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# Financial reporting lag during COVID-19: evidence from flash reporting in Japan

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**ABSTRACT**

Considering the waiver of the 'within 45 days' rule in Japan following the coronavirus disease 2019 pandemic, we conduct empirical investigations to examine the determinants of the timeliness of flash reporting and the market reactions to flash announcements during the pandemic, distinguishing between early and late filers and between the 'waiver' year (2020) and 'non-waiver' year (2021). We find that the complexity of operations and earnings news explain the observed reporting delay. We also find significant support for market reactions to flash reporting, suggesting that the waiver policy has no significant impact on the information value of flash announcements.

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Flash reporting; financial reporting; COVID-19; pandemic; Japan; market reactions

## 1. Introduction

The novel coronavirus disease 2019 (COVID-19) pandemic has hit almost every part of the world; its severity might be more than any other crisis since World War II. As COVID-19 spreads, its full implications on businesses remain unknown (Goodell 2020). In response to COVID-19, financial regulatory authorities worldwide have taken extraordinary measures affecting corporate activities, including annual general meetings, insolvency proceedings, and disclosure rules. In Japan, following the Declaration of a State of Emergency by the Japanese government in April 2020, the Tokyo Stock Exchange (the Exchange) advised listed firms to disclose information that affects investors' decisions in a timely and appropriate manner regardless of the 'within 45 days' rule.[1] Without this concession, delaying firms would be subject to rules 601 and 808 of the Securities Listing Regulations, which would result in delisting these firms from the Exchange. While the Exchange has rarely delisted Japanese firms because of reporting delays, recent cases, such as Goh Iron Works Co. and Sorghum Japan Holdings Corporation in 2017 and 2018, suggest that it is willing to exercise its powers when necessary. The pandemic and the inevitable concession resulted in a significant increase in Japanese firms announcing their flash annual earnings reports after the 45-day threshold, from 21 firms between 2016 and 2019 to 296 firms in 2020.[2] As the pandemic leads us to distinguish between 'early' and 'late' reporting firms, an interesting and important question arises: which factors explain the timeliness of corporate reporting during the pandemic. To answer this question, we investigate the attributes of firms that are quicker or slower in filing their financial reports in the unprecedented COVID-19 reporting environment. We also analyse stock price reactions to early and late reporting firms to examine the difference in the market valuation of these groups of firms.

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Our study is important for capital market participants, particularly investors and regulators dealing with corporate disclosure policies. As information in financial statements is relevant to investors (Charitou, Clubb, and Andreou 2000), the timeliness of this information promotes the decision relevance quality of financial statements. Professional accountants, users of accounting information, and regulatory bodies have highlighted the timeliness requirements of financial information. For example, Section 2(23) of the revised International Financial Reporting Standards Conceptual Framework (IASB 2018) identifies timeliness as one of the four enhancing qualities of financial information.

The timeliness of accounting information affects its relevance and faithful representation. Additionally, the extant literature provides evidence that the timeliness of financial reports is associated with future restatements (Blankley, Hurtt, and MacGregor 2014), mitigates the incidence of insider trading, leaks, and rumours (Owusu-Ansah 2000), and reduces information asymmetry (Healy and Palepu 2001). The timeliness of earnings announcements also seems to have a contagion effect on the share price of firms in the same industry (Yu and Webb 2017), consistent with the informational role of earnings announcements. Delays in releasing financial statements increase the uncertainty associated with investment decisions (Ashton, Willingham, and Elliott 1987).

Using 1212 observations for firms listed on the 1st Section of the Exchange for the financial year ending 31 March 2020, we investigate the determinants of the timeliness of financial flash reports (also known as *Kessan Tanshin* in Japanese). We examine the incentives and complexity of accounting explanations for our observed reporting lag during the COVID-19 pandemic. Additionally, we applied a standard event study methodology to investigate market reactions to flash reporting. Using abnormal returns around the announcements of flash reports as a proxy for information content, we then examine whether there is a significant difference between the information content of flash reporting between early and late filers in the year 2020; the 'within 45 days' waiver year. We also examine whether the information content of flash reporting differs significantly from that of 2021, a 'clean' year with no 'within 45 days' waiver.

Our first incentive-based explanation posits a negative association between the reporting lag and good news information. We use dividends and earnings news as proxies to find support for the hypothesised association between the reporting lag and positive earnings change. Consistent with the literature suggesting no relationship between reporting lag and dividend news (e.g. Conroy, Eades, and Harris 2000), we find that reporting lag and dividend news are insignificantly different from zero. Our second incentive-based explanation draws on proprietary cost theory, which suggests a positive association between reporting lag and the proprietary cost associated with early reporting. We find no evidence that the proprietary cost explanation plays a role in explaining the reporting lag during COVID-19. Finally, consistent with our expectations, the results indicate a statistically significant positive relationship between financial reporting lag and the complexity of accounting. These results are similar for ordinary least squares (OLS) and logistic regressions. Overall, our research presents early evidence on the determinants and market consequences of the timeliness of reporting during COVID-19 and provides a context-based contribution to the strand of the literature examining financial reporting lag. This study contributes to the literature by evaluating one of the financial reporting guidelines for combating macroeconomic crises such as pandemics.

The remainder of this paper is organised as follows. Section 2 presents the research hypotheses; Section 3 discusses the methodology adopted in the research and presents the empirical model; Section 4 presents and discusses the results. Section 5 provides a summary and conclusion of the study.

## 2. Literature review and hypotheses development

The literature on corporate disclosure abounds with attempts to explain the audit report lag and the timeliness of financial reporting. Early studies before 2000 mainly focused on developed countries

(Dyer and McHugh 1975 for Australia; Ashton, Graul, and Newton 1989 for Canada; Bamber, Bamber, and Schoderbek 1993 for the US). More recent work has shifted the focus to emerging markets (Haw, Qi, and Wu 2000 for China; Owusu-Ansah 2000 for Zimbabwe; Al-Ajmi 2008 for Bahrain). However, the determinants and consequences of financial reporting lag for Japanese firms have received no attention in the literature. A recent meta-analysis by Abdillah, Mardijuwono, and Habiburrochman (2019) reports a consensus among researchers on the role of firm and auditor characteristics in the timeliness of earnings reports. We anchor our study to the literature and test four hypotheses to examine the determinants and market reactions to the financial reporting lag for Japanese firms in the pandemic context. The following subsections review the literature and propose the research hypotheses.

### 2.1. Incentives for timely reporting

The literature suggests that managers and auditors have personal and organisational incentives to manage the release of annual reports (for managers see Dye and Sridhar 1995; Kothari, Shu, and Wysocki 2009; Owusu-Ansah 2000; for auditors see Bamber, Bamber, and Schoderbek 1993). For example, Owusu-Ansah (2000) advances two arguments to support a negative relationship between good news and reporting lag. He argues that firms' managers with good news are incentivised to report promptly. Their managerial capital value depends on their performance, which they signal through the early release of positive news. Complementing this argument and from the audit perspective, poorly performing firms are late reporters, requiring more auditing attention, given the potential reputational and litigation risks associated with their audits.

