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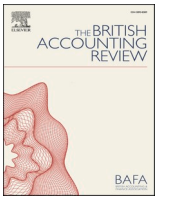

Validating implied cost of capital with realized returns by using alternative measures of cash-flow news

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ABSTRACT

We outline analytically that, when testing different implied cost of capital (ICC) estimates for validation by employing the Vuolteenaho (2002) framework, the cash-flow news in the validation framework should be defined in a way that considers the ICC model specific assumed sequence of future cash flows. This is based on market expectations, as proxied by analysts' forecasts. We then propose adjusting the cash-flow news proxies accordingly and implement these adjustments empirically. Consistent with the theoretical predications, the results from these tests show that ICC estimates are significantly positively related to realized returns. Informed by these findings, we employ the adjusted cash-flow news proxies in the validation framework and compare the correlation between adjusted and unadjusted for analysts' bias ICC estimates with realized returns. These tests show no difference in these correlations. This suggests that it is not the analysts' bias that weakens the validity of ICC estimates, as argued by prior literature. It is the proxies used in the validation framework that lead to the suggestion in removing analysts' optimism from ICC estimates. Overall, our proposed alternative framework not only unlocks the gate for extensive use of the existing ICC estimates but also enables future researchers to develop more reliable and meaningful ICC estimates.

1. Introduction

This paper offers a reliable and easy to calculate alternative proxy for cash-flow news, to be used when validating implied cost of capital (ICC) estimates. We provide evidence in favor of the construct validity of ICC estimates, when this proxy is used, eliminating the literature's concerns that has held back the wider application of the ICC method.

A fundamental issue in valuation is the estimation of cost of capital or systematic risk effects. Unlike *ex-post* cost of capital techniques, *ex-ante* techniques estimate the so-called ICC as the internal rate of return that equates current price to earnings forecasts and a long-term growth rate. This method has intuitive appeal and reverse engineers the residual income valuation model (RIM) and the abnormal earnings growth model (AEGM), using observable market prices and analysts' earnings forecasts. The ICC measures, if valid, represent an attractive option for empirical researchers addressing important research questions, and a number of studies apply ICC

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measures as a proxy for priced risk (e.g., Chen et al., 2011; Elzahar et al., 2015; Francis et al., 2004, 2005; Hail & Leuz, 2006; Mazzi et al., 2017; Mohanram & Rajgopal, 2009; Ortiz-Molina & Phillips, 2014).

A strand of the literature attempts to test the ICC validity by applying different research designs (i.e., Botosan & Plumlee, 2005; Easton & Monahan, 2005; Botosan et al., 2011; Larocque, 2013; Mohanram & Gode, 2013; Easton & Monahan, 2016). This literature concludes that, because "… empirical finance and accounting literature has largely failed to agree on a set of risk factors" (Easton & Monahan, 2016, p. 60) (hereafter EM 2016), the most appropriate validation method for ICC estimates is to test their relation with future realized returns (Easton & Monahan, 2005) (hereafter EM). This is now the standard approach (e.g., Larocque, 2013; Mohanram & Gode, 2013) and it is in this strand of the literature that this study is positioned in.

EM employ empirically a framework based on the return decomposition developed analytically by Vuolteenaho (2002) and propose evaluation of ICC estimates by their ability to explain realized returns, while controlling for cash-flow and discount-rate news. To model cash-flow news, EM rely on an autoregressive process that estimates return-on-equity from year two onwards and they apply the same cash-flow news proxy for all ICC proxies they test for validation. They do not find construct validity for ICC estimates for the entire cross-section of firms. Instead, some ICC measures are reliable only for specific sub-samples. Given this, researchers are skeptical about the use of ICC estimates (c.f., Penman, 2016).

In an attempt to solve the construct validity puzzle, more recent studies (e.g., Larocque, 2013; Mohanram & Gode, 2013) focus on the impact of analysts' bias, given the ICC estimates reliance on analysts' forecasts. Following labor-intensive computations that require extensive data, reduced sample size, and a "comprehensive model" (see Mohanram & Gode, 2013, p. 445), these studies propose techniques that account and correct for analysts' optimism bias. When they validate the resulting ICC by applying the EM approach, they report a significantly positive relation between ICC and future realized returns.[1]

In parallel to the developments in the accounting literature, recent finance literature emphasizes the importance of cash-flow vs. return news (e.g., Chen et al., 2013; Chen & Zhao, 2009; Khimich, 2017), conceptually and within the general implementation of the Vuolteenaho (2002) framework. Cash-flow news is a "proxy for changes in investors' expectations about future cash flows" (Patatoukas, 2014, p. 142). It is the dominant factor driving firm-level stock returns, whereas expected-return news is less important at the firm level (cf. Vuolteenaho, 2002). When implementing the Vuolteenaho (2002) framework, Khimich (2017), Da et al. (2014) and Da and Warachka (2009) propose revision in analysts' forecasts as a more accurate proxy of market expectations.

On reflection of these developments and particularly the specific model design about the expected future cash flows in each ICC measure, we revisit the EM approach for validating ICC proxies. We consider that each ICC model relies on a different sequence of future cash flows and none of the commonly used models makes use of an autoregressive approach for return-on-equity (which has been criticized recently also by the accounting literature (e.g., Penman, 2016; Penman & Yehuda, 2019). This reveals a potential inconsistency in the existing approach of testing ICC measures for validation.

We conjecture that the cash-flow news proxy in the regression of realized returns on expected returns, when testing ICC proxies for validation, has to capture market expectations about future cash flows, as these have been modelled in each individual ICC proxy tested for validation. The Botosan et al.'s (2011) design is consistent with this intuition, albeit only for validating the ICC approach developed by Botosan and Plumlee (2002).[2]

We advance the pertinent streams of the accounting literature in various ways. First, we outline how the cash-flow news proxy in the regression of realized returns on expected returns has to be aligned with the specification of the cash flows in each individual ICC approach tested for validation. Based on this, we develop easily implemented measures of cash-flow news proxies based on revisions in analysts' forecasts for each commonly used ICC method (i.e., earnings-to-price ratio, PEG, RIM (Claus & Thomas, 2001), and AEGM (Gode & Mohanram, 2003)). We show empirically that after controlling for such cash-flow news and return news, commonly used ICC estimates are positively related to realized returns and, in fact, the coefficient of the implied risk premia is not statistically different from the theoretical benchmark of one for three out of the four ICC approaches assessed in our main analyses. Thus, our results and the arguments and findings in the recent literature indicate that the EM's conclusion of invalid ICC measures is perhaps an overstatement. Thereby, we advance the ICC *validation* literature by introducing an alternative proxy for cash-flows news and show that ICC estimates are reliable proxies for expected returns when analysts' forecasts are used. Analysts' forecasts are readily available and, thus, easy to use, depending on the research question and sample.

Subsequently, we account and correct for analysts' optimism bias when explicitly validating ICC estimates (e.g., Larocque, 2013; Mohanram & Gode, 2013) and explore any resulting inferences arising from applying our proposed refinement. We show that adjusting for analysts' forecast bias in the ICC measures themselves effectively reduces the importance of cash-flow news in the validation framework. Consequently, the potential misspecification in the existing validation approach becomes less important when the ICC measures tested for validation have accounted for analysts' bias. In support of that, when using our proposed method, adjusted for analysts' optimism ICC estimates are not superior to their unadjusted counterparts. Thus, we offer a subsequent methodological

[1] Studies also propose to use mechanical forecasting models as a substitute to analysts' forecasts (Hou et al., 2012; Li & Mohanram, 2014). However, there is mixed evidence as to how mechanical forecasts are advantageous (Feng, 2014; Gerakos & Gramacy, 2012) over analysts' forecasts. Mechanical forecasts rely on sophistical models that gather earnings forecasts based on historical data only, while the literature shows that analysts' forward-looking data contains useful information (e.g., O'Brien, 1988).

[2] They rely on analysts' forecasts of dividends and target prices, while the cash-flow news specification is based on firm-level changes in target prices as provided by ValueLine. However, ValueLine is limited to a small US firms sub-set and cannot be used for international samples. It covers c.1,700 US stocks while I/B/E/S covers c.22,000 stocks across 100 countries. This might explain why EM's (2005) approach is the most commonly used in empirical research (see, e.g., Nekrasov & Ogneva, 2011; Mohanram & Gode, 2013).

contribution to the ICC *estimation* literature: the labor-intensive computations that adjust for analysts' optimism are not necessary.[3] Overall, our methodological contributions provide construct validity for ICC estimates, unlocking the gate for the use of ICC to answer important research questions.

The rest of this paper is structured as follows. Section 2 first outlines the relation between expected and realized returns and subsequently discusses pertinent ICC research and recent finance literature on revision of analysts' forecasts and cash-flow news. Section 3 elaborates on the ICC model specific cash-flow dynamics and outlines model-specific cash-flow news validation proxies. Section 4 empirically validates common ICC estimates. Section 5 discusses our results and compares them with prior research that adjusts for analysts' bias. Section 6 concludes.

2. The relation between expected and realized returns and the role of cash-flow news

In this section, first, we outline the key features of the Vuolteenaho (2002) return decomposition framework. Then, we explain how EM have operationalised this for the purposes of validating ICC measures and the conclusions they reach. Subsequently, we reflect on evidence in more recent finance literature that provides insights on the measurement of cash-flow news when operationalizing the return decomposition framework more broadly.

2.1. Decomposition of realized returns and EM's empirical proxies

The seminal work by Campbell and Shiller (1988a, 1988b) and Campbell (1991) decomposes unexpected stock returns into two components: Cash-Flow and Discount-Rate news. Vuolteenaho (2002) advances this work by tailoring the framework to the firm level. Specifically, Vuolteenaho (2002) reformulates the decomposition utilizing the clean surplus relation and ROE as a basic cash-flow variable (Khimich, 2017). As such, Vuolteenaho (2002) shows that firm-specific realized returns (RET) can be decomposed into three components: (i) the expected return (RP), (ii) the changes in expectations about the sequence of future cash-flow news ($CNEWS$), and (iii) changes in expectations about future discount rates (discount-rate news, $DNEWS$). This is expressed as follows:

$$RET_{i,t+1} = RP_{i,t} + CNEWS_{i,t+1} \quad DNEWS_{i,t+1}, \tag{1}$$

where RP is formed at the end of year t, and both types of news reflect changes in expectations occurring during year $t+1$, i.e. the return interval. In theory, the coefficient for RP should be equal to one (Botosan et al., 2011). With regard to measurement of cash-flow news in particular, Vuolteenaho's (2002) decomposition relies on the return-on-equity:

$$CNEWS_{i,t+1} = \Delta E_{t+1} \sum_{t=1}^{\infty} p^{\tau \ 1} \left[ROE_{i,t+\tau}\right], \tag{2}$$

Where $\Delta E_{t+1}[.]$ equals $E_{t+1}[.] \quad E_t[.]$, p is a number slightly less than 1, and ROE is the natural log of 1 plus the accounting rate of return-on-equity.

Informed by this, EM employ empirically the Vuolteenaho (2002) framework (Eq. (1)) as a means of validating ICC estimates that have been computed based on different models. For the empirical implementation, EM define cash-flow news as follows[4]

$$CNEWS_{i,t+1} = \ ROE_{i,t} \quad FROE_{i,t,t}\big) + \ FROE_{i,t+1,t+1} \quad FROE_{i,t,t+1}\big) + \frac{p}{1 \quad p * w_t} * \ FROE_{i,t+1,t+2} \quad FROE_{i,t,t+2}\big), \tag{3}$$

where ROE denotes return-on-equity, $FROE_{i,j,k}$ the forecasted return-on-equity for fiscal year k, which is based on the forecasts made in December of year j, and w_t is the expected persistence of ROE as of time t. This definition of cash-flow news assumes that the ROE follows a first-order autoregressive process after year $t+1$.[5] However, as explained by Khimich, 2017: p., 34), "empirical implementation of the return decomposition requires an expectation model of future cash flows and discount rates". This suggests a significant difference between what is outlined in Vuolteenaho (2002) and how the framework has been applied for validating ICC measures. Vuolteenaho's (2002) theoretical model relies on the change in the accounting return-on-equity over an infinite horizon, whereas EM proxy this by the expected forecast return-on-equity for a maximum of two years in combination with a first order auto-regression process (VAR I). This implies a truncation error in the cash-flow news definition, unless the VAR captures the "true" behavior of the future return-on-equity (cf. Khimich, 2017; Penman & Yehuda, 2019). The appropriateness of EMs assumption, thus, depends on whether the VAR (I) reasonably captures market expectations.

[3] We do not argue that analysts' estimates are not biased or that an adjustment in other contexts relating to analysts' forecasts is not needed. Our discussion and corresponding contribution relate to the specific models applied to adjust for analysts' optimism in the context of ICC estimates.

[4] Since we follow the same approach as in EM to measure DNEWS, the exact formula is presented in section 4.1 below.

[5] The parameter w_t is estimated using the following pooled cross-sectional and time-series regression: $ROE_{i,\tau+1} = w_{0,t} + w_t * ROE_{i,\tau}$, where τ is a number between t and t-9 and the sample consists of all firm-year observations in the same Fama and French (1992; 1997) industry with all necessary data in years t-9 through t. EM (2005) use an industry-specific persistence factor to account for the impact of accounting methods, competition, and risk across industries. The capitalization factor p varies with the price-to-dividend ratio, and EM estimate $p = 0.988$ for non-dividend paying stocks, $p = 0.957$, $p = 0.921$, $p = 0.927$, and $p = 0.924$ for the fourth, third, second, and first quantiles of price-to-dividend ratio for dividend paying stocks, respectively.

