eq. (3). The event study results conditional on *TI* index (Table 3) continue to hold for the (−10, −3) pre-event window (unreported). We then proceed to replicate the TSLS results reported in Table 4 (Panel C) using this version of abnormal returns. Results are reported in Panel A of Table 9. The coefficient on the *TI* index is 0.0038 (0.0066) with a *p*-value of 0.071 (0.045) for the *CAFXR[−10, −3]* (*CAFXR[−20, −3]*) case. Both coefficients remain approximately the same as in the baseline case of Table 4, Panel C.

In a second, more stringent test, we regress FX actual returns on local equity returns in the period (−20, +20) for each event so that the residuals of these regressions can be used as abnormal FX returns. We find that the coefficients on the local equity returns are not statistically different from zero (10% level or better) in 71% of all the

Table 7

Abnormal *TRMI buzz* and *sentiment* (low *TI* downgrades) conditional on Sovereign Downgrade News (*SDN*; rumors).

This table presents event studies of sovereign debt rating downgrades on two daily variables: *TRMI Buzz* (log), and *TRMI Sentiment*. TRMI stands for Thomson Reuters MarketPsych Indices. *Buzz* is a nonnegative number that captures the total number of references (words and phrases) to a specific country. *Sentiment* is a multi-dimensional, normalized index capturing macro-related, political, and other news that affect sentiment. The *SDN* (Sovereign Downgrade News) indicator variable is assigned the value of one when there is at least one news item relevant to the downgrade before the announcement, and zero otherwise. Results are shown for low *TI* Corruption Perception index events split in subsamples with $SDN=0$ and $SDN=1$ for the FM rating agency. FM rating changes comprise the union of all rating changes by the three rating agencies (Fitch, Moody's, and S&P), which are not led (contaminated) by rating actions (rating and outlook changes) by any of the three rating agencies, in the 20 trading days prior to their official announcements. Relative day is the trading day relative to the event day (day 0). Panel A shows the cumulative average abnormal index for the TRMI variables, *CAAI*, for news related to the downgraded country. There are 51 observations for $SDN=0$ (no rumors) and 21 observations for $SDN=1$ (rumors). Panel B shows the cumulative average abnormal index for the TRMI variables, *CAAI*, for news related to the US, since foreign exchange returns are measured relative to the US dollar. There are 67 observations for $SDN=0$ (no rumors) and 21 observations for $SDN=1$ (rumors). *p*-values are based on the Kolari and Pynnönen (2010) approach. ***, **, and * denote statistical significance (SS) at the 1%, 5%, and 10% level, respectively.

Panel A: Cumulative average abnormal TRMIs (CAAI) for downgraded country

Event window	*SDN = 0 (No rumors)*						*SDN = 1 (Rumors)*					
	TRMI buzz (local)			*TRMI sentiment (local)*			*TRMI buzz (local)*			*TRMI sentiment (local)*		
	CAAI	*p*-value	SS	*CAAI*	*p*-value	SS	*CAAI*	*p*-value	SS	*CAAI*	*p*-value	SS
Pre-event												
(−20,−3)	−2.082	0.060	*	−0.105	0.880		−1.799	0.334		−0.045	0.801	
(−10,−3)	−0.804	0.205		0.004	0.995		−0.277	0.727		0.040	0.970	
(−5,−3)	−0.149	0.779		−0.007	0.766		0.110	0.983		0.030	0.702	
At the event												
(0,+1)	0.524	0.042	**	−0.118	0.002	***	0.719	0.114		−0.093	0.002	***
(−1,+1)	0.620	0.117		−0.128	0.003	***	1.008	0.109		−0.099	0.017	**
After the event												
(+2,+5)	−0.135	0.803		0.059	0.866		0.089	0.917		0.048	0.837	
(+2,+10)	−0.329	0.772		0.012	0.881		0.513	0.807		0.053	0.859	
(+2,+20)	1.075	0.577		−0.036	0.923		−0.360	0.562		0.222	0.545	
Around the event												
(−5,+5)	0.418	0.655		−0.105	0.196		1.334	0.529		−0.050	0.526	
(−10,+10)	−0.432	0.838		−0.142	0.395		1.371	0.703		−0.034	0.728	
(−20,+20)	−0.306	0.962		−0.299	0.652		−1.025	0.641		0.049	0.976	