Haw, Qi, and Wu (2000) explain the reporting lag and managerial incentive relationship from the perspectives of the stakeholder theory and internal reporting hypothesis. Stakeholder theory (e.g. Freeman 2010) suggests that when regulatory requirements deny managers the opportunity to hide bad news, they take the alternative option of delaying its release to neutralise its price effect and circumvent stakeholders' monitoring influence. Consistent with the managerial capital argument, the internal reporting hypothesis suggests that the link between compensation and performance in corporate settings motivates managers to 'buy time' to respond to poor performance criticism by delaying annual reports release. In contrast, an early release of the annual report is preferred because outstanding performance is associated with positive stock price movements. Thus, while the two theories advanced different explanations for management's preference for early or delayed reporting, they both predict a negative (positive) relationship between good (bad) news and reporting lag. In line with these theories and empirical evidence, we test the following hypothesis stated in its alternative form.

H1: There is a negative relationship between the COVID-19 reporting lag and good news information.

Several studies suggest managers consider the proprietary cost of timely reporting (e.g. Verrecchia 1983). According to this theory, firms trade off the costs of early release, including franchise value loss, against their benefits. Botosan and Stanford (2005) report that managers of firms reporting excess profits face higher proprietary costs. However, Sengupta (2004) finds mixed evidence of the role of proprietary cost in the reporting lag of earnings releases. More importantly, the view of proprietary cost theory is valid for understanding the behaviour of Japanese firms. The Japanese management literature highlights that modern Japanese firms tend to act based on a relative comparison with rival firms in the industry to which they belong (Yamaguchi 2020). In line with this theory, we test the following incentive-based hypothesis:

H2: There is a positive relationship between COVID-19 reporting lag and proprietary cost.

### 2.1.1. Complexity of accounting[3]

The role of accounting complexity on the timeliness of financial reporting may be more pronounced during the pandemic. This is because government policies regarding social distancing and lockdown measures meant considerable disruptions in various business areas such as logistics, onsite inspection for audits, and workforce absences. Hence, reporting or audit delays were inevitable. Therefore, it is worth examining the accounting complexity impact on the timeliness of earnings announcements during the COVID-19 pandemic.

Pre-COVID-19 research studies on the relationship between accounting complexity and reporting lag have yielded mixed results. Ashton, Willingham, and Elliott (1987) provide early evidence of a positive relationship between operations complexity and audit delays. Sengupta (2004) examines a related issue in his study on the determinants of quarterly earnings reporting lag. He provides evidence of a positive relationship between the reporting lag and operations complexity (see also Habib and Bhuiyan 2011; Jaggi and Tsui 1999). However, conflicting results have been reported by (Blankley, Hurtt, and MacGregor (2014), who find a negative relationship between the operations complexity and audit reporting lag. Al-Ajmi (2008) compounds the research problem by reporting evidence of no relationship between accounting complexity and reporting delays regarding firms listed on the Bahrain Stock Exchange. Notwithstanding these contradictions, we state our hypothesis in the alternative form:

H3: There is a positive relationship between the COVID-19 reporting lag and accounting complexity.

### 2.1.2. Information content of flash reporting

There has been much debate in the accounting literature concerning the value relevance of accounting information (see Barth, Li, and McClure 2021 for the latest literature review). Several studies provide evidence that the announcement of financial information, including annual and interim reports, results in a significant market reaction, as investors use the information content of such announcements to reassess the prospect of announcing firms (Griffin 2003; Li and Ramesh 2009; Nakajima and Inaba 2021; Nishizaki, Takano, and Takeda 2014). Others, such as Dontoh, Radhakrishnan, and Ronen (2004), report that the usefulness of accounting declines in stock trading as many industries have shifted from highly capital-intensive to knowledge-intensive.

The hypotheses below test this enquiry in the context of the COVID-19 pandemic. Few studies have investigated how investors respond to disclosing corporate earnings information during the pandemic. We develop hypothesis H4a as the baseline to examine share price responses to the release of reports during the pandemic crisis.

H4a: There is a significant market reaction to the release of flash reports by Japanese firms during COVID-19.

The next three hypotheses, H4b, H4c, and H4d, are relevant to the market situation during COVID-19. Alford, Jones, and Zmijewski (1994) and others argue that the timeliness of filing disclosure documents affects stock price. Notably, investors consider late submission of earnings disclosure as a sign of management inefficiency (Cao, Chen, and Higgs 2016). Hence, the stock prices of late filing firms fall, and they are often the targets for short-sellers (Dai et al. 2021). While such phenomena were observed under normal circumstances, the market setting during the pandemic considerably differed in at least two aspects: bad news became the norm in many industries, and regulatory bodies temporarily removed the penalty of late submission of corporate discloser documents in many countries, including Japan.[4]

A few studies on disclosure theory, such as Kim, Pandit, and Wasley (2016) and Brennan, Edgar, and Power (2021), suggest that 'bad news is better than no news' in a crisis because any information disclosure helps reduce investors' uncertainty so that they are desperate for more information. This

leads to the conjecture that the availability of disclosed information outweighs the relative earliness or lateness of report submission dates. Based on this view, the three hypotheses reveal how investors perceive timely disclosure in an unprecedented pandemic setting.

H4b: There is a significant market reaction to the release of flash reports after the statutory 45 days (i.e. late filers) during COVID-19.

H4c: There is a significant difference in the market reaction to flash reports for early and late flash filers during the COVID-19 pandemic.

H4d: There is a significant difference in the market reaction to the release of flash reports during the 'waiver' year (2020) and the 'non-waiver' year (2021).

## 3. Methodology

### 3.1. Sample selection

We extracted our sample from the Nikkei Value Search and Timely Disclosure Network (TDnet) operated by the Exchange. Our study population consisted of 2,166 firms listed in the 1st Section of the Exchange. However, to focus on firms whose reporting is susceptible to COVID-19, we required the sample firms to have a fiscal year ending on 31 March 2020. This restriction limited our observations to 1,474 firms. We exclude financial firms and real estate investment trusts consistent with accounting and finance studies. Consequently, our final sample for the regression (event study) analysis consisted of 1,212 (1,281) observations, covering more than 80% of Japanese non-financial firms.