2.2. EM's results

As discussed earlier, in theory, the coefficient for *RP* should be equal to one. However, as Botosan et al. (2011) explain, in empirical research, when *RP* captures cross-sectional variation in expected returns (i.e., the coefficient of *RP* is positive and significant), the *RP* proxy could be considered a valid one. When regressing realized returns on expected returns as outlined earlier, EM find that none of the assessed implied risk premia reports a positive and statistically significant coefficient.[6] In fact, the coefficients of all but one of the implied risk premia ICC proxies tested are negative (and in fact some are statistically significant). The exception is the coefficient for the ICC based on the RIM (Claus & Thomas, 2001) approach, which is positive but statistically insignificant. The coefficients of CNEWS and DNEWS are significantly positive, in line with expectations. This evidence seems somewhat puzzling given the large empirical evidence indicating that earnings, analysts' forecasts of earnings or book values, at least partly, explain returns (e.g. Penman, 2016). In further analysis, EM show that, for sub-samples consisting of the observations with the lowest absolute ROE forecast error, some ICC estimates show a positive significant coefficient. Based on these findings, EM caution against the use of ICC estimates as a reliable proxy for expected returns. However, for sub-samples, they claim that some of the ICC estimates are reliable.[7]

More recent research in accounting applies the EM framework when assessing ICC estimates and shows some improvements when adjusting for predictable analysts' bias (e.g., Larocque, 2013; Mohanram & Gode, 2013). In light of the proposed refinement when validating ICC proxies computed with analysts' forecasts below, in section 5, we revisit this literature and explore whether the proposed adjustments for predictable analysts' bias are necessary.

2.3. Recent literature on cash-flow news

The return decomposition framework is widely accepted and applied in asset-pricing research (for reviews see, e.g., Campbell et al. (1997) and Cochrane (2001)). In fact, after the publication of the EM study, finance literature provides some new insights on to which we draw when revisiting the implementation of the Vuolteenaho (2002) framework as a means of validating ICC proxies below.

Chen and Zhao (2009) and Maio and Philip (2015), *inter alia*, show that the return decomposition framework is highly sensitive to the empirical proxies chosen and potential model misspecifications, often leading to opposing conclusions. Chen and Zhao (2009, p. 5214) state that "[w]hile model misspecification is always a potential problem any empirical model faces, it is likely to be more damaging for the return decomposition approach." Additionally, Khimich (2017) and Da et al. (2014) (and see also Da & Warachka, 2009) show that revision in analysts' forecasts performs well and, importantly, outperforms the commonly used VAR to model market expectations. More specifically, measuring cash-flow news by analysts' forecast revisions and applying a simple reversal strategy, Da et al. (2014) generate significant risk-adjusted returns. Thus, they conclude revisions in analysts' forecasts provide a good proxy for market expectations. Khimich (2017) comparatively assesses various cash-flow news proxies, providing evidence that the revision in analysts' forecasts outperforms the VAR approach in measuring cash-flow news i.e., it is a better proxy for markets' expected cash-flows than the VAR method (because of the advantage of forward-looking analysts' data than the VAR implied future cash-flow pattern). In line with this, Penman (2016) and Penman and Yehuda (2019) also criticize the autoregressive process assumption.

These findings and statements are particularly relevant in the context of the Vuolteenaho (2002) framework as a means of validating ICC proxies because Vuolteenaho (2002) finds that cash-flow news is the dominant factor driving stock returns, whereas expected return news is less important. These findings raise the opportunity for one to reconsider the commonly used validation approach that assumes a VAR process after year $t+1$ to be a reasonable proxy for market expectations when assessing ICC.[8] It is an open question whether a different proxy for cash-flow news would result in different conclusions about the validity of ICCs. As such, we revisit the operationalization of the existing validation framework accordingly.

3. Introducing model consistent cash-flow news proxies for the validation of ICC estimates

3.1. Return decomposition and cash-flow news

Botosan et al. (2011, section 7) outline that, by construction, all ICC models are a function of current price ($P_{i,t}$) and series of

[6] While return news is expected to have a negative impact on the returns (i.e., when empirically implementing Eq. (1)), EM multiply the discount rate news by minus one to expect a positive sign in their multivariate analyses.

[7] It is commonly acknowledged that each of the regressors in Equation (1) above contain a measurement error. EM (2005: 505) explain that this "implies the bias in a particular regression coefficient is a complex function of the measurement errors in all of the regressors (e.g., Rao, 1973)". To account for this issue, in addition to their main analysis, EM use a refinement of the approach discussed by Garber and Klepper (1980) and Barth (1991) and estimate measurement error variances. This additional approach isolates the portion of the bias in the coefficient of the $RP_{i,t}$ that is solely attributable to the measurement error in $RP_{i,t}$. This is discussed in detail in Section II (pages 506 and 507) in EM (2005). We refrain from repeating the equations and discussion for brevity. However, we have also conducted these additional tests and we report the modified noise variables, along with the results of our main analysis.

[8] Because EM use revisions in analysts' forecasts regarding years one and two, Khimich (2017) classifies them as using the RAF approach. However, these analysts' forecasts are used only as a starting point for then implementing VAR. So, technically EM employ a combination of the RAF and VAR approaches. This is why Khimich (2017, p. 35) explains that "the RAF method outlined in this paper computes the revision in analysts' forecasts after time $t = 2$ slightly differently than Easton and Monahan (2005)." Given that a major part of the firm value is captured after year two, the VAR is the dominant factor within EM's approach.

expected future cash flows ($CF_{i,t}$):

$$RP_{i,t} = f\ P_{i,t}, CF_{i,t}). \tag{4}$$

Similarly, Khimich (2017) shows that current stock price is a function of cash flows and expected return ($RP_{i,t}$). As outlined earlier, the cash-flow news represents the change in the expected future cash-flows sequence between t (when the expected returns are estimated) and $t+1$ (when the realized returns are observed). Thus, the cash-flow news proxy captures changes in the expected cash flows:

$$CNEWS_{i,t+1} = \ CF_{i,t+1} \quad CF_{i,t}) \tag{5}$$

Moreover, the return news (i.e., discount-rate news) is the expected risk premium in period $t+1$ less the expected risk premium at t. Alternatively, it is a function of the price and expected cash flows sequence at $t \ + \ 1$, less a function of the price and expected cash flows sequence at t (EM and, Mohanram & Gode, 2013). This is exactly the same function as for the estimation of the risk premium in Eq. (4). Hence discount-rate news can be expressed as:

$$DNEWS_{i,t+1} = RP_{i,t+1} \quad RP_{i,t} = f\ P_{i,t+1}, CF_{i,t+1}) \quad f\ P_{i,t}, CF_{i,t}). \tag{6}$$

Combining these elaborations, we propose that Eq. (1) can be reformulated as follows:

$$RET_{i,t+1} = f\ P_{i,t}, CF_{i,t}) + \ CF_{i,t+1} \quad CF_{i,t}) + [\ \ \{f\ P_{i,t+1}, CF_{i,t+1}) \quad f\ P_{i,t}, CF_{i,t})\}]. \tag{7}$$

From this, it is evident that the expected cash flows in Eq. (7) determine all three independent variables in the return decomposition.

Then we consider that, theoretically and intuitively, there is only one sequence of cash flows that forms market expectations for each company. However, each of the commonly used firm-level ICC models assume a specific sequence of future expected cash flows and they differentiate primarily in the way they use expected cash flows in the short- and the long-run (e.g. Ohlson, 2005; Ohlson & Gao, 2006) (for a discussion on how they approximate cash flows, see section 3.2 below).

Against this backdrop, we argue that a clear-cut validation of any specific ICC estimate (first part on the right-hand side of Eq. (7)) goes along with an evaluation of how the specific ICC model employed for the computation of this proxy uses future (expected) cash flows in a truncated horizon. Thus, from an empirical point of view, given the truncated horizon in which competing ICC models are employed, model-specific cash-flow news ensures one specification of future cash flows in the (different parts of the) return decomposition. This allows for unambiguous empirical evidence as to how valid any specification of future cash flows and, in turn, the corresponding ICC model, is. This is also in line with EM (2005, p. 512) who assert that "expectations and changes in expectations are inextricably linked …; hence, cn_{it+1} [cash-flow news] reflects revisions in the expectation underlying er_{it+1} [the expected returns]".

This leads us to the central observation that the specification of the cash-flow news proxy EM use could have influenced the empirical implementation of the return decomposition, resulting in the insignificant results EM obtain for their full sample (see discussion by Khimich (2017) and Chen and Zhao (2009) on the sensitivity of the return decomposition to the inputs used). Respectively, their findings of significant positive coefficients when they build sub-samples depending on the extent of the absolute forecast error could have also been influenced.[9]

Based on this and the recent evidence on the superiority of using revisions in analysts' forecasts method for the estimation of cash-flow news discussed above, in the next section, we elaborate on how such model-specific cash-flow news can be developed empirically for testing the validity of different ICC models with the return decomposition as an alternative.

3.2. Outline and empirical implementation of model-specific cash-flow news proxies

The ICC represents the internal rate of return when equating the current stock price to future expected cash flows. To derive those, the ICC literature employs accounting-based valuation models, mainly versions of the residual income model (RIM) and the abnormal earnings growth model (AEGM). Common across the commonly used models is the fact that they rely on the current price, forecast of (analysts') earnings within a finite horizon and a long-term growth assumption (for detailed review see Echterling et al. (2015)).

Given the wide range of different ICC models, we test the validity of the most commonly used ICC estimates (i.e., earnings-to-price ratio ($RP^{E/P}$), PEG (RP^{PEG}), RIM (Claus & Thomas, 2001) (RP^{RIM}), and AEGM (Gode & Mohanram, 2003) (RP^{AEGM})).[10], [11] Our proposition for model-specific cash-flow news proxies when validating ICC estimates, with the return decomposition framework, is

[9] This criticism also reveals that the approach followed by Botosan et al. (2011) does not provide clear-cut evidence since they (i) use an inconsistent definition of cash-flow news for commonly used ICC measures (except that of Botosan and Plumlee (2002), which also relies on target prices), and (ii) apply a discount-rate news proxy that relies on the change in market beta and the risk-free rate (for a discussion of the discount rate news see also EM (2016) and our discussion in Appendix B.4).

[10] Following Mohanram and Gode (2013), we use the implied risk premia for all ICC approaches, i.e., we subtract the prevailing risk-free rate.

[11] These models are reviewed in Appendix A.1. The Gebhardt et al. (2001) approach uses the RIM, however, it mainly deviates from Claus and Thomas (2001) in the terminal value specification. Easton (2004) also develops AEGM variations whose most important variation is the PEG ratio (which we analyse). Overall, we do not include various variations of RIM and AEGM for brevity and clarity purposes. Our proposed cash-flow news, though, could also be applied in these methods.

informed by previous research that outlines the dynamics of the various accounting-based valuation models (e.g. Ohlson, 2005; Ohlson & Gao, 2006). In fact, we take into consideration the fact that the long-term cash-flow news matter and, in fact, theoretically should matter more than the short-term cash-flow news (Ohlson, 2005; Ohlson & Gao, 2006). This is because the terminal value in accounting-based valuation models captures the largest proportion of the equity value (Penman & Sougiannis, 1998).

Based on this prior work, Panel A of Table 1 summarizes the main assumptions of each of the four ICC models we consider in this study. This shows that all models rely on earnings forecasts as a proxy for market expected future cash flows. This is not against the return decomposition framework by Vuoltneenaho (2002, p. 235) who states that "[w]hether one chooses to think about infinite-horizon cash-flow fundamentals in terms of dividend growth or ROE is a matter of taste, however." The use of earnings forecasts is mainly driven by data constraints given that earnings and earnings development is in the focus of analysts and thus earnings forecasts are most widely available proxies for future earnings (e.g. Claus & Thomas, 2001). While the reliance on analysts' earnings forecasts is pertinent for all models, the models differ in terms of the number of periods for which earnings forecasts are required and used and the growth assumption beyond the specific forecast horizon (i.e. the long-term growth assumption, which is also affected by the model choice e.g. RIM vs. AEGM).

As discussed earlier, EM's cash-flow news proxy is based on the forecasted return-on-equity for the short-term, in combination with a VAR (I) process for the long-term cash-flow news. Thus, they consider short-term and long-term portions of the cash-flow news proxy.[12] However, they apply the same cash-flow news proxy for all ICC measures they test for validation, given that they assume that "cash flow news is the component of realized returns corresponding to the change in investors' expectation about future cash flows" (EM, p. 511) and thus link it to the realized returns only, without reference to the ICC proxy tested for validation. While the use of one specific cash-flow news assumption is theoretically defendable, it is not in line with how the ICC models mainly differ in the way they approximate future cash flows (theoretically as well as when empirically implemented - see above). For the discount-rate news, they importantly link it to both, expected and realized returns, given that they use the change in the estimate from each ICC estimate tested for validation. Specifying cash-flow news without any reference to the ICC model brings difficulties in interpreting the empirical outcome. More specifically, when using different specifications of cash flows across the individual components (e.g., ICC, CNEWS, and DNEWS), it is not clear which of the separate parts absorbs the variance in cross-sectional analyses.

As an alternative, we outline an empirical specification of cash-flow news dependent on the sequence of cash flows applied within each respective model, resulting in different cash-flow news proxies for each model. In line with EM and Botosan et al. (2011), our proxies include a short- and a long-term proportion, capturing changes in the expected cash flows. The short-term cash-flow news is measured as the percentage change in earnings surprises during the realized return period (CF_{ST}). This is the same across all models, given that all use and thus all are affected by short-term changes. The empirical specification for the long-term cash-flow news is based on the percentage change in analysts' forecasts considering the model specific period for which forecasts are used (see Panel A of Table 1) (CF_{LT} *level*).,[13] [14] To account for the growth component inherently in three of the four models tested for validation, the percentage change in expected analysts' growth is further included (CF_{LT} *growth*).[15] How this reflects the assumptions of each model is presented in Panel B of Table 1. The formulae for the computations are provided in Appendix A.2.