Panel B: Cumulative average abnormal TRMIs (CAAI) for US

Event window	*SDN = 0 (No rumors)*						*SDN = 1 (Rumors)*					
	TRMI buzz US			*TRMI sentiment US*			*TRMI buzz US*			*TRMI sentiment US*		
	CAAI	*p*-value	SS	*CAAI*	*p*-value	SS	*CAAI*	*p*-value	SS	*CAAI*	*p*-value	SS
Pre-event												
(−20,−3)	0.018	0.841		0.032	0.818		1.019	0.274		−0.114	0.140	
(−10,−3)	0.136	0.751		0.024	0.685		0.713	0.135		−0.044	0.248	
(−5,−3)	0.041	0.864		0.014	0.449		0.398	0.029	**	0.002	0.975	
At the event												
(0,+1)	0.031	0.979		0.001	0.976		0.233	0.049	**	−0.006	0.589	
(−1,+1)	0.087	0.690		0.008	0.769		0.314	0.068	*	−0.010	0.661	
After the event												
(+2,+5)	−0.113	0.432		0.007	0.986		0.481	0.109		−0.017	0.672	
(+2,+10)	−0.359	0.270		0.033	0.888		1.108	0.047	**	0.053	0.760	
(+2,+20)	−0.336	0.517		0.002	0.954		1.429	0.382		−0.021	0.230	
Around the event												
(−5,+5)	0.068	0.931		0.029	0.658		1.304	0.025	**	−0.049	0.354	
(−10,+10)	−0.082	0.694		0.038	0.841		2.246	0.043	**	−0.101	0.289	
(−20,+20)	−0.177	0.691		0.072	0.841		2.874	0.230		−0.102	0.443	

events (70% of the events in low institutional quality countries). This finding leads us to conclude that the local stock returns do not seem to explain exchange rate changes during the (−20, +20) testing period.

5.2.2. Information through the CDS market

We next use CDS spreads as an additional source of public information following Della Corte et al. (2016) who find that CDS spreads can predict FX rate changes at a monthly frequency. To exclude the possibility that our results are driven by market information through the CDS traded instrument, we keep all events for which an active CDS market does not exist. The sample drops from 195 to 121 events and produces qualitatively similar event study results as the ones in Table 3. We do not report these results for brevity considerations but report their regression-equivalents next.

Table 8

Abnormal Google Trends for low institutional quality (TI) events.

This table presents event studies conducted on the daily volume of Google searches for the name of the country being downgraded. Data from Google Trends are used, available since January 1st, 2004. The estimation period used is the calendar day window (−60,−31), relative to the announcement date (day 0) of the country's sovereign debt downgrade. Testing period comprises the calendar day window (−30, +30). Abnormal Google Trends are the mean-adjusted Google Trends in the testing period minus the average Google Trends in estimation period. Panel A shows the results for the subsample of high Corruption Perception Transparency Index (*TI*) score for the FM rating agency. FM rating changes comprise the union of all rating changes by the three rating agencies (Fitch, Moody's and S&P), which are not led by rating actions (rating and outlook changes) by any of the three rating agencies, in the 20 trading days prior to their official announcements. Panel B shows the subsample of low *TI* events with no rumors of a downgrade ($SDN=0$) before the event. Sample size is denoted by n. Statistical significance (SS) at the 1%, 5%, and 10% level, with ***, **, and * respectively.