### 3.2. Model specification – firm-level characteristics

To investigate the relationship between reporting lag and our explanatory variables, we estimate the following cross-sectional model using ordinary least squares (OLS) with robust standard errors:

$$\begin{aligned} LEADTIME_j = {} & \beta_0 + \beta_1 SQRSUB_j + \beta_2 INVTA_j + \beta_3 DIVNEWS_j \\ & + \beta_4 EARNSNEWS_j + \beta_5 CAPREQ_j + \beta_6 LISTAGE_j + \sum \beta_k CONTROLS_j + \varepsilon_j \end{aligned} \tag{1}$$

where *LEADTIME* is the number of days between the fiscal year-end and the financial flash report announcement date. We operationalise complexity (H3) using two variables. Several studies suggest that the number of subsidiaries contributes to a firm's operational, auditing, and accounting complexities (Chan, Luo, and Mo 2016; Ng and Tai 1994). Following Jaggi and Tsui (1999), our first measure of complexity is the square root of the number of subsidiaries (SQRSUB). Other studies suggest that firms whose assets include a high proportion of inventory require special attention and verification procedures and take more time to report (Khoufi and Khoufi 2018). Therefore, we define our second measure of complexity as the ratio of inventory to total assets (INVTA). Our incentive variables (H1 and H2) are composites of four variables. H1 examines the role of good news information as an incentive for timely reporting.

Following Al-Ajmi (2008), we measure dividend news (DIVNEWS) as the change in dividend per share. The model also incorporates earnings news (EARNSNEWS), measured as the change in earnings per share. H2 examines the role of proprietary cost theory in the timeliness of reporting. Following Leventis and Weetman (2004), our first measure of proprietary cost is capital requirement (CAPREQ), measured as the ratio of property, plant, and equipment to total assets.

Table 1. Definition of variables.

| Hypothesis | Variable Names | Measures | Source | Exp. Sign |
|---|---|---|---|---|
| *INCENTIVES* | | | | |
| | DIVNEWS | Change in dividend per share | Al-Ajmi (2008) | – |
| | EARNSNEWS | Change in the ratio of net profit to total assets | Al-Ajmi (2008) | – |
| | CAPREQ | Capital requirement – the ratio of fixed assets to total assets | Leventis and Weetman (2004) | – |
| | LISTAGE | Listing status age – natural logarithm of the number of years since IPO | Prencipe (2004) | – |
| *COMPLEXITY* | | | | |
| | SQRSUB | The square root of the number of subsidiaries | Ng and Tai (1994); Chan, Luo, and Mo (2016); Jaggi and Tsui (1999) | + |
| | INVTA | Inventories to total assets | Khoufi and Khoufi (2018) | + |
| *CONTROLS* | | | | |
| | BIG4 | A dummy variable coded '1' if a big four audit firm audits the firm, and '0' otherwise. | Leventis et al. (2004) Al-Ajmi (2008) | – |
| | LEVERAGE | Long-term debt to total assets | Owusu-Ansah (2000); Leventis et al. (2004) | ± |
| | INDUSTRY | A dummy variable coded '1' for an industrial firm and '0' otherwise. | Jaggi and Tsui (1999); Leventis et al. (2004); Ashton, Willingham, and Elliott (1987) | ± |
| | SIZE | Natural Logarithm of total assets | Al-Ajmi (2008); Jaggi and Tsui (1999); Owusu-Ansah (2000) | – |
| | BOARDIND | The proportion of independent directors | Sengupta (2004) | ± |
| DEPENDENT VARIABLE: | | | | |
| | LEADTIME | The number of days between the fiscal year-end and the financial flash report announcement date | | |

Additionally, following Prencipe (2004), we measure proprietary cost as the listing status age (LISTAGE), defined as the natural logarithm of the number of years since the initial public offering.

The literature suggests firm-level variables that help explain the timeliness of financial reporting. Drawing from this literature, we control for the impact of leverage, firm size, board independence, auditor quality, and industry classification in the model. Table 1 presents the definitions of the variables.

### 3.3. Event study methodology

We adopt the standard event study methodology to test our hypothesis regarding market reactions to delayed flash reports (Brown and Warner 1985). We compute abnormal returns as the difference between the actual and expected returns, where the expected return is the normal return without conditioning on the flash announcement event. Specifically, the abnormal return for stock $i$ on day $t$ is expressed as

$$AR_{it} = R_{it} - E(R_{it}|X_t) \tag{2}$$

where $R_{it}$ and $E(R_{it}|X_t)$ are the actual and expected returns, respectively, for stock $i$ and date $t$; $X_t$ is the conditioning information for the predictive model, $AR_{it}$ is the abnormal return of stock $i$ on day $t$. Assuming the market model, we derive the expected return of stock $i$ on day $t$ as

$$R_{it} = \hat{\alpha} + \hat{\beta} Rm_{it} \tag{3}$$

where $\hat{\alpha}$ and $\hat{\beta}$ are the estimated parameters of the market model, $Rm_{it}$ is the return on market index. That is,

$$AR_{it} = R_{it} - \left(\hat{\alpha} + \hat{\beta} Rm_{it}\right) \tag{4}$$

Following Lagasio and Brogi (2021), because we investigate the possible short-term effects of flash reporting, we calculate daily logarithmic returns, which are continuously compounded:

$$R_{it} = \ln(P_{it}) - \ln(P_{it-1}) \tag{5}$$

We use the Nikkei 225 Index returns as a proxy for market returns.[5] We estimated the market model parameters for each security (event) by taking 110 to 11 days before the reporting date as the estimation window.[6] To examine how fast the market responds to the information content of flash reports, we define event windows of various sub-lengths to investigate the evolution of abnormal returns over a wide window. Our choice of a wide (21 days) event window enables us to explore possible information leakage and market drift around the release of flash reports. Therefore, for a sample of N securities, we estimate the cross-sectional average abnormal returns (AAR) for period $t$ as

$$AAR_t = \frac{1}{N}\sum\nolimits_{i=1}^{N} AR_{it} \tag{6}$$

We then calculated cumulative average abnormal return (CAAR) as:

$$CAAR_b = \sum\nolimits_{i=a}^{b} AAR_t \tag{7}$$

A significant deviation of abnormal returns from a zero mean value is consistent with flash reporting information content (value relevance) during the pandemic. We test for the significance of abnormal returns using the standard parametric tests commonly used in the literature: the cross-sectional $t$ statistic, Patell's (1976) $z$ statistics, Boehmer, Masumeci, and Poulsen's (1991) $t$ statistics, and the time series test (Brown and Warner 1985). However, past studies (e.g. Fama 1976) suggest that daily returns do not follow a normal distribution, which is a vital assumption underlying parametric tests. Additionally, Corrado (1989) and Campbell and Wasley (1996) suggest that nonparametric rank tests are better specified and more statistically robust than parametric t-tests in detecting abnormal daily returns. Thus, we also conduct tests of the significance of abnormal returns using the Corrado (1989) nonparametric rank test and the sign-test statistic (Cowan 1992). Event study analyses were conducted using Event Study Metrics software.