4. Data, variable computation and empirical results

4.1. Data and variables

Our sample selection process is in line with Easton et al. (2002), Easton (2004), and Mohanram and Gode (2013) and summarized in Table 2. We begin by including all firms covered by I/B/E/S traded on NASDAQ and NYSE for the years 1981–2014. Forecasts of earnings and long-term growth rates are collected from I/B/E/S, data on fundamentals is drawn from Worldscope, and the risk-free rate is collected from Datastream. In line with Easton et al. (2002) and Easton (2004), we assume that the dividends in period t

[12] Botosan et al. (2011) also consider short-term and long-term portions of the cash-flow news proxy. However, they follow an alternative approach. They measure short-term cash-flow news as the earnings surprise during the realized return period and the revision in analysts' forecasts of target prices during the realized return period serves as the long-term proxy. The former is measured as the earnings surprise scaled by price and the latter by the change in target price scaled by price.

[13] As a robustness check, we have also run the tests by including changes in earnings in period $t = 2$ and $t = 3$ in the cash-flow news proxy for the Claus and Thomas (2001) model. The results are virtually identical in relation to the ICC. However, we also observe that there is an almost perfect correlation between changes in expected earnings in periods $t = 2$, $t = 3$, and $t = 4$, caused by using the I/B/E/S long-term growth rate. This then results in very high levels of multicollinearity in the model. This does not allow for reliable inferences regarding the coefficients of the CF news proxies.

[14] Given how this proxy is defined, when empirically implemented, there may be some observations with negative denominators. However, these would apply only for the first forecast period available. This may create some noise in the interpretation of the overall results. As such, as a further robustness check, we delve in our data and identify 1,001 firm-year observations (out of 43.465 i.e., 2.3% of our sample) with a negative denominator. To explore any influence these may have on the interpretation of the findings presented in the paper, we repeat the analysis we present in Panel A of Table 5 by excluding these observations. The results of this analysis reveal that our results and conclusions from the study remain the same.

[15] We note that the RIM relies on residual earnings and their growth, while the AEGM relies on earnings and their growth. Residual earnings also rely on earnings, since they are calculated as the earnings less a charge (cost of capital) at the beginning-of-year book value. We therefore do not differentiate the cash-flow news proxy between RIM and AEGM.

Table 1
Model implied cash flow dynamic and outline of cash-flow news specification for different ICC measures.

Panel A: Model implied cash flow dynamic

RP metric	Cash-flow approximation	Cash-flow horizon used	(model implied) growth assumption beyond cash-flow horizon
$RP^{E/P}$	Earnings	Two	No
RP^{PEG}	Earnings	Two	Yes
RP^{RIM}	Earnings	Four	Yes
RP^{AEGM}	Earnings	Two	Yes

Panel B: Derived cash-flow news proxies

RP metric	CF_{ST}	CF_{LT}	
		level	growth
$RP^{E/P}$	ΔE_1	ΔE_2	
RP^{PEG}	ΔE_1	ΔE_2	ΔgE_2
RP^{RIM}	ΔE_1	ΔE_4	ΔgE_4
RP^{AEGM}	ΔE_1	ΔE_2	ΔgE_2

Panel A of this table shows the model implied dynamic for the four commonly used ICC models. It outlines the main cash-flow approximator, the forecast horizon applied by ICC research and whether a growth component is used after the forecast horizon, i.e. whether the models rely on a growth component. Panel B of this table summarizes the corresponding cash-flow news proxies derived from the model implied dynamics. $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM} stand for the four implied risk premia models under investigation (see Appendix B for more details). E_t stands for the earnings forecasts for period t and g indicates growth. Δ indicates change. CF_{ST} represents the short-term cash flows and CF_{LT} represents the long-term cash flows.

equal the dividends in the current period (i.e., $DPS_t = DPS_0$ for $t = 1, \quad 2, \quad 3, \quad 4$ (so data for years up to 2018 are needed)), and we proxy the risk-free rate by the ten-year US T-Bill rates. This sample selection process results in a total sample of 60,244 firm-year observations. It then decreases to 43,465 observations due to data absence to calculate all proxies i.e., the 12-month buy-and-hold returns or realized returns, cash-flow news, and return news. The resulting sample is comparable to previous research when focusing on the overlapping years (e.g., EM).

The realized returns (RET_1) is the buy-and-hold returns for 12 months following the calculation of the ICC proxies. We subtract the risk-free rate to be consistent with the ICC estimates. In accordance with EM, we transform all proxies to be continuously compounded i.e., we take the natural log of one plus for all proxies.

The discount-rate news (DNEWS) is defined as in EM: $DNEWS_{i,t+1} = \frac{p}{(1\ p)} * (RP_{i,t+1} \quad RP_{i,t})$ i.e., the discount-rate news is given by the change in the (implied) expected return proxy. p is a capitalization factor that allows the sensitivity of realized returns to a given change in the discount rate to vary across stocks, on the basis of variation in expected growth. This specification implies that the return news proxy varies across the expected return proxies (ICC models) and embeds the assumption that changes in the discount rate are permanent. Hence, they equal the change in the estimate of the expected return over the realized return interval.

4.2. Descriptive statistics

Table 3 provides descriptive statistics. Panel A focuses on the implied risk premia estimates and realized returns. The mean (median) implied risk premia estimates lie above (below) all the realized returns (except for RP^{AEGM}). This is because of the much higher standard deviation of the realized returns (0.4152) than that of all the implied risk premia (maximum of 0.0415). The distributions of realized returns and implied risk premia estimates are comparable to those of EM.[16] Panel B reports descriptive statistics for the various cash-flow and discount-rate news proxies. In line with EM, the median values are consistently positive, while for some approaches (in our case 2/4) the mean values are negative.[17]

Now we turn to the cash-flow news proxies. First, the mean and median values for the cash-flow news proxy of EM (i.e., Eq. (3), $CNEWS^{EM}$) are negative (mean 0.0581; median 0.0212). Negative values are also observable across our developed cash-flow news proxies. However, the absolute values of our proxies are, on average, higher than those of EM, particularly observable for the combined short- and long-term news measure. This is plausible given that both short- and long-term news are added and that the long-term news is multiplied by $\frac{p}{(1\ p)}$.[18] A negative value for the cash-flow news indicates that cash flows are downwardly adjusted during the return

[16] A comparison to a more recent paper, such as that of Mohanram and Gode (2013), is not possible given that we report log values while they report untransformed values. However, untabulated results of the untransformed values show a comparable distribution of our sample to that of Mohanram and Gode (2013).

[17] While EM transform the discount rate news by minus one only for the multivariate analysis, we do this for our whole empirical analysis, including the descriptive statistics and correlations.

[18] Untabulated results show that the mean and median values of our proposed cash-flow news proxies are significantly different from those of the EM proxy.

Table 2
Sample selection process.

Sample selection filters (mostly similar to Easton et al., 2002; Easton, 2004; Mohanram & Gode, 2013)	# of firm-years
I/B/E/S-covered firms from 1981 to 2014 with one- and two-year forecasts of earnings and long-term growth rate or one- to five-[a] year-ahead forecasts of earnings at the last Thursday of December in each year (median values) (I/B/E/S codes: EPSXMD stands for the earnings forecasts, where X is 1, 2, 3, 4, 5; LTMD stands for the long-term growth rate forecast)	85,255
Less observations with no common shareholder equity (Worldscope code: WC03501) or number of common shares outstanding available (WC05301)	5,508
Less observations with no data for current price (I/B/E/S: IBP), earnings (WC05202), or dividends (WC05110)	138
Less observations with negative book value of equity	1,486
Less observations with negative two-year-ahead forecast of earnings	6,058
Less observations with three-year-ahead forecasts smaller than two-year-ahead forecasts of earnings	3,840
Less observations at the top and bottom 1% of the price-to-book ratio and the earnings-to-book ratio as well as the top 1% of the two-year-ahead earnings-to-price ratio	2,310
Less observations with a non-converging algorithm for firm-level ICC estimates or estimates smaller than zero or greater than 60%	4,075
Less observations with $P_0 < 0.5$ or $P_0 > 500$	1,596
Sample for which all ICC approaches can be estimated	*60,244*
Less observations for which either one-year ahead buy-and-hold return (Datastream: RI), cash-flow news proxy, or return news proxy cannot be calculated	16,185
Less observations at the top 1% of the one-year ahead buy-and-hold return	594
Sample used in the main analysis	*43,465*

[a] Since we follow Easton et al. (2002) and Easton (2004), we do not use the forecasts for the year ending just a few days after the collection of the forecasts. Consequently, for our analysis we apply a maximum of four years of forecasts.

interval. This is, at least to some extent, attributable to analysts' optimism (see also the discussion in EM). Overall, a value different from zero for both types of news highlights the importance of considering cash-flow and return news when assessing the relation between expected and realized returns.

4.3. Correlations

Table 4 provides correlations among the various risk premia metrics as well as the return and the cash-flow news. Pearson (Spearman) correlation coefficients are reported above (below) the diagonal (the following discussion focuses on Pearson correlations). All of the implied risk premia metrics are highly correlated (ranging between 0.50 and 0.86). This is not surprising given that all rely on the same idea i.e., equating the current price with future expected cash flows (cf. Botosan et al., 2011). In comparison, none of the implied risk premia estimates show a high correlation with realized returns (largest value for $RP^{E/P}$ is 0.07), which is in line with prior research (e.g., Botosan et al., 2011; Mohanram & Gode, 2013).

Additionally, and in line with expectations and prior research (e.g., EM), all the discount-rate news and cash-flow news proxies are highly correlated with the realized returns (correlations range between 0.20 for $DNEWS^{AEGM}_{i,t+1}$ and 0.38 for $DNEWS^{E/P}_{i,t+1}$ and between 0.18 for $CNEWS^{EM}_{i,t+1}$ and 0.43 for $CNEWS^{E/P}_{i,t+1}$). Moreover, the correlations among the cash-flow news proxies we develop with realized returns (ranging between 0.24 and 0.43) are significantly higher (based on unreported results) compared to those of EM (0.18). This suggests that our cash-flow news proxies outperform those of EM.

Considering the correlation between both types of news, Table 4 shows that EM's cash-flow news are not correlated with the return news (correlations range between −0.01 for $DNEWS^{AEGM}_{i,t+1}$ and −0.07 for $DNEWS^{RIM}_{i,t+1}$). In comparison, our developed cash-flow news proxies are highly correlated with the return news (correlations range between −0.23 ($DNEWS^{E/P}_{i,t+1}$) and −0.72 ($DNEWS^{AEGM}_{i,t+1}$)). While this discussion focuses on the combined measure, the same applies for the separate components.

This observation raises two questions. First, should we expect such (high) negative correlation between the cash-flow and return news? Second, would this cause any multicollinearity problems when conducting our empirical analysis?

Regarding the first, recall what both news types represent. Cash-flow news is changes in expectations of the cash-flows sequence, whereas return news is expected returns changes, between t and t+1. Based on Eq. (6), the return news change depends on two variables: the price change and the change in the expected future cash flows sequence between t and t+1. Consequently, changes in the expected cash flows sequence are used in both types of news. Thus, a high (negative) correlation between the two news components is expected. Similarly, Vuolteenaho (2002) reports correlations between the return and cash-flow news of about 41% to 47%. Accordingly, the correlation of the EM proxy, which is close to zero, is difficult to explain and defend. Regarding the second question, the correlation coefficient values are well below the |0.8| threshold, thus, do not indicate multicollinearity (Gujarati, 2003, p. 359). We also address potential multicollinearity when discussing our multivariate results.

4.4. Main analysis

We employ two different regressions. First, and in line with Botosan et al. (2011), we include separate cash-flow news proxies:

Table 3
Descriptive statistics.

Panel A: Expected returns and realized returns.

	N	Mean	Median	SD	5th	25th	75th	95th
$RP^{E/P}$	43,465	0.0365	0.0322	0.0382	0.0170	0.0121	0.0840	0.1050
RP^{PEG}	43,465	0.0537	0.0494	0.0415	0.0063	0.0282	0.0740	0.1276
RP^{RIM}	43,465	0.0460	0.0415	0.0304	0.0066	0.0275	0.0598	0.1000
RP^{AEGM}	43,465	0.0693	0.0622	0.0397	0.0201	0.0447	0.0859	0.1426
RET_1	43,465	0.0154	0.0614	0.4152	0.7117	0.1648	0.2594	0.5886

Panel B: Cash-flow and return news proxies.