Panel A: Abnormal Google Trends for downgrade events with low TI score

Event window	Low TI (n = 40)		
	Abnormal Google Trends	p-value	SS
(−30,−5)	9.195	0.679	
(−17,−12)	6.645	0.335	
(−15,−5)	8.983	0.440	
(0,10)	21.508	0.172	
(0,20)	52.033	0.053	*
(0,30)	72.558	0.058	*

Panel B: Abnormal Google Trends for downgrade events with low TI score and no rumors

Event window	Low TI & SDN = 0 (n = 30)		
	Abnormal Google Trends	p-value	SS
(−30,−5)	25.356	0.370	
(−17,−12)	7.833	0.387	
(−15,−5)	18.439	0.209	
(0,10)	36.406	0.065	*
(0,20)	86.783	0.008	***
(0,30)	128.361	0.005	***

Table 9, Panel B shows the TSLS results using the baseline sample with no available CDS traded securities. The dependent variables are cumulative abnormal (mean-adjusted) FX returns in the (−10, −3) and (−20, −3) period. The statistical significance of the coefficient estimates is now at the 5% and 10% level respectively, with the magnitude of the coefficient estimates being larger than the results for the full sample in Table 4, Panel C, indicating again that our previous results hold in the absence of potential public information from CDS markets.

5.2.3. Information through equity fund flows

Another measure possibly containing public information is equity fund flows. The negative foreign exchange abnormal returns we show in the pre-event period might be driven by equity outflows from the local stock markets and not by private information in the FX markets. Gande and Parsley (2014) use monthly equity flow data for the period 1996 to 2002 and show an increase in outflows in the month of a sovereign downgrade with less corrupt countries experiencing lower outflows.

We obtain the EPFR fund flows database used by Gande and Parsley (2014), which is now available at a daily frequency since 2007.[22] Consistent with our analysis, we use equity fund flow data at a daily frequency, resulting in 39 downgrade events (18 and 21 in high and low *TI* score, respectively). Adding outlook changes increases the sample size from 39 to 105 (52 and 53 in high and low *TI* score, respectively). We use the ratio of net flows at day t to asset value at t-1 as the net equity flow at day t (*EFF*). We then conduct event studies on variable *EFF* by subtracting the average *EFF* in the estimation period of (−270, −21) from the actual *EFF* in the testing period of (−20, +20) to compute the *abnormal EFF*. Next, we calculate cumulative average abnormal *EFFs* (*CAEFFs*) over several windows conditional on transparency.

Results (unreported) show no statistically significant *CAEFFs* before the downgrade announcements for both samples (rating changes only, rating and outlook changes). The absence of significant outflows before the official announcement is consistent with the private information hypothesis. Future research can shed more light into the potential association between equity flows and currency markets around sovereign downgrades.

[22] Gelos and Wei (2005) also use this dataset at a monthly frequency (http://www.epfrglobal.com/).

Table 9

Controlling for other sources of market-based information.

This table reports TSLS regressions of cumulative abnormal FX (*CAFXR*) returns on the institutional quality proxy (*TI index*) for FM downgrades (raw FX returns relative to the USD are winsorized at the 1st and 99th percentiles). FM rating changes comprise the union of all rating changes by the three rating agencies (Fitch, Moody's, and S&P), which are not led (contaminated) by rating actions (rating and outlook changes) by any of the three rating agencies, in the 20 trading days prior to their official announcements. In Panel A, *CAFXRs* are estimated by adjusting for dollar risk, carry risk, and local stock index returns. In Panel B, *CAFXRs* are adjusted for the mean FX in the estimation period and are used only for those countries which do not have an active CDS market in the period of one year before the downgrade announcement. The cumulative abnormal FX return (*CAFXR*) over $(-10, -3)$ and $(-20, -3)$ are used as the dependent variable. Three IVs are used (similar to Table 4 Panel C). The first one is a *landlocked* indicator variable that takes the value of one if the country is surrounded by land and zero otherwise. The other two IVs are from Alesina et al. (2003) and they capture how fractionalized ethnicity and religion are within a country (*Ethnicity fractionalization* and *Religion fractionalization* that takes values between zero and one). Dependent variables are *CAFXR* $(-20, -3)$ and *CAFXR* $(-10, -3)$ from sovereign debt downgrades of the FM rating agency. The following outputs are reported: the regressions "Coefficient" and the robust z-value and p-value of the regression coefficient (regression's constant is not shown). UID stands for under-identification test, OID stands for the over-identification test, and WID is the weak-identification tests. Furthermore, the Stock-Yogo weak-identification test critical values that correspond to the relative bias and size of the IVs (10% maximal values). Statistical significance (SS) is shown at the 1%, 5%, and 10% levels, using ***, **, and *, respectively.