## 4. Results

### 4.1. Descriptive statistics

Table 2 presents the descriptive statistics of the variables in the regression model. For comparison and to provide further support for the importance of this study, we also report the reporting lag for 2019. The mean reporting lag for 2020 was 43 days, ranging from 7 to 86 days. While the mean reporting lag was within the regulatory requirement of 45 days, it was significantly higher than the 39 days reported for 2019 (t = 13.12). We interpret this as evidence of a COVID-19 reporting lag. The table also shows that our sample firms have an average of 22 subsidiaries and an INVTA ratio of approximately 11.8%. The independent directors are 34.8% of the board size, comparable to China's (see e.g. Chan, Luo, and Mo 2016) and in line with one-third suggested by Japan's Corporate Governance Code.[7] On average, the sample firms have earnings per share 43% lower than that reported in 2019. However, the reduction in dividends per share in 2020 is approximately 2% on average, supporting the sticky nature of dividends. Additionally, the Big Four auditing firms account for approximately 81% of the sample firms, suggesting a highly concentrated audit market.

**Table 2.** Descriptive statistics.

| Variables | No of obs. | Mean | Std Dev. | Min. | Max. | Nos. of delays |
|---|---|---|---|---|---|---|
| Panel A: LEADTIME | | | | | | |
| 2016 | 1,316 | 42.571 | 37.253 | 1 | 470 | 59 |
| 2017 | 1,330 | 41.852 | 34.162 | 3 | 452 | 28 |
| 2018 | 1,336 | 39.737 | 24.275 | 2 | 410 | 21 |
| 2019 | 1,259 | 38.630 | 7.524 | 2 | 86 | 24 |
| 2020 | 1,259 | 42.738 | 8.388 | 7 | 86 | 296 |
| Panel B: INDEPENDENT VARIABLES | | | | | | |
| SQRSUB | 1,212 | 4.729 | 4.062 | 0 | 38.601 | |
| INVTA | 1,212 | 0.118 | 0.106 | 0 | 0.789 | |
| BIG4 | 1,212 | 0.808 | 0.394 | 0 | 1 | |
| DIVNEWS | 1,212 | −0.021 | 0.485 | −1 | 9 | |
| EARNSNEWS | 1,212 | −0.425 | 2.185 | −17.67 | 10.420 | |
| CAPREQ | 1,212 | 0.282 | 0.179 | 0.001 | 0.949 | |
| LISTAGE | 1,212 | 3.480 | 0.759 | 0.693 | 4.263 | |
| SIZE | 1,212 | 11.721 | 1.601 | 7.275 | 19.472 | |
| BOARDIND | 1,212 | 0.348 | 0.111 | 0.071 | 0.833 | |
| LEVERAGE | 1,212 | 1.149 | 1.253 | 0.021 | 25.012 | |
| INDUSTRY | 1,212 | 0.521 | 0.500 | 0 | 1 | |

The table presents the descriptive statistics for the dependent and independent variables. The final sample comprises 1212 observations 2020. All variables are defined in Table 1.

On average, the sample firms have a leverage ratio of 1.15, a capital requirement ratio of 28.2%, a log of total assets of 11.72, and a log of listing age of 3.48.

Table 3 presents the Pearson correlation coefficients and collinearity statistics. The coefficients are low (below 0.50) except that between SQRSUB and SIZE (the natural logarithm of total assets), which exceeds the rule of thumb of 0.70, as suggested in the literature (e.g. Anderson, Sweeney, and Williams 1993). However, we note that none of our collinearity statistics violates the thresholds Menard (2002) and Studenmund (2001) recommended.

### 4.2. Regressions results

Table 4 presents the results of the empirical model. We report the unstandardised and standardised coefficient estimates for each independent variable in the model. The model is significant with an F-statistics of 7.46 ($p < 0.00$) and an adjusted $R^2$ of 6%. Hypothesis H1 predicts a negative association between reporting lag and good news information. Consistent with this hypothesis, we find a statistically negative relationship between reporting lag (LEADTIME) and management incentives as measured by EARNSNEWS ($t = -3.15$, $p < 0.00$). A one standard deviation increase in EARNSNEWS is associated with a decrease of 8.8% in the standard deviation in LEADTIME. A 1% increase in the ratio of net profit to total assets (EARNSNEWS) would result in a reduction in the reporting lag by about two days.[8] This result supports the idea that management sees benefits in the early release of positive information, possibly influencing market participants. However, while the coefficient on DIVNEWS is negative, it is statistically insignificant, suggesting that earnings-based incentives dominate the dividend-based motivation of management release of information. Hypothesis H2 suggests a proprietary cost incentive-based argument against timely reporting. However, we find no support for this theory in the present study. The coefficients on both proprietary cost measures (CAPREQ and LISTAGE) are insignificantly different from zero at conventional levels.

Hypothesis H3 predicts a positive association between reporting lag and accounting complexity. As shown in Table 4, the coefficient on complexity (SQRSUB) is statistically positively significant ($t = 2.94$, $p < 0.00$). A one-standard deviation increase in SQRSUB is associated with a decrease of 0.143 standard deviations or 14.3% of the standard deviation in LEADTIME. This finding supports the theory that firms with more subsidiaries are more complex in operations, accounting, and

**Table 3.** Correlation matrix of dependent and independent variables.

| | | A | B | C | D | E | F | G | H | I | J | K | L | Collinearity statistics | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | | | | | | | | | | | | | | VIF | Tolerance |
| A | LEADTIME | 1.000 | | | | | | | | | | | | - | - |
| B | SQRSUB | 0.037 | 1.000 | | | | | | | | | | | 2.60 | 0.384 |
| C | INVTA | **0.091** | −0.014 | 1.000 | | | | | | | | | | 1.14 | 0.877 |
| D | BIG4 | −0.028 | **0.137** | −0.040 | 1.000 | | | | | | | | | 1.06 | 0.945 |
| E | DIVNEWS | **−0.078** | −0.017 | **−0.065** | −0.021 | 1.000 | | | | | | | | 1.11 | 0.904 |
| F | EARNSNEWS | **−0.128** | −0.010 | −0.053 | −0.050 | **0.303** | 1.000 | | | | | | | 1.13 | 0.883 |
| G | CAPREQ | −0.045 | **0.063** | **−0.161** | −0.006 | −0.007 | −0.038 | 1.000 | | | | | | 1.18 | 0.846 |
| H | LISTAGE | −0.016 | **0.252** | 0.034 | **−0.059** | −0.032 | −0.019 | **0.202** | 1.000 | | | | | 1.34 | 0.749 |
| I | SIZE | 0.047 | **0.769** | 0.029 | **0.164** | −0.005 | 0.032 | **0.206** | **0.354** | 1.000 | | | | 2.94 | 0.341 |
| J | BOARDIND | **0.070** | **0.164** | −0.023 | **0.100** | −0.002 | −0.041 | **−0.098** | **−0.078** | **0.105** | 1.000 | | | 1.07 | 0.935 |
| K | LEVERAGE | **0.076** | **0.301** | 0.037 | 0.045 | −0.016 | **−0.084** | **0.159** | −0.034 | **0.331** | **0.076** | 1.000 | | 1.26 | 0.791 |
| L | INDUSTRY | **0.108** | **0.085** | **0.243** | −0.028 | −0.056 | **−0.072** | **0.068** | **0.305** | **0.085** | 0.042 | **−0.175** | 1.000 | 1.25 | 0.797 |

The table shows the Pearson correlations between the dependent and independent variables and the collinearity statistics. The sample comprises 1212 observations for the year 2020. Correlations differing from zero at the 0.05 level or better (two-tailed test) are in bold font. All variables are defined in Table 1.