N = 43,465	Mean	Median	SD	5th	25th	75th	95th
$CNEWS^{EM}$	0.0581	0.0212	0.5601	0.3522	0.0992	0.0222	0.1473
($ROE_{i,t}$ $FROE_{i,t,t}$)	0.0319	0.0010	0.2119	0.1332	0.0172	0.0045	0.0376
($FROE_{i,t+1,t+1}$ $FROE_{i,t,t+1}$)	0.0092	0.0037	0.0608	0.0862	0.0243	0.0078	0.0469
($FROE_{i,t+1,t+2}$ $FROE_{i,t,t+2}$)	0.0107	0.0050	0.0655	0.0959	0.0274	0.0086	0.0516
$CNEWS^{E/P}_{i,t+1}$	4.5052	0.6672	22.4990	41.4409	4.9507	1.2067	16.1988
ST	0.1339	0.0305	0.4860	0.9675	0.2162	0.0588	0.3302
LT-level	0.1010	0.0370	0.3660	0.7334	0.2029	0.0584	0.3049
$CNEWS^{PEG,AEGM}_{i,t+1}$	3.7797	0.9718	20.6881	38.7035	6.1769	2.1290	19.1621
ST	0.1339	0.0305	0.4860	0.9675	0.2162	0.0588	0.3302
LT-level	0.1010	0.0370	0.3660	0.7334	0.2029	0.0584	0.3049
LT-growth	0.0476	0.0163	0.7166	1.1705	0.3285	0.2177	1.0941
$CNEWS^{RIM}_{i,t+1}$	2.7325	0.6544	18.1634	32.3352	5.0461	2.0026	18.3253
ST	0.1339	0.0305	0.4860	0.9675	0.2162	0.0588	0.3302
LT-level	0.0784	0.0351	0.3726	0.7077	0.2063	0.0859	0.3899
LT-growth	0.0344	0.0000	0.5318	0.8402	0.1823	0.1086	0.7621
$DNEWS^{E/P}_{i,t+1}$	0.1495	0.0067	1.9383	2.6445	0.3100	0.2586	1.8759
$DNEWS^{PEG}_{i,t+1}$	0.0100	0.0256	1.9808	2.7061	0.3107	0.3510	2.7055
$DNEWS^{RIM}_{i,t+1}$	0.0462	0.0184	1.4585	.21142	0.2468	0.2517	1.8270
$DNEWS^{AEGM}_{i,t+1}$	0.0171	0.0294	2.0066	2.7145	0.3282	0.3761	2.7374

Panel A of this table provides descriptive statistics for the four implied risk premia estimates and realized returns. The four implied risk premia estimates, i.e., implied cost of capital less the prevailing risk-free rate, $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, are estimated as outlined in Appendix B. RET_1 is the one-year-ahead buy-and-hold return following the calculation of the implied risk premia, less the risk-free rate. Panel B of this table provides descriptive statistics for the cash flow as well as discount-rate news proxies. $CNEWS^{EM}$ is the cash-flow news as outlined in Equation (2) (according to Easton & Monahan, 2005). $CNEWS^{E/P}_{i,t+1}$ ($CNEWS^{PEG,AEGM}_{i,t+1}$) is the approach-specific cash-flow news proxy for the $RP^{E/P}$ (RP^{PEG} and RP^{AEGM}). $CNEWS^{RIM}_{i,t+1}$ is the approach-specific cash-flow news proxy for the RP^{RIM}. ST, LT-level, and LT-growth stand for short-term, long-term, and the long-term growth components of the cash-flow news proxy, respectively. The developed *CNEWS* are calculated as outlined in Appendix A.2. $DNEWS^{z}_{i,t+1}$ stands for the approach-specific discount-rate news proxy, as outlined in Section 4.1 (according to Easton & Monahan, 2005), where z stands for the specific implied risk premia approach, i.e., $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, respectively. *SD* stands for the standard deviation and *N* for the number of observations. All variables are measured as the natural log of one plus the respective variable, according to Easton and Monahan (2005).

$$RET_{i,t+1} = \beta_0 + \beta_1 RP_{i,t} + \beta_2 CNEWS^{ST}_{i,t+1} + \beta_3 CNEWS^{LT\ level}_{i,t+1} + \beta_4 CNEWS^{LT\ growth}_{i,t+1} + \beta_5 DNEWS_{i,t+1} + \varepsilon_{i,t+1}. \quad (13)$$

Second, in line with EM, we repeat the procedure with the summed cash-flow news by using the capitalization factor *p*. This puts more weight on long-term than short-term changes in expected cash flows:

$$RET_{i,t+1} = \alpha_0 + \alpha_1 RP_{i,t} + \alpha_2 CNEWS_{i,t+1} + \alpha_3 DNEWS_{i,t+1} + \vartheta_{i,t+1}. \quad (14)$$

Table 5 reports the results from the first regression (i.e., Eq. (13)), where we use the individual parts of the cash-flow news for our developed proxies (Panel A) and the EM proxy (Panel B). The coefficients and the adjusted R^2 are the mean of the annual values and the statistic is according to Fama and MacBeth (1973). Panel A shows that all risk premia metrics have a highly significant positive coefficient when using our developed cash-flow news. A test of difference indicates that all coefficients except for $RP^{E/P}$ (t = -2.44) are not statistically significantly different from one (RP^{PEG} $t = 1.80$; RP^{RIM} $t = 0.65$; RP^{AEGM} $t = 1.40$). Additionally, in line with expectations, all cash-flow and return news proxies exhibit a statistically significant and positive coefficient. This provides construct validity of the ICC estimates. In comparison, in Panel B, only the $RP^{E/P}$ coefficient is positive and statistically significant (at the 10% level), while all others are insignificant and the coefficient for the RP^{PEG} is negative. Moreover, only the third component of the cash-flow news is positive and statistically significant.

Panels C and D in Table 5 show the results of the second regression (i.e., Eq. (14)), where we use the combined cash-flow news for our developed proxies (Panel C) and the proxy of EM (Panel D). The results in Panel C confirm those from our main analysis reported in Panel A, for two out of four ICC models assessed. They provide construct validity (highly statistically significant positive coefficient for

Table 4
Correlations.

		RP				RET1	CNEWS															DNEWS			
		E/P	PEG	RIM	AEGM		EM	EM_1	EM_2	EM_3	E/P	E/P ST	E/P LT	PEG, AEGM	PEG, AEGM ST	PEG, AEGM LT-level	PEG, AEGM LT-growth	RIM	RIM ST	RIM LT-level	RIM LT-growth	E/P	PEG	RIM	AEGM
RP	**E/P**	1.00	0.51	0.79	0.50	**0.07**	0.10	0.05	0.13	0.17	0.14	0.14	0.20	0.12	0.14	0.20	0.07	0.10	0.14	0.19	0.04	0.25	0.19	0.21	0.17
	PEG	0.46	1.00	0.59	0.86	**−0.01**	0.12	0.05	0.12	0.24	0.26	0.16	0.30	0.45	0.16	0.30	0.45	0.20	0.16	0.25	0.09	0.16	0.38	0.19	0.36
	RIM	0.81	0.60	1.00	0.62	**0.01**	0.11	0.05	0.16	0.22	0.20	0.14	0.24	0.19	0.14	0.24	0.12	0.26	0.14	0.33	0.18	0.20	0.20	0.31	0.19
	AEGM	0.47	0.85	0.61	1.00	**0.01**	0.09	0.02	0.09	0.22	0.22	0.13	0.27	0.40	0.13	0.27	0.41	0.16	0.13	0.22	0.09	0.16	0.36	0.19	0.38
RET1		0.07	0.00	0.04	0.02	1.00	**0.18**	0.01	0.36	0.40	**0.43**	0.39	0.47	**0.24**	0.39	0.47	0.02	**0.34**	0.39	0.46	0.12	**0.38**	**0.22**	**0.25**	**0.20**
CNEWS	**EM**	0.16	0.24	0.21	0.22	0.42	1.00	0.57	0.50	0.53	0.30	0.34	0.40	0.16	0.34	0.40	0.00	0.20	0.34	0.36	0.04	**−0.05**	**−0.02**	**−0.07**	**−0.01**
	EM_1	0.04	0.07	0.05	0.05	0.08	0.50	1.00	**0.08**	0.07	0.03	0.04	0.04	0.03	0.04	0.04	0.02	0.02	0.04	0.03	0.00	0.01	0.02	0.01	0.02
	EM_2	0.14	0.15	0.18	0.15	0.47	0.79	0.23	1.00	**0.87**	0.53	0.73	0.69	0.23	0.73	0.69	0.07	0.35	0.73	0.63	0.07	0.08	0.01	0.12	0.02
	EM_3	0.19	0.29	0.26	0.29	0.51	0.83	0.21	0.84	1.00	0.63	0.60	0.81	0.36	0.60	0.81	0.04	0.43	0.60	0.73	0.11	0.13	0.09	0.18	0.07
	E/P	0.17	0.29	0.23	0.27	0.49	0.78	0.20	0.80	0.94	1.00	**0.62**	**0.80**	0.54	0.62	0.80	0.00	0.66	0.62	0.69	0.10	**−0.23**	0.14	0.33	0.12
	E/P ST	0.14	0.15	0.16	0.15	0.46	0.76	0.23	0.97	0.80	0.80	1.00	**0.76**	0.24	1.00	0.76	0.12	0.39	1.00	0.65	0.06	**−0.12**	0.01	0.16	0.03
	E/P LT	0.19	0.29	0.24	0.29	0.51	0.81	0.20	0.82	0.97	0.96	0.83	1.00	0.42	0.76	1.00	0.00	0.52	0.76	0.87	0.11	**−0.19**	0.11	0.25	0.08
	PEG,AEGM	0.11	0.47	0.19	0.45	0.27	0.40	0.11	0.32	0.49	0.51	0.32	0.49	1.00	**0.24**	**0.42**	**0.63**	0.55	0.24	0.51	0.16	0.24	**−0.74**	0.36	**−0.72**
	PEG,AEGM ST	0.14	0.15	0.16	0.15	0.46	0.76	0.23	0.97	0.80	0.80	1.00	0.83	0.32	1.00	**0.76**	0.12	0.39	1.00	0.65	0.06	0.12	**0.01**	0.16	**0.03**
	PEG,AEGM LT-level	0.19	0.29	0.24	0.29	0.51	0.81	0.20	0.82	0.97	0.96	0.83	1.00	0.49	0.83	1.00	0.00	0.52	0.76	0.87	0.11	0.19	**−0.11**	0.25	**−0.08**
	PEG,AEGM LT-growth	0.06	0.40	0.12	0.40	0.05	0.05	0.01	0.06	0.10	0.09	0.07	0.09	0.79	0.07	0.09	1.00	0.16	0.12	0.19	0.19	0.10	**−0.58**	0.15	**−0.56**
	RIM	0.12	0.23	0.26	0.21	0.36	0.49	0.11	0.49	0.59	0.62	0.49	0.60	0.54	0.49	0.60	0.27	1.00	**0.39**	**0.66**	**0.55**	0.16	0.26	**−0.49**	0.26
	RIM ST	0.14	0.15	0.16	0.15	0.46	0.76	0.23	0.97	0.80	0.80	1.00	0.83	0.32	1.00	0.83	0.07	0.49	1.00	**0.65**	0.06	0.12	0.01	**−0.16**	0.03
	RIM LT-level	0.18	0.25	0.30	0.25	0.48	0.68	0.16	0.69	0.83	0.81	0.69	0.84	0.59	0.69	0.84	0.26	0.77	0.69	1.00	0.28	0.20	0.23	**−0.38**	0.21
	RIM LT-growth	0.04	0.12	0.16	0.13	0.15	0.15	0.02	0.13	0.19	0.18	0.12	0.18	0.26	0.12	0.18	0.28	0.73	0.12	0.35	1.00	0.01	0.10	**−0.20**	0.10
DNEWS	**E/P**	0.31	0.19	0.26	0.20	0.49	0.14	0.07	0.11	0.16	0.16	0.12	0.18	0.18	0.12	0.18	0.13	0.12	0.12	0.19	0.01	1.00	0.65	0.83	0.59
	PEG	0.20	0.40	0.22	0.40	0.24	0.10	0.04	0.01	0.15	0.14	0.01	0.15	0.68	0.01	0.15	0.74	0.25	0.01	0.27	0.17	0.60	1.00	0.64	0.97
	RIM	0.24	0.22	0.34	0.22	0.33	0.19	0.07	0.16	0.23	0.23	0.16	0.24	0.27	0.16	0.24	0.20	0.37	0.16	0.39	0.24	0.81	0.59	1.00	0.61
	AEGM	0.17	0.37	0.20	0.42	0.26	0.06	0.03	0.03	0.10	0.10	0.03	0.10	0.63	0.03	0.10	0.71	0.22	0.03	0.21	0.17	0.55	0.94	0.57	1.00

This table provides the mean of annual correlations (Pearson Product Moment above and Spearman Rank Order below the diagonal). The four implied risk premia estimates, i.e., implied cost of capital less the prevailing risk-free rate, $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, are estimated as outlined in Appendix B. RET_1 is the one-year-ahead buy-and-hold return following the calculation of the implied risk premia, less the risk-free rate. $CNEWS^{EM}$ is the cash-flow news as outlined in Equation (2) (according to Easton & Monahan, 2005). *EM_1* stands for the first part of the $CNEWS^{EM}$, i.e., $(ROE_{i,t} \quad FROE_{i,t,t})$, EM_2 for the second part, i.e., $(FROE_{i,t+1,t+1} \quad FROE_{i,t,t+1})$, and EM_3 for the third part, i.e., $(FROE_{i,t+1,t+2} \quad FROE_{i,t,t+2})$. $CNEWS^{E/P}_{i,t+1}$ ($CNEWS^{PEG,AEGM}_{i,t+1}$) is the approach-specific cash-flow news proxy for the $RP^{E/P}$ (RP^{PEG} and RP^{AEGM}). $CNEWS^{RIM}_{i,t+1}$ is the approach-specific cash-flow news proxy for the RP^{RIM}. ST, LT-level, and LT-growth stand for short-term, long-term, and the long-term growth components of the cash-flow news proxy, respectively. The developed *CNEWS* are calculated as outlined in Appendix A.2. $DNEWS^{z}_{i,t+1}$ stands for the approach-specific discount-rate news proxy as outlined in Section 4.1 (according to Easton & Monahan, 2005). z stands for the specific implied risk premia approach, i.e., $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, respectively. All variables are measured as the natural log of one plus the respective variable according to Easton and Monahan (2005). The sample size is 43,465 observations.

the RP) for the RP^{PEG} and RP^{AEGM}.[19] The coefficients for the cash-flow and return news proxies are again significantly positive. Not surprisingly, the results in Panel D are almost identical to those reported in Table 4 of EM (2005, p. 518). The coefficients of the expected return proxies are negative (three out of four are significantly negative), while those for the cash-flow and return news are significantly positive.