Panel A: TSLS (second stage) regression of pre-event abnormal FX returns (adjusted for dollar risk, carry risk, and local stock index returns) on institutional quality

$n = 142$	$CAFXR[-10, -3]$				$CAFXR[-20, -3]$			
Institutional quality	Coefficient	z-value	p-value	SS	Coefficient	z-value	p-value	SS
TI index	0.0038	1.800	0.071	*	0.0066	2.01	0.045	**
Test statistics								
UID (Kleibergen-Paap rk *LM* statistic)		40.656	0.000	***		40.656		***
OID (Hansen *J*-statistic)		5.780	0.0556			3.367	0.000	
WID (Kleibergen-Paap rk Wald *F*-statistic)		46.464				46.646	0.1857	
WID (Cragg-Donald Wald *F*-statistic)		16.605				16.605		
Stock-Yogo WID 10% relative bias		9.080				9.080		
Stock-Yogo WID 10% size	22.300	22.300				22.300		

Panel B: TSLS (second stage) regression of pre-event (mean-adjusted) abnormal FX returns (without CDS traded instruments) on institutional quality

$n = 121$	$CAFXR[-10, -3]$				$CAFXR[-20, -3]$			
Institutional quality	Coefficient	z-value	p-value	SS	Coefficient	z-value	p-value	SS
TI index	0.0052	2.140	0.033	**	0.0070	1.760	0.078	*
Test statistics								
UID (Kleibergen-Paap rk *LM* statistic)		21.253	0.000	***		21.253	0.000	***
OID (Hansen *J*-statistic)		3.000	0.223			0.986	0.611	
WID (Kleibergen-Paap rk Wald *F*-statistic)		19.088				19.088		
WID (Cragg-Donald Wald *F*-statistic)		12.500				12.500		
Stock-Yogo WID 10% relative bias		9.080				9.080		
Stock-Yogo WID 10% size		22.300				22.300		

Overall, we interpret these empirical findings as showing that local currencies tend to depreciate prior to public sovereign debt rating downgrades, especially in low institutional quality countries, even after controlling for market-based sources of public information.

6. Policy implications

Our results are consistent with leaked information generating trading that affects currency (FX) markets. In particular, such events are more likely to take place in countries with lower institutional quality with the FX rate depreciating ahead of official sovereign debt rating announcements. Our results are consistent with MMNP who find evidence of leakage of a similar type in a larger cross-section of countries and show this can destabilize stock markets.

One immediate question is where the leaked information could be coming from. There has been a lot of attention recently on this topic, in particular from European regulators. The European Securities and Markets Authority (ESMA) published a report in 2013[23] stating that ESMA ". . . is concerned that confidential information has been passed on to third parties who should not be privy to it." With respect to timeliness, ESMA observed ". . . significant and frequent delays between the decision taken by the rating committee and the publication of sovereign ratings. In particular, there were instances of publication of ratings more than five days after the rating decision had been approved by the rating committee and, in at least one case even two weeks after the date of the committee." Our results validate the fears of ESMA that information leakage can be an important problem to be tackled during

[23] See https://www.esma.europa.eu/sites/default/files/library/2015/11/2013-1780_esma_identifies_deficiencies_in_cras_sovereign_ratings_processes.pdf.

a sovereign debt rating consultation process. Partly as a response to these concerns, ESMA now forces sovereign debt rating agencies to issue their reports after the close of business on a predetermined Friday afternoon, unless exceptional circumstances warrant earlier action. Interestingly, many rating agencies now follow this voluntarily for sovereign debt rating announcements in non-European countries.