**Table 4.** Determinants of timeliness of financial flash reports.

$$LEADTIME_j = \beta_0 + \beta_1 SQRSUB_j + \beta_2 INVTA_j + \beta_3 DIVNEWS_j + \beta_4 EARNSNEWS_j + \beta_5 CAPREQ_j + \beta_6 LISTAGE_j + \sum \beta_k CONTROLS_j + \varepsilon_j$$

| | | Unstandardised coef. | | Standardised coef. | | |
|---|---|---|---|---|---|---|
| | Exp. sign | *B* | Standard error | Beta | t-value | p-value |
| Intercept | + | 49.88 | 2.76 | | 18.10 | 0.00*** |
| *COMPLEXITY:* | | | | | | |
| SQRSUB | + | 0.30 | 0.10 | 0.143 | 2.94 | 0.00*** |
| INVTA | + | 4.53 | 1.89 | 0.057 | 2.40 | 0.01** |
| *INCENTIVES:* | | | | | | |
| DIVNEWS | − | −0.66 | 0.61 | −0.038 | −1.07 | 0.28 |
| EARNSNEWS | − | −0.34 | 0.11 | −0.088 | −3.15 | 0.00*** |
| CAPREQ | + | −1.47 | 1.43 | −0.031 | −1.03 | 0.30 |
| LISTAGE | + | 0.35 | 0.33 | 0.032 | 1.07 | 0.29 |
| BIG4 | − | −0.47 | 0.63 | −0.022 | −0.74 | 0.46 |
| SIZE | + | −1.10 | 0.26 | −0.209 | −4.16 | 0.00*** |
| BOARDIND | ± | 4.14 | 2.26 | 0.055 | 1.83 | 0.07* |
| LEVERAGE | ± | 0.75 | 0.21 | 0.112 | 3.56 | 0.00*** |
| INDUSTRY | ± | 1.68 | 0.50 | 0.100 | 3.35 | 0.00*** |
| No. of Obs. | | | Adjusted$R^2$ | F | | p-value |
| 1212 | | | 0.06 | 7.46 | | 0.000 |

The table reports the results of the OLS regression of the determinants of financial flash reporting lag for 2020. The dependent variable (LEADTIME) is defined as the number of days between a firm's fiscal year-end and the release date of the financial flash report. ***, ** and * indicate significant at 0.01, 0.05 and 0.1 levels, respectively. All p-values are based on robust standard errors. The independent variables are defined in Table 1.

auditing, thus taking more time to report. Furthermore, in support of this hypothesis, we find that firms with a higher ratio of INVTA are associated with a higher reporting lag; the coefficient on INVTA is positive and statistically significant ($t = 2.40$, $p < 0.05$). A one standard deviation increase in INVTA is associated with a 5.7% decrease in the standard deviation of LEADTIME. Additionally, a 1% increase in INVTA would increase reporting lag by more than six days.[9] These findings are consistent with the claim that firms with overseas subsidiaries have suffered more from the current global pandemic (Financial Services Agency 2020).

Except for auditor quality (BIG4), all the control variables are statistically significant, at least at the 10% level. Our results support a significantly negative relationship between reporting lag and firm size ($t = -4.16$, $p < 0.00$). A one standard deviation in SIZE is associated with a 20.9% decrease in the standard deviation in LEADTIME, and a 1% increase in SIZE would decrease reporting lag by about two days.[10] This result suggests that large firms have more efficient accounting information systems for timely reporting (Owusu-Ansah 2000). We also find support for the role of board independence in reporting lag. The coefficient on board independence (BOARDIND) is positive and statistically significant ($t = 1.83$, $p < 0.10$). A one standard deviation increase in the BOARDIND is associated with a 5.5% decrease in the standard deviation in LEADTIME. This result supports the evidence in the literature on the negative consequences of a high proportion of non-executive directors (Haniffa and Cooke 2002), such as excessive monitoring (Baysinger and Butler 1985). Consistent with Conover, Miller, and Szakmary (2008), we find support for a positive relationship between reporting lag and leverage ($t = 3.56$, $p < 0.00$). A one standard deviation increase in LEVERAGE is associated with a decrease of 11.2% in the standard deviation in LEADTIME. This positive relationship agrees with the perspective that higher leverage is associated with a greater probability of bankruptcy and litigation risk and thus requires more attention, leading to a longer reporting lag (Owusu-Ansah 2000).

Finally, our results also show that industrial (manufacturing) firms are associated with longer reporting lags ($t = 3.35$, $p < 0.00$), consistent with the suggestion that manufacturing firms have defining processes and characteristics that contribute to longer reporting lags (Afify 2009). While the result for BIG4 is unexpected, given the pervasive role of auditors throughout the fiscal year, the 'flash' nature of reporting considered in this study may suggest that auditors have less influence at this stage of financial reporting.

### 4.3. Event study results

We report the results of our analysis of market reactions for various event windows in Tables 5 and 6. Focusing on the narrow event window (−1, 1) and announcement date (day 0) in Panel A of Table 5, the results suggest that flash reporting is value-relevant during COVID-19. We report a significantly positive CAAR of approximately 0.3%, amounting to an annualised return of approximately 198% for the sample of 1281 announcements. These results generally hold for both parametric and nonparametric tests and support our hypothesis (H4a) of a significant market reaction to flash reports, providing evidence that the market relied on flash reports as a vehicle for price discovery during COVID-19.

To investigate whether early filers drive reported significant positive abnormal returns, we separately examine abnormal returns for the two subsamples of early and late filers.[11] As reported in Panels B and C of Table 6, we find evidence of a significantly positive abnormal return for both subsamples except for the narrow event window for late filers. The test of the difference in means (medians) between the returns of the subsamples is also insignificantly different from zero (Table 6). Thus, while we expect timeliness to matter, given the insignificant difference between the market reactions of the two categories of filers, it seems to us that the market is forgiving for 'approved lateness'. These results support our hypothesis (H4b) that there is a significant market reaction to the late release of flash reports. However, contrary to expectations, the results reject the hypothesis (H4c), which predicts a significant difference between the market reaction to early and