Additionally, as we discuss in section 2.3, we also estimate the modified noise variable according to EM (2005), with reference to Garber and Klepper (1980) and Barth (1991). This is reported in the last column in each panel of Table 5. The values in relation to each of the $RP_{i,t}$ tested against realized returns are far below 0.01 and thus confirm the results that the assessed ICC estimates constitute valid proxies for the expected return.[20]

When comparing the results between Panel A (C) and Panel B (D), one observes a positive sign for both types of news irrespective of the definition of cash-flow news. However, arguably, our developed proxies used in Panel A (C) perform better, given that the *t*-statistics are considerably higher than the corresponding ones in Panel B (D). Furthermore, the adjusted R^2 using our cash-flow news proxies, which range between 36% and 48%, are substantially higher than those using the EM proxy, which range between 11% and 38%. These findings provide empirical evidence in support of our analytical discussion, which suggests that cash-flow news have to be consistently defined.[21]

A comparison of the results (especially for the adjusted R^2) between Panels A and B (separate cash-flow news components) and Panels C and D (combined cash-flow news) shows that a linear combination of the individual cash-flow news components is superior to a non-linear combination i.e., less weight on long-term news is better.[22] Consequently, we argue for the use of the individual components. Nevertheless, our results provide evidence that even when imposing a specific capitalization for the cash-flow news, two out of four approaches (AEGM and PEG) reveal a positive significant sign.[23]

We also address potential multicollinearity with the maximum value of the variance inflation factor (VIF) across the annual regressions. The results indicate no problem of multicollinearity in any results in Panels A, C, and D in Table 5.[24]

4.5. Additional analysis and robustness checks

Appendix B discusses in detail and reports in four Tables the findings of five sets of additional tests we conduct for checking the sensitivity of our findings. First, to alleviate any concerns that our findings are driven by the different/longer period in our study compared to that in EM, we repeat our analysis by using the same sample period as EM do. Second, to further investigate potential multicollinearity for our developed cash-flow news proxies, we apply a procedure in line with Yu (2008). Third, to alleviate any potential concerns of mechanical relation between our cash-flow news and return news proxies, which may cause the positive coefficients of RP, we substitute the firm-specific return news with an average return news based on the average of the four ICC proxies. This is consistent with one of Botosan et al.'s (2011) sensitivity tests. Fourth, we consider evidence in relevant studies that rely only on cash-flow news when considering firm-level realized returns (e.g., Ogneva, 2012). Fifth, we repeat our analysis by using Botosan et al.'s (2011) definition of the firm-specific return news proxy (i.e., the change in market beta). The findings discussed and reported in the corresponding four tables in Appendix B demonstrate that the findings presented in Table 5 are not sensitive to any of these research design choices we have made in our main tests.

[19] All proxies are statistically distinguishable from the hypothetical benchmark of one ($RP^{E/P}$ t = -5.06; RP^{PEG} t = -2.27; RP^{RIM} t = -4.11; RP^{AEGM} t = -3.94).

[20] As EM point out (see details on page 507), the modified noise variable changes with different RP metrics as expected, but not with different cash flow news estimation.

[21] In untabulated tests, we further elaborate on the need for consistency of cash-flow news applied and find, in line with our expectations, that a consistent definition is of importance. More specifically, we have repeated the analysis presented in Panel A of Table 5 by, first, using the ST cash-flow news only and, second, by using the LT cash-flow news only. When using short-term cash-flow news only, the coefficient on the ICC is insignificant, when, however, using the long-term level and growth, the coefficients for the ICC across all four models tested turn significantly positive. The results from these tests are in line with the arguments in the literature (e.g. Ohlson, 2005; Ohlson & Gao, 2006) and those we put forward in this study: it is the long-term cash-flow news that really matter and thus their modeling and measurement when validating ICC measures matter as well.

[22] We acknowledge that, arguably, the EM capitalization factor we apply may be somewhat inappropriate for our more recent and updated sample.

[23] This observation might be explained by the generally agreed assertion that the *AEGM* and *PEG* are more sensitive to future growth assumptions than the *RIM* or the *E/P* (e.g., Gode & Mohanram, 2003). Given that the capitalization method puts more weight on the long-term part, which includes long-term growth, this approach might work better for the *AEGM* and *PEG* than for the *RIM* or *E/P*.

[24] The max VIFs in Panels A and C range between 1.89 and 5.37, while those in Panel D range between 1.34 and 1.43. All values are well below ten, which is an indication of no problem of multicollinearity (Gujarati, 2003, p. 262). In Panel B, where we apply the EM (2005) proxy for cash-flow news, we observe VIFs higher than ten, which indicates multicollinearity. To further investigate this issue, we remove the second component of the cash-flow news, i.e., $FROE_{i,t+1,t+1}$ $FROE_{i,t,t+1}$ because of the high correlation with $FROE_{i,t+1,t+2}$ $FROE_{i,t,t+2}$ (see Table 4 herein). The results for the *RP*, the remaining *CNEWS*, and *DNEWS* are virtually identical, and the VIFs drop to values below two. Thus, multicollinearity does not drive the results obtained in Panel B. Additionally, in the case of multicollinearity, all coefficients in the regression would be insignificant (cf. EM 2016). However, there is a consistent, positive, statistically highly significant relation between the news proxies and realized returns (see Panels A and C).

Table 5
Implied risk premia and realized returns.

Panel A: Regression results for individual CNEWS as developed in this paper

RP metric	Intercept	RP	CNEWS-components			DNEWS	Adj. R^2	max. VIF	Mod. Noise Var.
			ST	LT-level	LT-growth				
$RP^{E/P}$	0.0725	0.5994	0.0326	0.5274		0.1094	48.00	3.50	0.0003
t-value	(3.54)***	(3.65)***	(2.87)***	(26.25)***		(18.36)***			
RP^{PEG}	0.0231	1.3262	0.0431	0.5122	0.1704	0.0816	38.65	3.80	0.0014
t-value	(1.06)	(7.31)***	(3.49)***	(24.67)***	(10.72)***	(10.38)***			
RP^{RIM}	0.0355	1.1259	0.0634	0.5552	0.0506	0.1298	45.69	3.79	0.0006
t-value	(1.82)**	(5.79)***	(4.93)***	(24.10)***	(6.90)***	(19.85)***			
RP^{AEGM}	0.0054	1.2201	0.0404	0.4946	0.1489	0.0675	36.16	3.89	0.0012
t-value	(0.25)	(7.71)***	(3.09)***	(22.97)***	(9.33)***	(11.57)***			

Panel B: Results for individual CNEWS according to Easton and Monahan (2005)

RP metric	Intercept	RP	$CNEWS^{EM}$-components			DNEWS	Adj. R^2	max. VIF	Mod. Noise Var.
			$(ROE_{i,t}$ $FROE_{i,t,t})$	$(FROE_{i,t+1,t+1}$ $FROE_{i,t,t+1})$	$(FROE_{i,t+1,t+2}$ $FROE_{i,t,t+2})$				
$RP^{E/P}$	0.0543	0.3368	0.0312	0.0292	2.6591	0.0989	38.14	11.05	0.0003
t-value	(2.54)**	(1.89)*	(-1.60)	(-0.14)	(14.65)***	(17.23)***			
RP^{PEG}	0.0569	0.0508	0.0254	0.2716	2.6830	0.0541	25.97	11.95	0.0014
t-value	(2.73)***	(-0.23)	(-1.16)	(-1.14)	(12.34)***	(7.84)***			
RP^{RIM}	0.0530	0.2129	0.0365	0.0091	2.6500	0.0904	30.28	11.27	0.0006
t-value	(2.55)**	(0.86)	(-1.77)*	(-0.04)	(12.35)***	(13.15)***			
RP^{AEGM}	0.0458	0.1312	0.0251	0.2501	2.6419	0.0457	24.77	11.95	0.0012
t-value	(2.21)**	(0.67)	(-1.17)	(-1.02)	(12.04)***	(9.35)***			

Panel C: Results for combined CNEWS as developed in this paper

RP metric	Intercept	RP	CNEWS	DNEWS	Adj. R^2	max. VIF	Mod. Noise Var.
$RP^{E/P}$	0.0647	0.1295	0.0102	0.1184	44.68	1.89	0.0003
t-value	(2.93)***	(0.75)	(13.93)***	(20.40)***			
RP^{PEG}	0.0389	0.6562	0.0186	0.1785	42.65	5.37	0.0014
t-value	(1.84)**	(4.34)***	(14.64)***	(21.16)***			
RP^{RIM}	0.0540	0.0883	0.0146	0.1574	37.03	3.20	0.0006
t-value	(2.67)***	(0.40)	(11.49)***	(22.67)***			
RP^{AEGM}	0.0386	0.4466	0.0167	0.1526	37.49	3.84	0.0012
t-value	(1.90)*	(3.18)***	(13.20)***	(24.08)***			

Panel D: Results for combined CNEWS according to Easton and Monahan (2005)

RP metric	Intercept	RP	$CNEWS^{EM}$	DNEWS	Adj. R^2	max. VIF	Mod. Noise Var.
$RP^{E/P}$	0.0493	0.0648	0.2342	0.0885	21.72	1.43	0.0003
t-value	(2.21)**	(-0.31)	(7.35)***	(15.51)***			
RP^{PEG}	0.0704	0.7977	0.2078	0.0491	11.67	1.40	0.0014
t-value	(3.32)***	(-3.38)***	(7.02)***	(8.83)***			
RP^{RIM}	0.0630	0.5187	0.2314	0.0745	14.37	1.40	0.0006
t-value	(2.94)***	(-1.83)*	(7.06)***	(9.69)***			
RP^{AEGM}	0.0733	0.5723	0.2113	0.0437	10.54	1.34	0.0012
t-value	(3.45)***	(-2.82)***	(7.05)***	(9.03)***			

This table provides results of regressions of realized returns on expected returns as well as cash flow and return news. The coefficient and the adjusted R^2 are the mean of annual values, and the statistic is according to Fama and MacBeth (1973). Panels A and B provide results with individual cash-flow news proxies and Panels C and D provide results with the summarized cash-flow news proxies (see Appendix A.2 for more details). The four implied risk premia estimates, i.e., implied cost of capital less the prevailing risk-free rate, $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, are estimated as outlined in Appendix B. RET_1 is the one-year-ahead buy-and-hold return following the calculation of the implied risk premia, less the risk-free rate. $CNEWS^{EM}$ is the cash-flow news as outlined in Equation (2) (according to Easton & Monahan, 2005). $CNEWS^{E/P}_{i,t+1}$ ($CNEWS^{PEG,AEGM}_{i,t+1}$) is the approach-specific cash-flow news proxy for the $RP^{E/P}$ (RP^{PEG} and RP^{AEGM}). $CNEWS^{RIM}_{i,t+1}$ is the approach-specific cash-flow news proxy for the RP^{RIM}. ST, LT-level, and LT-growth stand for short-term, long-term, and the long-term growth components of the cash-flow news proxy, respectively. The developed CNEWS are calculated as outlined in Appendix A.2. $DNEWS^{z}_{i,t+1}$ stands for the approach-specific discount-rate news proxy as outlined in Section 4.1 (according to Easton & Monahan, 2005). z stands for the specific implied risk premia approach, i.e., $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, respectively. All variables are measured as the natural log of one plus the respective variable according to Easton and Monahan (2005). Max. VIF is the maximum value of variance inflation factor across the annual regressions. Mod. Noise Var. is the modified noise variable computed as in Easton and Monahan (2005) (see footnote 7). T-value is the test of difference from zero. ***, **, and * indicate significance at the 1%, 5%, and 10% level, respectively. The sample size is 43,465 observations.

5. Application of the EM procedure and analysts' bias

We further elaborate on EM's findings of significant positive coefficients when they build sub-samples depending on the extent of the absolute forecast error. Moreover, we discuss findings of more recent studies that show significant positive coefficients using EM's methodology when correcting for analysts' forecast error (e.g., Mohanram & Gode, 2013; Nekrasov & Ogneva, 2011). With regard to the latter, the discussion reflects on the design applied in Mohanram and Gode (2013) (hereafter MG 2013), although the arguments are also relevant to older papers.

Table 6
Implied risk premia and realized returns: investigating the impact of analysts' bias.

Panel A: Regression results using the observations with the lowest third of absolute forecast error								
RP metric	Intercept	RP	CNEWS-components			DNEWS	Adj. R^2	max. VIF
			ST	LT-level	LT-growth			
$RP^{E/P}$	0.0823	0.7576	0.0266	0.6767		0.1731	47.71	2.87
t-value	(4.51)***	(3.56)***	(0.64)	(17.07)***		(9.58)***		
RP^{PEG}	0.0166	2.0127	0.0293	0.7522	0.1863	0.1358	36.03	3.48
t-value	(0.83)	(8.73)***	(-0.74)	(16.03)***	(10.02)***	(4.77)***		
RP^{RIM}	0.0227	2.0019	0.1005	0.6678	0.0175	0.1786	44.26	2.39
t-value	(1.16)	(7.76)***	(2.24)**	(19.02)***	(1.87)*	(16.45)***		
RP^{AEGM}	0.0227	2.0711	0.0405	0.7516	0.1602	0.1054	32.13	3.61
t-value	(-1.10)	(9.09)***	(-1.04)	(15.53)***	(9.25)***	(7.39)***		
Panel B: Regression results using the observations with the highest third of absolute forecast error								
RP metric	Intercept	RP	CNEWS-components			DNEWS	Adj. R^2	max. VIF
			ST	LT-level	LT-growth			
$RP^{E/P}$	0.0466	0.5497	0.0370	0.4626		0.0897	47.17	4.04
t-value	(1.97)**	(3.17)***	(3.47)***	(23.49)***		(18.31)***		
RP^{PEG}	0.0012	1.0420	0.0467	0.4394	0.1724	0.0731	38.95	4.21
t-value	(-0.04)	(5.62)***	(3.99)***	(22.28)***	(9.65)***	(8.98)***		
RP^{RIM}	0.0243	0.6735	0.0599	0.4816	0.0631	0.1075	43.92	3.70
t-value	(1.13)	(3.42)***	(5.31)***	(21.17)***	(5.68)***	(15.27)***		
RP^{AEGM}	0.0152	0.9878	0.0453	0.4234	0.1514	0.0589	36.89	4.14
t-value	(-0.58)	(6.13)***	(3.74)***	(21.51)***	(8.55)***	(9.75)***		
Panel C: Regression results using adjusted earnings forecasts								
RP metric	Intercept	A_RP	CNEWS-components			A_DNEWS	Adj. R^2	max. VIF
			A_ST	A_LT-level	A_LT-growth			
$RP^{E/P}$	0.0673	0.8258	0.0046	0.5035		0.0795	51.96	4.87
t-value	(3.18)***	(5.50)***	(-0.38)	(18.15)***		(9.81)***		
RP^{PEG}	0.0198	1.5690	0.030	0.4911	0.0844	0.0329	47.25	5.22
t-value	(0.82)	(7.37)***	(-2.15)**	(19.78)***	(7.49)***	(7.50)***		
RP^{RIM}	0.0081	1.8503	0.0016	0.6188	0.0240	0.0982	55.40	5.68
t-value	(0.39)	(9.99)***	(-0.16)	(21.87)***	(2.03)**	(10.39)***		
RP^{AEGM}	0.0142	1.0132	0.0358	0.4646	0.0521	0.0222	44.27	5.18
t-value	(0.61)	(6.83)***	(-2.41)**	(17.35)***	(5.34)***	(5.49)***		