Our results also indicate that information leakage might be a very serious concern for regulators, given the size of the FX market and the corresponding derivatives associated with it. ESMA virtually transformed sovereign rating actions from unscheduled to scheduled events, partly controlling for potential leakage of information by turning the spotlight onto the period before the announcement. However, since the consultation between the CRAs and the rated entity is still in place, the potential for leakage of information still exists. Hence, one radical option would be to avoid the consultation process altogether. For example, since July 1, 2017 the UK Office for National Statistics stopped sending prereleases of economic data to government ministers, citing the negative detriment to public trust from early access to such statistics.[24] Avoiding the consultation period altogether could also increase the reputational cost of inaccurate analysis and could therefore strengthen the quality of the sovereign debt rating analysis. A second option would be to limit the window between the consultation and the official announcement. Moreover, following uniform worldwide rules with regards to the timing of official announcements might further safeguard information confidentiality.

7. Conclusion

During the consultation process between CRAs and local government officials, prior to the official announcement of sovereign debt rating downgrades, leakage of information is likely, especially in countries with lower institutional quality. Using the universe of changes in sovereign downgrades by the three largest CRAs since 1988, we test whether leaked private information during the consultation process can influence FX markets. We find evidence consistent with information leakage before the public announcement of a sovereign rating downgrade, generating a depreciation in the local currency (relative to the US dollar) prior to the official announcement.

We also find evidence of a causal link between institutional quality and negative pre-announcement abnormal FX returns, showing that sovereign debt risk is priced into FX markets at the daily level. More importantly, we find evidence that private information can move FX markets, and therefore information leakage can be a concern to regulators and could potentially apply to other important economic data releases. The recent ban of prereleasing economic data to ministers in the UK is consistent with these concerns, and further research with more granular datasets can shed further light on these issues.

Appendix A. Rating scale

Rating scales used by Fitch, Moody's, and S&P are transformed into a common numerical scale, as shown in the table below. Since outlooks and watchlist inclusions/exclusions are typically an intermediate step before a rating action, we treat them with a weight of half a notch. For example, a positive (negative) outlook would decrease (increase) the numerical rating by 0.5, e.g. from 7 to 6.5 (7.5). The lowest investment grade rating has a numerical value of 10.

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

Appendix B. Countries of interest

To examine the impact of sovereign rating changes on currency markets, the currency of the country of interest needs to have a minimum amount of nominal exchange variability relative to the US dollar. This variability needs to exist for at least one year before the official announcement of the sovereign debt rating change and at least 40 days after it. We use Ilzetzki et al. (2008) and Rose (2014) to make this selection as described in Section 3.2 of the paper. The 71 countries of interest are

[24] In the June 15, 2017, letter to the chair of the UK Statistics Authority Board, the National Statistician explicitly mentions the relevant tradeoffs: "the public benefit likely to result from prerelease access to ONS statistics is outweighed by the detriment to public trust in those statistics likely to result from such access." https://www.ons.gov.uk/aboutus/transparencyandgovernance/onsdatapolicies.

1	Argentina	26	Ireland	51	Poland
2	Australia	27	Israel	52	Portugal
3	Azerbaijan	28	Italy	53	Romania
4	Belarus	29	Jamaica	54	Russia
5	Botswana	30	Japan	55	Rwanda
6	Brazil	31	Kazakhstan	56	Serbia
7	Cambodia	32	Kenya	57	Seychelles
8	Canada	33	Korea	58	Singapore
9	Cape Verde	34	Kuwait	59	South Africa
10	Chile	35	Latvia	60	Sri Lanka
11	China	36	Lithuania	61	Sweden
12	Colombia	37	Malaysia	62	Taiwan
13	Costa Rica	38	Malta	63	Thailand
14	Croatia	39	Mauritius	64	Trinidad & Tobago
15	Czech Republic	40	Mexico	65	Tunisia
16	Denmark	41	Morocco	66	Turkey
17	Dominican Republic	42	New Zealand	67	Uganda
18	Finland	43	Nicaragua	68	Ukraine
19	Ghana	44	Nigeria	69	United Kingdom
20	Guatemala	45	Norway	70	Uruguay
21	Honduras	46	Pakistan	71	Vietnam
22	Hungary	47	Papua New Guinea		
23	Iceland	48	Paraguay		
24	India	49	Peru		
25	Indonesia	50	Philippines		

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