**Table 5.** Results of market model estimation of and tests of cumulative average abnormal returns (CAARs).

| Event Window | CAAR | Pos: Neg | T-test time-series | Prob. | T-test cross-sectional | Prob. | Patell z | Prob. | Boehmer, Masumeci, and Poulsen (1991) | Prob. | Corrado rank | Prob. | Sign test | Prob. |
|---|---|---|---|---|---|---|---|---|---|---|---|---|---|---|
| Panel A – All Filers (2020) | | | | | | | | | | | | | | |
| (−10 . . . 10) | 0.012 | 682: 599 | 3.983 | 0.000*** | 3.853 | 0.000*** | 4.2717 | 0.000*** | 4.118 | 0.000*** | 1.088 | 0.277 | 3.556 | 0.000*** |
| (−1 . . . 1) | 0.003 | 626: 655 | 2.803 | 0.005*** | 1.859 | 0.063* | 2.327 | 0.020** | 1.478 | 0.134 | 0.775 | 0.438 | 0.424 | 0.671 |
| (0 . . . 0) | 0.003 | 704: 577 | 5.089 | 0.000*** | 4.053 | 0.000*** | 5.647 | 0.000*** | 4.099 | 0.000*** | 1.982 | 0.048** | 4.786 | 0.000*** |
| Panel B – Late Filers (2020) | | | | | | | | | | | | | | |
| (−10 . . . 10) | −0.002 | 186: 190 | −0.328 | 0.743 | −0.357 | 0.721 | −0.184 | 0.854 | −0.183 | 0.855 | −0.889 | 0.374 | 0.324 | 0.746 |
| (−1 . . . 1) | 0.003 | 184: 192 | 1.193 | 0.233 | 0.866 | 0.387 | 1.280 | 0.201 | 0.829 | 0.407 | 0.832 | 0.406 | 0.118 | 0.906 |
| (0 . . . 0) | 0.004 | 203: 173 | 2.912 | 0.004*** | 2.522 | 0.012** | 2.968 | 0.003*** | 2.302 | 0.021** | 1.622 | 0.105 | 2.078 | 0.038** |
| Panel C – Early Filers – 2020 | | | | | | | | | | | | | | |
| (−10 . . . 10) | 0.017 | 496: 409 | 5.071 | 0.000*** | 4.696 | 0.000*** | 5.201 | 0.000*** | 4.979 | 0.000*** | 1.556 | 0.120 | 4.022 | 0.000*** |
| (−1 . . . 1) | 0.003 | 442: 463 | 2.583 | 0.010** | 1.649 | 0.548 | 1.943 | 0.052* | 1.224 | 0.221 | 0.475 | 0.635 | 0.429 | 0.668 |
| (0 . . . 0) | 0.003 | 501: 404 | 4.174 | 0.000*** | 4.696 | 0.000*** | 4.805 | 0.000*** | 3.400 | 0.001*** | 1.434 | 0.152 | 4.354 | 0.000*** |
| Panel D – Results of estimation of cumulative average abnormal returns (CAARs) using market model (All filers 2021) | | | | | | | | | | | | | | |
| (−10 . . . 10) | −0.010 | 587: 686 | −3.851 | 0.000*** | −3.827 | 0.000*** | −3.687 | 0.000*** | −3.430 | 0.001*** | −0.817 | 0.414 | −1.744 | 0.081*** |
| (−1 . . . 1) | 0.001 | 664: 609 | 1.379 | 0.168 | 0.781 | 0.435 | 3.585 | 0.000*** | 1.840 | 0.066*** | 0.467 | 0.640 | 2.574 | 0.010*** |
| (0 . . . 0) | 0.002 | 690: 583 | 4.474 | 0.000*** | 3.146 | 0.002*** | 6.223 | 0.000*** | 3.810 | 0.000*** | 1.264 | 0.206 | 4.032 | 0.000*** |

The table presents the results of the tests of cumulative average abnormal returns (CAARs) for different event windows for flash reports. Panel A shows the results for 2020 filers, while Panel B and C show the results for late and early filers, respectively. The results for 2021 announcements are presented in Panel D. We estimated CAAR as the residual of a market model regression cumulated over the event windows, with Nikkei 225 value-weighted index as the market index; p-values are two-tailed. The estimation window is −11 to −110 days relative to the announcement date (day 0). Significance at the 1%, 5% and 10% level is denoted by ***, ** and *, respectively.

**Table 6.** Test of difference of means and medians cumulative average abnormal returns (CAARs).

| Panel A – Early filers and late filers | | | | | | | | |
|---|---|---|---|---|---|---|---|---|
| | Early Filers (N = 905) | | Late Filers (N = 376) | | Diff. in Means | | Diff. in Medians | |
| | Mean | Median | Mean | Median | t-stat | p-value | z-stat | p-value |
| Signed Abnormal Returns: | | | | | | | | |
| (−10 . . . 10) | .017 | .012 | −.002 | −.002 | −2.888 | 0.004 | −2.697 | 0.007*** |
| (−1 . . . 1) | .003 | −.001 | .003 | −.002 | −0.209 | 0.835 | −0.044 | 0.965 |
| (0 . . . 0) | .003 | .002 | .004 | .002 | 0.775 | 0.286 | 0.367 | 0.714 |
| Absolute Abnormal Returns: | | | | | | | | |
| (−10 . . . 10) | .082 | .064 | .078 | .058 | −0.913 | 0.361 | −0.672 | 0.501 |
| (−1 . . . 1) | .044 | .032 | .042 | .028 | −0.738 | 0.461 | −0.594 | 0.552 |
| (0 . . . 0) | .020 | .015 | .020 | .014 | −0.149 | 0.882 | 0.176 | 0.860 |
| Panel B – 'waiver' year (2020) and 'non-waiver' year (2021) | | | | | | | | |
| | 'waiver' year 2020 (N = 1281) | | non-waiver" year 2021 (N = 1273) | | Diff. in Means | | Diff. in Medians | |
| | Mean | Median | Mean | Median | t-stat | p-value | t-stat | p-value |
| Signed Abnormal Returns: | | | | | | | | |
| (−10 . . . 10) | .012 | .006 | −.009 | −.007 | 5.401 | 0.000*** | 4.960 | 0.000*** |
| (−1 . . . 1) | .003 | −.001 | .002 | .001 | 0.778 | 0.436 | −0.063 | 0.949 |
| (0 . . . 0) | .003 | .002 | .002 | .002 | 0.771 | 0.441 | 0.835 | 0.404 |
| Absolute Abnormal Returns: | | | | | | | | |
| (−10 . . . 10) | .080 | .063 | .065 | .047 | 5.643 | 0.000*** | 6.035 | 0.000*** |
| (−1 . . . 1) | .043 | .031 | .043 | .032 | 0.015 | 0.988 | −0.281 | 0.779 |
| (0 . . . 0) | .015 | .020 | .012 | .018 | 2.786 | 0.005*** | 4.359 | 0.000*** |

Panel A presents the result of the tests of difference in means (medians) of signed and absolute abnormal returns for early filers and late filers in 2020. Panel B presents the result of the tests of difference in means (medians) of signed and absolute abnormal returns for 2020 and 2021. We estimated CAAR as the residual of a market model regression cumulated over the event windows, with Nikkei 225 value-weighted index as the market index. Significance at the 1%, 5% and 10% level is denoted by ***, ** and *, respectively.