This table provides results of regressions of realized returns on expected returns as well as cash flow and return news. Panel A consists of the bottom third of the distribution of the absolute forecast error and Panel B consists of the observations of the top third of the distribution of the absolute forecast error. Panel C applies adjusted earnings forecasts when determining the implied risk premia, the cash-flow news and the return news (all adjusted variables are thus denoted as A_). The coefficient and the adjusted R^2 are the mean of annual values, and the statistic is according to Fama and MacBeth (1973). The adjustment process is as outlined in Appendix C. The four implied risk premia estimates, i.e., implied cost of capital less the prevailing risk-free rate, $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, are estimated as outlined in Appendix B. RET_1 is the one-year-ahead buy-and-hold return following the calculation of the implied risk premia, less the risk-free rate. $CNEWS^{EM}$ is the cash-flow news as outlined in Equation (2) (according to Easton & Monahan, 2005). $CNEWS^{E/P}_{i,t+1}$ ($CNEWS^{PEG,AEGM}_{i,t+1}$) is the approach-specific cash-flow news proxy for the $RP^{E/P}$ (RP^{PEG} and RP^{AEGM}). $CNEWS^{RIM}_{i,t+1}$ is the approach-specific cash-flow news proxy for the RP^{RIM}. ST, LT-level, and LT-growth stand for short-term, long-term, and the long-term growth components of the cash-flow news proxy, respectively. The developed CNEWS are calculated as outlined inAppendix A.2. $DNEWS^{z}_{i,t+1}$ stands for the approach-specific discount-rate news proxy as outlined in Section 4.1 (according to Easton & Monahan, 2005). z stands for the specific implied risk premia approach, i.e., $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, respectively. All variables are measured as the natural log of one plus the respective variable according to Easton and Monahan (2005). Max. VIF is the maximum value of variance inflation factor across the annual regressions. T-value is the test of difference from zero. ***, **, and * indicate significance at the 1%, 5%, and 10% level, respectively. The sample size in Panels A and B is 14,488 observations, i.e., one third of 43,465. The sample size in Panel C is 34,519 observations, i.e. 43,465 observations less observations for which data to adjusted earnings forecasts is missing (see the adjustment process in Appendix C).

EM (Panel C of their Table 9) show that three ICC proxies are significantly positively associated with returns for a sub-sample that consists of the observations with the lowest absolute forecast error. They also report that the RIM approach (Claus & Thomas, 2001) shows a significant positive relation but only for observations in the middle third of the population.

Based on this, MG (2013, Panel B of their Table 4) show that common ICC estimates are not significantly positively related to realized returns after controlling for cash-flow and return news. Thus, they conclude that, "[c]onsistent with the findings of EM, none of the risk premium metrics have significant coefficients" (MG 2013, p. 459). We argue that this is not surprising, as they follow the commonly used research design for validating ICC measures. Second, after adjusting ICC estimates for predictable analysts' forecast errors and repeating the same validation procedure, MG (2013, Panel B of their Table 8) show that "removing predictable forecast errors improves the association between implied risk premium and realized returns after controlling for cash-flow news and discount-rate news" (MG 2013, p. 472), and that "all the theoretically motivated implied risk premium metrics have a significant positive coefficient" (MG 2013, p. 471–472).

When adjusting for analysts' optimism and/or using sub-samples consisting of firms with low absolute forecast errors, low absolute forecast errors are accompanied by low absolute news for that period. Hence, by construction, either when building sub-samples (EM) or when adjusting for analysts' optimism (MG 2013) while computing the ICC measures, the news that occurs between t and t+1 is less important. Consequently, the potential inconsistency we identify in the definition of the cash-flow news in commonly used validation framework becomes of lower importance.

We argue that it is the decreased influence of this weakness that allows EM and MG (2013) to provide empirical support for the construct validity of ICC estimates, when adjusting for analysts' optimism or when constructing sub-samples.

To elaborate on our argument, first, following EM, we build sub-samples based on the absolute forecast error for period t+1 scaled by price at t, and repeat the analysis presented in Panel A of Table 5. Panels A and B of Table 6 report these results.

Panel A consists of the bottom third of the distribution of the absolute forecast error, mirroring the results presented by EM (their Table 9). Consistent with our results (Table 5, Panel A), the coefficients for the risk premia are always positive and statistically significant (and only $RP^{E/P}$ t = -1.13 is not statistically different from one). From the cash-flow news proxies, only the coefficients for the long-term effects are positive and significant, while those for the short-term effects are insignificant in three out of four approaches. This is consistent with expectations: for the observations with the lowest absolute forecast error, the short-term cash-flow news is by construction low if not close to zero. The return news coefficient is, again as expected, positive and significant.

Panel B of Table 6 consists of the top third of the distribution of the absolute forecast error. As expected, the coefficients for the expected return, cash-flow news and return news are positive and statistically significant (three risk premia proxies are not different from one: RP^{PEG} t = 0.23; RP^{RIM} t = -1.65; RP^{AEGM} t = -0.07). Compared to Panel A, the coefficients for the short-term cash-flow news proxies are also positive and statistically significant. This is intuitive and not surprising, because observations with high absolute forecast errors exhibit cash-flow news for the period ending in t+1 which is captured by the short-term cash flows.[25]

These confirm that the consistent cash-flow news proxies we develop capture cash-flow news more accurately than the EM proxy. The latter only works if cash-flow news is less important i.e., for observations with a low absolute forecast error.

Second, when using adjusted analysts' forecasts, MG (2013) show positive and significant risk premia coefficient. However, Panel B of Table 8 in MG (2013), which reports the results for the adjusted ICC estimates, shows a negative sign for the return news for four out of six methods, although this should be positive. In fact, when MG (2013, Panel B of their Table 4) report results for unadjusted estimates, indeed they show a consistently positive sign for the discount-rate news. Additionally, their reported adjusted-R^2 for the unadjusted (adjusted) estimates range between 12.7% and 34.3% (8.85%–17.06%). The signs for return news and a decreasing adjusted-R^2 are again against expectations.

Third, following MG (2013) and Larocque (2013), we adjust earnings forecasts for predictable analysts' forecast errors and repeat our Table 5 (Panel A) analysis (using our cash-flow news proxies).[26] The results are reported in Panel C of Table 6 and show that, consistent with theoretical expectations, the coefficients for the expected returns, long-term cash-flow news, and return news are significantly positive. As in Panel A of Table 6, the short-term cash-flow news coefficients are no longer positive and significant. Thus, we confirm that adjusting for predictable forecast errors reduces the short-term forecast errors, which, in turn, decreases the short-term cash-flow news. Considering the negative sign for the return news and a decreasing adjusted-R^2 in MG (2013), the theoretically consistent results for all the independent variables in Panel C of Table 6 indicates that our cash-flow news definition provides more consistent results than those based on the EM procedure.

Finally, we further examine the benefit of adjusting for analysts' optimism, given our proposed adjustment in the cash-flow news proxies used in the validation framework. From Table 5 (Panel A) one can see that three out of four ICC methods show a coefficient not distinguishable from the benchmark of one. Instead, the results using adjusted forecasts (Table 6, Panel C) show that only two out of four approaches are not distinguishable from the benchmark of one ($RP^{E/P}$ t = -1.16; RP^{AEGM} t = 0.09). However, as discussed, the value of one is a theoretical benchmark, and a positive sign indicates validity of the expected return proxy. This is observable for all coefficients of the expected return proxies in Table 5 (Panel A, unadjusted analysts' forecasts) and in Table 6 (Panel C, adjusted analysts' forecasts). Thus, adjusting analysts' earnings forecasts for the estimation of ICC measures does not lead to superior ICC estimates to proxy expected returns.

Arguably, this evidence might be somewhat puzzling given the evidence that analysts' forecasts contain substantial biases (e.g.,

[25] In untabulated analysis, we also repeat the procedure for the middle third of the distribution of the absolute forecast error. In line with Panels A and B of Table 6, the results show positive and significant coefficients for the risk premia, the cash-flow news, and the return news.

[26] The procedure of how we correct for analysts' optimism is summarized in Appendix C.

Frankel & Lee, 1998; Hou et al., 2012). However, biases/errors in forecasts relative to reported earnings are not our concern. Our concern is whether analysts' expectations are in line with the market ones. Even so, if there is a systematic analysts' forecasts bias which may lead to overstated ICC estimates, this does not necessarily suggest that the ICC is a poor proxy and/or that adjusting for optimism is necessary. Measurement errors in the ICC have to be systematic and correlated with other variables (see also the discussion in Botosan et al., 2011).[27]

A further practical advantage when relying on unadjusted forecasts is the relaxation of extensive data requirements for adjusting analysts' forecasts e.g., lagged forecast errors or fundamental data to estimate future forecast errors (see also Appendix C and MG (2013)). Nor does it require a "comprehensive model" (MG 2013, p. 445).

6. Conclusion

Informed by recent finance literature that emphasizes the importance of cash-flow vs. return news (e.g., Chen et al., 2013; Chen & Zhao, 2009; Khimich, 2017), conceptually and within the general implementation of the Vuolteenaho (2002) framework, we provide evidence that the revision in analysts' forecasts outperforms the VAR approach in measuring cash-flow news i.e., it is a better proxy for markets' expected cash-flows than the VAR method. Thus, we provide an alternative solution to the construct validity puzzle in the ICC literature (Botosan et al., 2011; Botosan & Plumlee, 2005; Easton & Monahan, 2005, 2016; Larocque, 2013; Mohanram & Gode, 2013).

We focus on the second part of the news components i.e., the cash-flow news, and we outline model-specific cash-flow news proxies for the widely used ICC models, which focus on how each specific ICC model deals with expected cash-flows. Accordingly, our proposed cash-flow news proxies are aligned with the measurement of the cash flows in the expected return proxy and the return news proxy.

We show that commonly used ICC estimates are highly significantly positively related to realized returns, after applying the proposed specification of cash-flow news. We report that the coefficient of the implied risk premia is not statistically different from the theoretical benchmark of one, and is statistically positive for three out of the four ICC approaches assessed. Hence, our results speak to the strong validity of ICC estimates. We propose that ICC estimates are reliable proxies for expected returns and future research can employ this alternative cash-flow news specification when testing the validity of a newly introduced ICC measure.

Subsequently, we turn to prior literature that accounts and corrects for biases due to analysts' optimism. This stream of literature implements the commonly used validation framework and shows that ICC estimates adjusted for analysts' bias are significantly positively related to future realized returns (e.g., Larocque, 2013; Gode & Mohanram, 2003). However, we report that this evidence is driven by the definition of cash-flow news when applying the commonly used procedure, and not by unadjusted ICC estimates. Our results indicate that adjusted for analysts' optimism ICC estimates are not superior to their unadjusted counterparts, when construct validity tests have captured cash-flow news in the way we propose in this study. Thus, we recommend using unadjusted earnings forecasts for estimating ICC. This goes along with the advantage of less restrictive data requirements and labor-intensive estimations when estimating an expected return proxy with this approach (cf. Gode & Mohanram, 2003).

Data availability

The authors do not have permission to share data.

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APPENDICES.

Appendix A. Models applied and cash-flow news derivation

A.1: Models applied and tested for validation

For empirical implementation, we estimate ICC for the most widely used models in the literature, as briefly reviewed below. As is common, all implementations are applied on a per-share basis.

ICC based on common heuristics. First, we estimate PEG according to Easton (2004), which is derived by setting the long-term growth rate and the dividends in Eq. A.4 equal to zero. The *PEG* looks as follows (Eq. A.1):

[27] It may also be that existing approaches to detect and correct for analysts' optimism do not perform well i.e., they do not adequately adjust for any relevant systematic bias.

$$r^{PEG} = \sqrt{\frac{(EPS_2 \quad EPS_1)}{P_0}}. \tag{A.1}$$

Motivated by Ohlson and Johannesson (2016), we further include as a benchmark the forward earnings to price ratio as defined below (Eq. A.2)[28]:

$$r^{E/P} = \frac{EPS_2}{P_0}. \tag{A.2}$$

ICC based on RIM according to Claus and Thomas (2001). Claus and Thomas (2001) developed the most widely used model to estimate ICC using RIM. RIM estimates the cost of capital by equating current stock price to the sum of the current book value and the present value of future residual income. Claus and Thomas (2001) rely on five years of forecasts (T = 5)[29] and assume long-term growth rate (g_{CT}) at the expected inflation rate, which is set to $r_f \quad 3\%$ ($g_{CT} = r_f \quad 3\%$). The equation is as follows (Eq. A.3):

$$P_0 = BVPS_0 + \sum_{t=1}^{T \ 1} \frac{EPS_t \quad r^{CT} BVPS_{t \ 1}}{(1 + r^{CT})^t} + \frac{EPS_T \quad r^{CT} BVPS_{T \ 1}}{(r^{CT} \quad g_{CT})(1 + r^{CT})^{T \ 1}}. \tag{A.3}$$

where $BVPS_t$ is the book value per share at t, EPS_t are forecasted earnings per share for time t, P_0 is the price per share at the end of December, when the forecasts are collected, and r^{CT} is the resulting ICC from an iterative procedure to solve Equation A.3.

ICC based on AEGM according to Gode and Mohanram (2003). The cost of capital based on the AEGM as developed by Gode and Mohanram (2003) is based on the following equation (Eq. A.4):

$$r^{AEGM} = A + \sqrt{A^2 + \frac{EPS_1}{P_0} * \left[\left(\frac{EPS_2}{EPS_1} \quad 1\right) \quad g_{GM}\right]}, \tag{A.4}$$

where A is determined according to Eq. B.5:

$$A = \frac{1}{2}\left(g_{GM} + \frac{DPS_1}{P_0}\right), \tag{A.5}$$

where DPS_1 is the expected one-year-ahead dividend per share, and the rest of the variables are defined as above. Consistent with Gode and Mohanram (2003), we set long-term growth rate (g_{GM}) equal to the risk-free rate minus 3% ($g_{GM} = r_f \quad 3\%$).[30]

A.2 Cash-flow news proxies outlined and applied in this study

In line with EM and Botosan et al. (2011), we split the news proxy into j different news proxies:

$$CNEWS_{i,t+1} = \sum_{1}^{j} CF_{i,t+1,j} \quad CF_{i,t,j}). \tag{A.6}$$

The first news component is the same across all risk premia methods and deals with the issue that, during the period between t and t+1, a change in the expected cash-flows (i.e., earnings) for the period ending in t+1 could occur. This leads to:

$$CNEWS_{i,t+1} = \left(\frac{E_{i,t+1,t+1}}{E_{i,t,t+1}} \quad 1\right) + \sum_{2}^{j} CF_{i,t+1,j} \quad CF_{i,t,j}). \tag{A.7}$$

where $E_{i,u,v}$ denotes forecasted earnings for fiscal year v and is based on forecasts of earnings made in December of year u. We refer to this as the short-term (ST) cash-flow news.

Regarding the long-term (LT) cash-flow news, it is assumed to be ICC model-specific, depending on how each specific model uses forecasts. We start with the earnings-to-price ratio, which expresses ICC based on the level of forecasted earnings and the current price

[28] Review often applies EPS1 instead of EPS 2 in empirical models. Given the recent evidence in Ohlson and Johannesson (2016), we opted to use EPS 2. However, we note that due to the high correlation between EPS 1 and EPS 2 (above 95% for the sample), cross-sectional results would be qualitatively similar using EPS 1 instead of EPS 2.

[29] Due to the data collection procedure we follow, i.e., we use forecasts as of the last Thursday in December, we have a maximum of four years (T = 4) of forecasts.

[30] We do not adjust the short-term growth part in the *AEGM* $\left(\frac{EPS_2}{EPS_1} \quad 1\right)$ because of the arbitrary nature of any adjustment and to keep the model equivalent to the present value of the expected dividend model. Moreover, when adjusting the short-term growth to be the average between $\left(\frac{EPS_2}{EPS_1} \quad 1\right)$ and the I/B/E/S long-term growth rate, the results are nearly the same.

(i.e., $r^{E/P} = \frac{E_2}{P_0}$). As a result, cash-flows news should consider a relative change in the absolute forecast level in period two. The corresponding LT cash-flow proxy should, thus, focus on the earnings change in t+2 only. This leads to the following cash-flow news proxy:

$$CNEWS^{E/P}_{i,t+1} = \left(\frac{E_{i,t+1,t+1}}{E_{i,t,t+1}} \quad 1\right) + \left(\frac{E_{i,t+1,t+2}}{E_{i,t,t+2}} \quad 1\right). \tag{A.8}$$

Next, we focus on the PEG ratio: $r^{PEG} = \sqrt{\frac{(E_2 \quad E_1)}{P_0}}$. A corresponding cash-flow news proxy should account for the relative change in the absolute level of earnings in t+2 and the relative change in earnings growth in t. This leads to[31]

$$CNEWS^{PEG,AEGM}_{i,t+1} = \left(\frac{E_{i,t+1,t+1}}{E_{i,t,t+1}} \quad 1\right) + \left(\frac{E_{i,t+1,t+2}}{E_{i,t,t+2}} \quad 1\right) + \left(\frac{gE_{i,t+1,t+3}}{gE_{i,t,t+2}} \quad 1\right). \tag{A.9}$$

where $gE_{i,t,t+1} = \frac{E_{t+1} \quad E_t}{E_t}$ stands for the expected growth in earnings in t. Effectively, $gE_{i,t,t+1}$ varies with t.

Now we turn to the two common present value of expected dividends (PVED) equivalent models i.e., the RIM and AEGM. These models rely on the idea that earnings and their growth are used, in connection with a valuation anchor, to express value and in turn to estimate ICC.[32] We rely on the relative earnings level change and relative earnings growth change for the particular period (as done for the PEG). The period depends on how many years of forecasts the model generally applies. The AEGM (Gode & Mohanram, 2003) applies two years of forecasts, while the RIM (Claus & Thomas, 2001) relies on the longest forecast horizon available (i.e., four years).

Accordingly, the AEGM applies the same cash-flow news proxy as the PEG, i.e., Eq. (A.9), and the RIM uses the following CNEWS proxy:

$$CNEWS^{RIM}_{i,t+1} = \left(\frac{E_{i,t+1,t+1}}{E_{i,t,t+1}} \quad 1\right) + \left(\frac{E_{i,t+1,t+4}}{E_{i,t,t+4}} \quad 1\right) + \left(\frac{gE_{i,t+1,t+5}}{gE_{i,t,t+4}} \quad 1\right). \tag{A.10}$$

The fact that a change in the LT expected cash-flows, compared to a change in the ST expected cash-flows, should have a multiplied impact on the realized return cannot be ignored. EM deal with this by adding a capitalization factor (p) to the LT expected cash-flow changes (see p. 511 in EM).

We address this issue in two ways in our multivariate analysis. First, in line with Botosan et al. (2011), we add each component of the cash-flow news to the regression i.e., for the earnings-to-price ratio there are two cash-flow news components and for the others there are three cash-flow news components in the regression.[33] Second, we also follow EM and use the capitalization factor $\frac{p}{1\ p}$. We add to the short-term cash-flow news the average of the LT cash-flow news multiplied by $\frac{p}{1\ p}$.

Appendix B. Additional analysis and robustness checks

B.1 EM sample sizes and periods

To investigate whether the difference between our results and the EM results are driven by differences in sample sizes and periods, we repeat our analysis presented in Table 5 by using the same sample period as EM. So, we use observations between 1981 and 1998, which results in a total of 17,778 observations. The results reported in the Table below and are not different from those our main analysis (Table 5). Consequently, the different sample and period do not drive our results.

Table B.1
Implied risk premia and realized returns, with sample sizes and periods as in EM

Panel A: Regression results for individual CNEWS as developed in this paper

RP metric	Intercept	RP	CNEWS-components						Mod. Noise Var.
			ST	LT-level	LT-growth	DNEWS	Adj. R^2	max. VIF	
$RP^{E/P}$	0.0824	0.5385	0.0561	0.5424		0.1068	49.76	3.50	0.0007
t-value	(4.05)***	(2.84)***	(3.06)***	(18.75)***	(13.53)***				
RP^{PEG}	0.0489	1.2395	0.0663	0.5237	0.1637	0.0867	40.05	3.78	0.0017
t-value	(1.74)*	(5.39)***	(3.20)***	(16.67)***	(6.88)***	(6.75)***			
RP^{RIM}	0.0557	0.8795	0.0837	0.5586	0.0624	0.1280	44.71	3.78	0.0009

(*continued on next page*)

[31] The notation *PEG, AEGM* here indicates that this cash-flow news proxy is identical in these two ICC models (see discussion below).

[32] Strictly speaking, the *RIM* relies on residual earnings and their growth, while the *AEGM* relies on earnings and their growth. Residual earnings also rely on earnings, since they are calculated as the earnings less a charge (cost of capital) at the beginning-of-year book value. We therefore do not differentiate between *RIM* and *AEGM*.

[33] In fact, the procedure of Botosan et al. (2011) relaxes the assumption in EM on how future cash flows are capitalized into realized returns.

Table B.1 (*continued*)

Panel A: Regression results for individual CNEWS as developed in this paper

RP metric	Intercept	RP	CNEWS-components						Mod. Noise Var.
			ST	LT-level	LT-growth	DNEWS	Adj. R^2	max. VIF	
t-value	(2.67)***	(4.19)***	(3.77)***	(17.18)***	(5.34)***	(13.01)***			
RP^{AEGM}	0.0251	1.1583	0.0652	0.4989	0.1400	0.0656	36.44	3.88	0.0014
t-value	(0.96)	(6.33)***	(2.94)***	(14.69)***	(5.55)***	(7.73)***			

Panel B: Results for individual CNEWS according to Easton and Monahan (2005)

RP metric	Intercept	RP	$CNEWS^{EM}$-components						Mod. Noise Var.
			$(ROE_{i,t}$ $FROE_{i,t,t})$	$(FROE_{i,t+1,t+1}$ $FROE_{i,t,t+1})$	$(FROE_{i,t+1,t+2}$ $FROE_{i,t,t+2})$	DNEWS	Adj. R^2	max. VIF	
$RP^{E/P}$	0.0673	0.2500	0.0305	0.3468	2.6341	0.0970	40.62	11.05	0.0007
t-value	(3.21)***	(1.04)	(-0.84)	(1.11)	(8.76)***	(12.09)***			
RP^{PEG}	0.0705	0.0138	0.0233	0.0596	2.7768	0.06114	28.43	11.95	0.0017
t-value	(2.59)***	(-0.05)	(-0.57)	(0.16)	(7.89)***	(5.52)***			
RP^{RIM}	0.0678	0.1488	0.0413	0.4088	2.6419	0.0893	30.78	11.27	0.0009
t-value	(3.06)***	(0.44)	(-1.09)	(1.16)	(7.48)***	(8.82)***			
RP^{AEGM}	0.0592	0.1775	0.0251	0.1227	2.6766	0.0467	26.38	11.95	0.0014
t-value	(2.27)**	(0.74)	(-0.63)	(0.32)	(7.43)***	(6.90)***			

Panel C: Results for combined CNEWS as developed in this paper

RP metric	Intercept	RP	CNEWS	DNEWS	Adj. R^2	max. VIF	Mod. Noise Var.
$RP^{E/P}$	0.0699	0.1194	0.0127	0.1187	47.86	1.89	0.0007
t-value	(3.41)***	(0.53)	(11.86)***	(15.25)***			
RP^{PEG}	0.0452	0.7174	0.0232	0.2004	46.59	5.37	0.0017
t-value	(1.76)*	(3.50)***	(13.63)***	(15.71)***			
RP^{RIM}	0.0629	0.0095	0.01937	0.1736	40.03	3.21	0.0009
t-value	(3.09)***	(-0.04)	(12.80)***	(18.54)***			
RP^{AEGM}	0.0417	0.5212	0.0208	0.1668	40.16	3.86	0.0014
t-value	(1.66)*	(3.27)***	(11.52)***	(19.49)***			

Panel D: Results for combined CNEWS according to Easton and Monahan (2005)

RP metric	Intercept	RP	$CNEWS^{EM}$	DNEWS	Adj. R^2	max. VIF	Mod. Noise Var.
$RP^{E/P}$	0.0599	0.3563	0.2126	0.0862	21.59	1.23	0.0007
t-value	(2.96)***	(-1.21)	(3.89)***	(10.29)***			
RP^{PEG}	0.0712	0.9074	0.1972	0.0505	11.14	1.39	0.0017
t-value	(2.63)***	(-2.78)***	(3.82)***	(6.00)***			
RP^{RIM}	0.0814	0.9556	0.2115	0.0675	12.12	1.34	0.0009
t-value	(3.68)***	(-2.51)**	(3.80)***	(5.91)***			
RP^{AEGM}	0.0792	0.6225	0.2000	0.0415	9.27	1.32	0.0014
t-value	(2.94)***	(-2.40)**	(3.85)***	(6.17)***			

This table provides results of regressions of realized returns on expected returns as well as cash-flow and return news. The coefficient and the adjusted R^2 are the mean of annual values, and the statistic is according to Fama and MacBeth (1973). Panels A and B provide results with individual cash-flow news proxies and Panels C and D provide results with the summarized cash-flow news proxies (see Appendix A.2 for more details). The four implied risk premia estimates, i.e., implied cost of capital less the prevailing risk-free rate, $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, are estimated as outlined in Appendix A. RET_1 is the one-year-ahead buy-and-hold return following the calculation of the implied risk premia, less the risk-free rate. $CNEWS^{EM}$ is the cash-flow news as outlined in Equation (2) (according to Easton & Monahan, 2005). $CNEWS^{E/P}_{i,t+1}$ ($CNEWS^{PEG,AEGM}_{i,t+1}$) is the approach-specific cash-flow news proxy for the $RP^{E/P}$ (RP^{PEG} and RP^{AEGM}). $CNEWS^{RIM}_{i,t+1}$ is the approach-specific cash-flow news proxy for the RP^{RIM}. ST, LT-level, and LT-growth stand for short-term, long-term, and the long-term growth components of the cash-flow news proxy, respectively. The developed CNEWS are calculated as outlined inAppendix A.2. $DNEWS^{z}_{i,t+1}$ stands for the approach-specific discount-rate news proxy as outlined in Section 4.1 (according to Easton & Monahan, 2005). z stands for the specific implied risk premia approach, i.e., $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, respectively. All variables are measured as the natural log of one plus the respective variable according to Easton and Monahan (2005). Max. VIF is the maximum value of variance inflation factor across the annual regressions. Mod. Noise Var. is the modified noise variable, measured as in Easton and Monahan (2005) (see footnote 7). As EM point out (see details on page 507), the modified noise variable changes with different RP metrics as expected, but not with different cash-flow news estimation. T-value is the test of difference from zero. ***, **, and * indicate significance at the 1%, 5%, and 10% level, respectively. The sample size is 17,778 observations for years 1981–1998.

B.2 Further tests to check for multicollinearity

To further investigate potential multicollinearity for our developed cash-flow news proxies, we apply a procedure in line with Yu (2008). In the Table below (Panel A), we regress short-term on long-term (level) cash-flow news and substitute the long-term (level) cash-flow news with the residuals from that regression. The results are virtually identical to those in Panel A of Table 5, while the VIFs drop to values below 2.5. In Panel B of the Table below, we regress discount-rate news on cash-flow news and substitute the cash-flow news with the residuals from that regression. Again, results are virtually identical to those in Panel C of Table 5 and the VIFs drop below 1.5.