late reporters. Finally, we compared the market reaction to flash reports released during COVID-19 (2020) with those released in 2019 and 2021.[12] While we again find evidence of significantly positive abnormal returns in 2021 (Panel D, Table 5), a test of the difference in means (medians) does not provide evidence that these returns are significantly different from those reported in 2020 (Panel B, Table 5). This suggests that the waiver of the 'within 45 days' rule does not have a significant impact on market reactions. Thus, we reject the hypothesis of a significant difference between the market reactions for waiver and non-waiver years (hypothesis H4d).[13]

### 4.4 Robustness tests

The results support our proposition that the complexity of accounting and incentive-based theories explains the financial flash-reporting lag. However, two potential concerns may question these results. First, a reporting delay phenomenon may have existed before 2020. Second, following the first concern, firms with a history of reporting late before COVID-19 are the same firms that reported late in 2020. Suppose either or both concerns hold, the COVID-19 impact on the reporting lag is modest, and our primary measure of reporting lag (LEADTIME) is questionable. To address these concerns, we first compare the frequency of reporting lag in 2020 to the frequencies of late reporting for the four preceding years (2016–2019). Panel A (Table 2) shows the mean number of reporting days from 2016 to 2020 and the frequency of late reporting. Reporting firms took, on average, 43 days to report during the COVID-19 pandemic. While this is comparable to the means for 2016 and 2017, it is markedly different from the mean for 2018 and 2019. Additional evidence that COVID-19 amplified the reporting lag phenomenon is evident in the significant increase in the frequency of firms reporting late during the pandemic. Table 2 suggests a range of between five -and fourteen-fold increases in the number of firms reporting late in 2020 relative to the last four years.

**Table 7.** Test of difference of means and medians.

| | COVIDLAG1 = 1 | | COVIDLAG1 = 0 | | | | | |
|---|---|---|---|---|---|---|---|---|
| | N = 284 | | N = 928 | | Diff. in Means | | Diff. in Medians | |
| Variable | Mean | Median | Mean | Median | t-stat | p-value | z score | p-value |
| SQRSUB | 5.332 | 4.243 | 4.544 | 3.606 | −2.87 | 0.004 | −4.25 | 0.000 |
| INVTA | 0.129 | 0.120 | 0.114 | 0.095 | −2.04 | 0.041 | −3.16 | 0.002 |
| DIVNEWS | −0.104 | 0.000 | 0.004 | 0.000 | 3.30 | 0.001 | 4.66 | 0.000 |
| EARNSNEWS | −1.040 | −0.231 | −0.236 | 0.093 | 5.49 | 0.000 | 5.60 | 0.000 |
| LISTAGE | 3.543 | 3.828 | 3.461 | 3.555 | −1.59 | 0.111 | −2.43 | 0.015 |
| CAPREQ | 0.289 | 0.278 | 0.280 | 0.256 | −0.73 | 0.462 | −1.43 | 0.154 |
| BIG4 | 0.824 | 1.000 | 0.803 | 1.000 | −0.79 | 0.429 | −0.79 | 0.429 |
| SIZE | 11.831 | 11.593 | 11.687 | 11.479 | −1.32 | 0.187 | −1.42 | 0.154 |
| BOARDIND | 0.357 | 0.333 | 0.345 | 0.333 | −1.55 | 0.122 | −1.07 | 0.285 |
| LEVERAGE | 1.353 | 0.996 | 1.087 | 0.779 | −3.14 | 0.001 | −4.52 | 0.000 |
| INDUSTRY | 0.634 | 1.000 | 0.486 | 0.000 | −4.39 | 0.000 | −4.36 | 0.000 |

The table presents the results of the test of difference of means (and median) between late reporting firms.

Following Conover, Miller, and Szakmary (2008), we estimate logistic regressions for reporting lag with the same explanatory variables as in the empirical model stated in Section 3.2. For this purpose, we benchmarked reporting firms in 2020 against their reporting practices in 2019 to identify firms with a history of reporting early but late in 2020 because of the pandemic or the waiver of penalty for late reporting. We code COVIDLAG1 equals one for firms that delayed 2020 reporting but did not delay in 2019 and zero otherwise. We further examine the reporting lag in the context of the Exchange's requirement for firms to report no later than 45 days after their fiscal year-end. Thus, we defined late reporting as the difference between the days taken and allowed (45 days). We consider this a nondiscretionary reporting lag. Consequently, late reporting (COVIDLAG1) is coded as one for firms with a reporting lag greater than 45 days and zero otherwise.

Consistent with our earlier results, Table 7 shows that late reporters have less good news reporting and greater accounting complexity. The results provide further support for hypotheses H1 and H3. The estimated logistic regression coefficients (Table 8) generally support the conjecture that accounting complexity and good news incentives captured in earnings information explain the financial flash-reporting lag. Additionally, the coefficients on proprietary cost and auditor quality are insignificantly different from zero.

We revisited our analysis of market reactions to flash announcements as our choice of the market model because the return-generating model may have driven the reported results. Specifically, we repeated our tests for abnormal returns using market returns. Our results (unreported) were qualitatively like those in Tables 5 and 6. We find support for significant market reactions to flash reporting (H2), even for firms reporting after the statutory 45 days (H2), but significantly different reactions for late and early filers (H3). The results also reject the prediction of a significant difference in market reactions between waiver and non-waiver years (H4d).

Finally, we acknowledge that our results may suffer from a well-known endogeneity problem, resulting in biased coefficient estimates.[14] We use the lag approach to construct instrumental variables to address this issue. We regressed the dependent variable of each model against the previous year's firm characteristics (rather than contemporaneous values). Consistent with the findings of this study, the results (unreported) show that the coefficients on the lagged variables are qualitatively similar. Therefore, it is unlikely that our results suffer from an endogeneity problem.

## 5. Summary and conclusion

It is difficult to deny the benefits of timely earnings announcements in the stock market, specifically from investors' perspectives. The transparency of the stock market, in part promoted by timely

**Table 8.** Logistic regression of reporting lag.

| | COVIDLAG1 | | | COVIDLAG2 | | |
|---|---|---|---|---|---|---|
| | Odds Ratio | Coef. | z-value | Odds Ratio | Coef. | z-value |
| Intercept | .35 | −1.05 | −1.24 | 0.39 | −0.94 | −1.14 |
| *COMPLEXITY:* | | | | | | |
| SQRSUB | 1.06 | 0.06 | 2.01** | 1.06 | 0.06 | 2.03** |
| INVTA | 1.73 | 0.55 | 0.75 | 1.75 | 0.56 | 0.78 |
| *INCENTIVES:* | | | | | | |
| DIVNEWS | .65 | −0.43 | −1.10 | 0.67 | −0.40 | −1.06 |
| EARNSNEWS | .90 | −0.10 | −2.99*** | 0.91 | −0.10 | −2.92*** |
| LISTAGE | 1.02 | 0.02 | 0.21 | 1.05 | 0.05 | 0.45 |
| CAPREQ | 1.16 | 0.15 | 0.32 | 1.13 | 0.12 | 0.27 |
| *CONTROLS:* | | | | | | |
| BIG4 | 1.15 | 0.14 | 0.74 | 1.11 | 0.10 | 0.56 |
| SIZE | 0.88 | −0.13 | −1.61 | 0.87 | −0.14 | −1.76* |
| BOARDIND | 1.40 | 0.34 | 0.51 | 1.55 | 0.44 | 0.67 |
| LEVERAGE | 1.17 | 0.16 | 1.77* | 1.18 | 0.17 | 1.79* |
| INDUSTRY | 1.92 | 0.65 | 3.94*** | 1.79 | 0.58 | 3.57*** |
| No. of Obs. | 1212 | | | 1212 | | |
| Likelihood Ratio | −613.64 | | | −628.15 | | |
| Pseudo R-square | 0.05 | | | 0.05 | | |

The table reports the results of logistic regressions for determinants of the propensity to delay financial flash reports during COVID-19. The dependent variable (COVIDLAG1) takes the value of 1 for firms that delayed reporting in 2020 but not in 2019 and zero otherwise. The dependent variable (COVIDLAG2) takes the value of one for firms with reporting days greater than 45 days and zero otherwise. ***, ** and * indicate significance at 0.01, 0.05, and 0.1 levels, respectively. All p-values are based on robust standard errors. The independent variables are defined in Table 1.

reporting, has implications for the economy in general, helping attract considerable capital from overseas investors. Financial reporting is essential for management communication, and its timeliness is as important as its content. Understanding the motivation for early and late reporting and the characteristics of early and late reporters, especially during periods of uncertainty, should contribute to investors' ability to evaluate the firm value and ultimately contribute to the allocative efficiency of the capital market.

However, the COVID-19 pandemic challenges firms' ability to meet the obligation of prompt reporting, resulting in the Japanese regulatory authority's relaxation of the 'within 45 days' rule. Notwithstanding this challenge, several Japanese firms with the same fiscal year ending 31 March 2020 satisfied the 45-day regulatory requirement for publishing financial flash reports. This study investigated the determinants of flash reporting timeliness within the framework of incentive-based and accounting complexity theories. We also examined the market reactions to the announcements of flash reports, distinguishing between early and late filers and between the waiver year (2020) and the non-waiver year (2021).

We hypothesise that firms with good dividends and earnings news are associated with a lower reporting lag. Our results support the hypothesis that firms with positive earnings news report timelier to signal managerial capital, reduce the uncertainty related to firm performance, and influence stock valuation. However, we find no evidence to support the dividend news hypothesis and the conjecture that management delays financial flash reports to reduce proprietary costs. The results support our hypothesis on the association between reporting lag and accounting complexity.

Our findings on good news incentives and accounting complexity have important implications for firms. Regarding good news incentives, the question is how much the stock market incorporates good/bad news into the share price. According to the efficient market hypothesis, it is irrelevant for firms to time the market by manipulating the timing of corporate announcements. In contrast, managers may think that their announcements change the current direction of their share prices. There are many cases where share prices sometimes fall with good news and rise with bad news.

Perhaps, regardless of the information content of corporate announcements, timely announcements would enhance the firm's reputation among investors, thereby lowering the cost of capital.

From a managerial perspective, accounting complexity relates specifically to the flow of information from a parent company to its subsidiaries. Our findings suggest a bottleneck problem exists between the parent firm and its subsidiaries regarding financial information for late reporting firms.

Our analysis of abnormal returns around flash reporting suggests that flash reports retained their value relevance attributes during the pandemic, notwithstanding the unprecedented difficulties experienced by firms during this period. However, this positive abnormal return is driven by early filers, as we report positive but insignificant abnormal returns for late filers. However, the difference in abnormal returns between late and early filers and between the waiver year (2020) and non-waiver year (2021) is insignificantly different from zero, suggesting a mooted impact of the pandemic and overall market reactions.

This study contributes to the literature and should interest firms, investors, and policymakers. By empirically highlighting the determinants of reporting lag, our results provide focal points for firms to ensure the timely delivery of critical financial information to the public and market participants. For investors, our research offers a way to interpret the recent reporting lag by empirically isolating the underlying reasons for the delay. Our findings also help regulatory bodies to understand the impact of the COVID-19 pandemic on business activities.

Although almost every country faces the same challenges and introduces emergency measures related to stock markets, a limitation of the study is that our empirical research is conducted within a single country. Therefore, it overlooks country-specific factors, which we recommend should receive the attention of future studies. Finally, we recommend further studies on the other implications of timely reporting during COVID-19. Such studies should focus on the consequences of the pandemic on restatements, the auditing quality, and the tone of financial reporting.

## Notes

1. https://www.jpx.co.jp/news/1023/nlsgeu000004rbjm-att/nlsgeu000004rbm7.pdf.
2. Note that the number of listed firms in the 1st section of the Exchange is approximately 2,000. (See the Japan Exchange Group Number of Listed Companies/Shares | Japan Exchange Group (jpx.co.jp)).
3. We broadly define this concept to include the complexities of accounting, operations, and auditing.
4. The earnings of firms listed in the 1st Section of the Exchange declined by about 70% year-on-year in the first quarter of 2020.
5. The results remain the same; we used the return on the TOPIX Index.
6. Armitage (1995) suggests that estimation periods can comprise 100 to 300 days for daily observations.
7. The code does not require that one-third of the directors are independent but that listed firms appoint two or more independent directors.
8. We evaluated this at the average value of the INVTA. Specifically, $1\% \times 1212 \times (-0.425) \times (-0.34) = 1.75$ days, where 1212, −0.425, and −0.34 are the number of observations, the average value of EARNSNEWS, and unstandardised coefficient on EARNSNEWS, respectively.
9. We evaluated this at the average value of the INVTA. Specifically, $1\% \times 1212 \times (0.118) \times (4.53) = 6.48$ days, where 1212, 0.118, and 4.53 are the number of observations, the average value of INVTA, and unstandardised coefficient on INVTA, respectively.
10. We evaluated this at the average value of the SIZE. Specifically, $1\% \times 1212 \times (11.72) \times (1.1/100) = 1.56$ days, where 1212, 11.72, and 1.11 are the number of observations, the average value of SIZE, and unstandardised coefficient on SIZE, respectively.
11. Firms reporting after the 45 days allowed during the non-COVID-19 years are classified as late filers.
12. The 'within 45 days' rule was not waived in 2021, notwithstanding the continued impact of the pandemic.
13. Similar results were obtained when we compared with the year 2019.
14. We thank the anonymous reviewer for bringing this to our attention.

## Disclosure statement

No potential conflict of interest was reported by the author(s).

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