As a test to alleviate any potential concerns of mechanical relation between our cash-flow news and return news proxies, which may cause the positive coefficients of RP, we substitute the firm-specific return news with an average return news based on the average of the four ICC proxies. This consistent with one of Botosan et al.'s (2011) sensitivity tests. The results are reported in Panels C and D of the Table below and are comparable to those in Panels A and C of Table 5.

Table B.2
Implied risk premia and realized returns, after controlling for potential multicollinearity

Panel A: Regression results for individual CNEWS as developed in this paper

RP metric	Intercept	RP	CNEWS-components			DNEWS	Adj. R^2	max. VIF
			ST	Residuals	LT-growth			
$RP^{E/P}$	0.0606	0.6003	0.3238	0.5292		0.1098	48.08	1.50
t-value	(2.65)***	(3.68)***	(22.75)***	(26.63)***		(18.56)***		
RP^{PEG}	0.0105	1.3318	0.3267	0.5147	0.1706	0.0819	38.72	2.47
t-value	(0.43)	(7.29)***	(20.98)***	(10.75)***	(24.60)***	(10.50)***		
RP^{RIM}	0.0283	1.1282	0.3362	0.5570	0.0503	0.1306	45.80	1.84
t-value	(1.27)	(5.93)***	(19.69)***	(6.76)***	(24.13)***	(19.85)***		
RP^{AEGM}	0.0072	1.2262	0.3144	0.4971	0.1490	0.0677	36.22	2.24
t-value	(-0.31)	(7.64)***	(20.12)***	(9.34)***	(22.87)***	(11.75)***		

Panel B: Results for combined CNEWS as developed in this paper

RP metric	Intercept	RP	Residuals	DNEWS	Adj. R^2	max. VIF
$RP^{E/P}$	0.0325	0.1321	0.0102	0.0862	44.73	1.43
t-value	(1.33)	(0.77)	(13.99)***	(14.83)***		
RP^{PEG}	0.0068	0.6558	0.0187	0.0393	42.76	1.46
t-value	(-0.24)	(4.27)***	(14.71)***	(7.45)***		
RP^{RIM}	0.0261	0.0863	0.0146	0.0683	37.13	1.37
t-value	(1.07)	(0.39)	(11.54)***	(9.30)***		
RP^{AEGM}	0.0033	0.4458	0.0167	0.0368	37.57	1.43
t-value	(-0.13)	(3.13)***	(13.22)***	(8.13)***		

Panel C: Regression results for individual CNEWS as developed in this paper

RP metric	Intercept	RP	CNEWS-components			DNEWS-Avg.	Adj. R^2	max. VIF
			ST	LT-level	LT-growth			
$RP^{E/P}$	0.0544	0.8340	0.0054	0.5321		0.0877	39.97	3.56
t-value	(2.66)***	(4.56)***	(0.42)	(22.32)***		(15.73)***		
RP^{PEG}	0.0283	1.3599	0.0380	0.5444	0.1456	0.1110	44.24	3.81
t-value	(1.34)	(8.02)***	(3.20)***	(25.80)***	(12.30)***	(16.68)***		
RP^{RIM}	0.0141	1.4404	0.0395	0.5378	0.0244	0.1049	43.45	3.74
t-value	(0.70)	(7.41)***	(3.63)***	(23.07)***	(3.21)***	(17.28)***		
RP^{AEGM}	0.0134	1.2199	0.0351	0.5407	0.1403	0.1107	43.93	3.92
t-value	(0.67)	(8.41)***	(3.00)***	(26.05)***	(11.19)***	(17.64)***		

Panel D: Results for combined CNEWS as developed in this paper

RP metric	Intercept	RP	CNEWS	DNEWS-Avg.	Adj. R^2	max. VIF
$RP^{E/P}$	0.04721	0.4265	0.0099	0.0984	36.62	2.06
t-value	(2.20)**	(2.11)**	(11.16)***	(15.68)***		
RP^{PEG}	0.0278	0.7505	0.0146	0.1767	38.96	3.47
t-value	(1.35)	(4.86)***	(10.94)**	(23.97)***		
RP^{RIM}	0.0294	0.5354	0.0121	0.1146	32.67	2.72
t-value	(1.43)	(2.44)**	(9.54)***	(12.90)***		
RP^{AEGM}	0.0241	0.6193	0.0145	0.1762	38.78	3.44
t-value	(1.24)	(4.31)***	(10.50)***	(24.41)***		

This table provides results on additional tests of regressions of realized returns on expected returns as well as cash flow and return news. Panel A substitutes residuals from a regression of the short-term cash-flow news on the long-term (level) cash-flow news for the long-term (level) cash-flow news. Panel B substitutes residuals from a regression of cash-flow news on discount-rate news for the cash-flow news. Panels C and D use an average of the four measures discount-rate news as a proxy for the discount-rate news. The coefficient

and the adjusted R^2 are the mean of annual values and the statistic is according to Fama and MacBeth (1973). The four implied risk premia estimates, i.e., implied cost of capital less the prevailing risk-free rate, $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, are estimated as outlined in Appendix A. RET_1 is the one-year-ahead buy-and-hold return following the calculation of the implied risk premia, less the risk-free rate. $CNEWS^{EM}$ is the cash-flow news as outlined in Equation (2) (according to Easton & Monahan, 2005). $CNEWS^{E/P}_{i,t+1}$ ($CNEWS^{PEG,AEGM}_{i,t+1}$) is the approach-specific cash-flow news proxy for the $RP^{E/P}$ (RP^{PEG} and RP^{AEGM}). $CNEWS^{RIM}_{i,t+1}$ is the approach-specific cash-flow news proxy for the RP^{RIM}. ST, LT-level, and LT-growth stand for short-term, long-term, and the long-term growth components of the cash-flow news proxy, respectively. The developed CNEWS are calculated as outlined in Appendix A.2. *DNEWS Avg.* stands for the average of the four approach-specific discount-rate news proxies as outlined in Section 4.1 (according to Easton & Monahan, 2005). All variables are measured as the natural log of one plus the respective variable according to Easton and Monahan (2005). Max. VIF is the maximum value of variance inflation factor across the annual regressions. T-value is the test of difference from zero. ***, **, and * indicate significance at the 1%, 5%, and 10% level, respectively. The sample size is 43,465 observations.

B.3 Focusing on cash-flow news only

Vuolteenaho (2002) demonstrates that cash-flow news is largely firm-specific, while return news is primarily driven by macroeconomic factors. Accordingly, recent research relies only on cash-flow news when considering firm-level realized returns (e.g., Ogneva, 2012). To provide further robustness, we repeat the analysis reported in Table 5 without considering return news. Panels A and C in the Table below report these annual regressions using our developed cash-flow news and Panels B and D report the cash-flow news proxy of EM.

Table B.3
Implied risk premia and realized returns, with cash-flow news only

Panel A: Regression results for individual CNEWS as developed in this paper

RP metric	Intercept	RP	CNEWS-components			Adj. R^2	max. VIF
			ST	LT-level	LT-growth		
$RP^{E/P}$	0.0246	1.4680	0.0510	0.4444		27.95	3.42
t-value	(0.98)	(6.42)***	(3.50)***	(20.40)***			
RP^{PEG}	0.0046	1.6639	0.0580	0.4649	0.0613	28.15	3.78
t-value	(-0.18)	(7.21)***	(3.71)***	(20.16)***	(6.46)***		
RP^{RIM}	0.0256	2.1398	0.1060	0.3909	0.0268	27.49	3.74
t-value	(-1.01)	(8.28)***	(7.01)***	(18.29)***	(3.90)***		
RP^{AEGM}	0.0336	1.6977	0.0532	0.4661	0.0582	27.77	3.89
t-value	(-1.35)	(8.72)***	(3.52)***	(20.49)***	(5.60)***		

Panel B: Results for individual CNEWS according to Easton and Monahan (2005)

RP metric	Intercept	RP	$CNEWS^{EM}$-components					Adj. R^2	max. VIF
			($ROE_{i,t}$	$FROE_{i,t,t}$)	($FROE_{i,t+1,t+1}$	$FROE_{i,t,t+1}$)	($FROE_{i,t+1,t+2}$	$FROE_{i,t,t+2}$)	
$RP^{E/P}$	0.0124	1.1617	0.0307		0.3005		2.1943		21.29	11.04
t-value	(0.48)	(4.77)***	(-1.50)		(1.39)		(11.62)***			
RP^{PEG}	0.0164	0.7272	0.0338		0.1509		2.3214		19.82	11.86
t-value	(0.67)	(3.48)***	(-1.56)		(0.62)		(10.80)***			
RP^{RIM}	0.0017	1.2358	0.0337		0.3088		2.1963		19.96	11.21
t-value	(-0.06)	(4.68)***	(-1.69)*		(1.36)		(10.90)***			
RP^{AEGM}	0.0055	0.8789	0.0409		0.1054		2.3660		19.67	11.85
t-value	(-0.22)	(4.85)***	(-1.65)*		(0.44)		(11.26)***			

Panel C: Results for combined CNEWS as developed in this paper

RP metric	Intercept	RP	CNEWS	Adj. R^2	max. VIF
$RP^{E/P}$	0.0157	1.1382	0.0084	22.65	1.08
t-value	(0.61)	(4.49)***	(12.32)***		
RP^{PEG}	0.0134	1.0476	0.0063	9.10	1.43
t-value	(-0.53)	(4.37)***	(6.78)***		
RP^{RIM}	0.0181	1.4061	0.0089	15.26	1.18
t-value	(-0.71)	(4.77)***	(10.66)***		
RP^{AEGM}	0.0365	1.1755	0.0063	8.92	1.37
t-value	(-1.40)	(5.78)***	(6.51)***		

Panel D: Results for combined CNEWS according to Easton and Monahan (2005)

RP metric	Intercept	RP	$CNEWS^{EM}$	Adj. R^2	max. VIF
$RP^{E/P}$	0.0088	0.7132	0.2263	7.20	1.05
t-value	(0.34)	(2.68)***	(7.13)***		
RP^{PEG}	0.0259	0.0081	0.2162	5.70	1.08

(*continued on next page*)

Table B.3 (*continued*)

Panel D: Results for combined CNEWS according to Easton and Monahan (2005)					
RP metric	Intercept	RP	$CNEWS^{EM}$	Adj. R^2	max. VIF
t-value	(1.09)	(0.04)	(7.05)***		
RP^{RIM}	0.0122	0.3947	0.2209	5.90	1.06
t-value	(0.48)	(1.36)	(7.05)***		
RP^{AEGM}	0.0183	0.2027	0.2178	5.32	1.05
t-value	(0.74)	(1.13)	(7.00)***		

This table provides results of regressions of realized returns on expected returns as well as cash-flow news. Panels A and B provide results with individual cash-flow news proxies and Panels C and D provide results with the summarized cash-flow news proxies (see Appendix A.2 for more details). The coefficient and the adjusted R^2 are the mean of annual values and the statistic is according to Fama and MacBeth (1973). The four implied risk premia estimates, i.e., implied cost of capital less the prevailing risk-free rate, $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, are estimated as outlined in Appendix A. RET_1 is the one-year-ahead buy-and-hold return following the calculation of the implied risk premia, less the risk-free rate. $CNEWS^{EM}$ is the cash-flow news as outlined in Equation (2) (according to Easton & Monahan, 2005). $CNEWS^{E/P}_{i,t+1}$ ($CNEWS^{PEG,AEGM}_{i,t+1}$) is the approach-specific cash-flow news proxy for the $RP^{E/P}$ (RP^{PEG} and RP^{AEGM}). $CNEWS^{RIM}_{i,t+1}$ is the approach-specific cash-flow news proxy for the RP^{RIM}. ST, LT-level, and LT-growth stand for short-term, long-term, and the long-term growth components of the cash-flow news proxy, respectively. The developed CNEWS are calculated as outlined in Appendix A.2. All variables are measured as the natural log of one plus the respective variable according to Easton and Monahan (2005). Max. VIF is the maximum value of variance inflation factor across the annual regressions. T-value is the test of difference from zero. ***, **, and * indicate significance at the 1%, 5%, and 10% level, respectively. The sample size is 43,465 observations.

Panels A and C confirm the results of our main analyses. Panel A shows that the coefficients on the risk premia estimates are positive and significant for all ICC estimates applied. Additionally, the coefficients for all cash-flow news proxies are highly significantly positive. In comparison to the results in Table 5, which additionally account for return news, the average adjusted R^2 are substantially smaller across all ICC approaches (e.g., 45.69 for the RP^{PEG} in Panel A of Table 5 vs. 28.15 in Panel A of this Table). This indicates that neglecting discount-rate news leads to an omitted variable bias, because the discount-rate news is able to explain a substantial part of the variation. The results in Panel C, which are based on the summed cash-flow news, confirm those in Panel A. The results in Panels B and D using the EM cash-flow news proxy show positive signs (all are significant in Panel B and one value is significant in Panel D) for the expected return proxies. However, the reported adjusted R^2 in Panel B (D) are substantially lower than those in Panel A (C), which again indicates that our cash-flow news proxies outperform those of EM.

B.4 Discount-rate news as in Botosan et al. (2011)

We repeat our analysis by using the Botosan et al. (2011) definition of return news. We use the change in market beta as the firm-specific return news proxy.[34] In contrast to Botosan et al. (2011), we do not include the change in the risk-free rate, given that we use firms with December year-end only, and the change in the risk-free rate is thus a cross-sectional constant (see also the discussion of EM (2016)). The results of this analysis are presented in the Table below.

Table B.4
Implied risk premia and realized returns, with discount-rate news as in Botosan et al. (2011)

Panel A: Regression results for individual CNEWS as developed in this paper								
RP metric	Intercept	RP	CNEWS-components			DNEWS-β	Adj. R^2	max. VIF
			ST	LT-level	LT-growth			
$RP^{E/P}$	0.0274	1.3636	0.0543	0.4225		0.0024	28.94	3.39
t-value	(1.05)	(6.22)***	(3.42)***	(21.78)***		(-0.07)		
RP^{PEG}	0.0002	1.6054	0.0610	0.4435	0.0584	0.0138	29.30	3.97
t-value	(0.01)	(7.35)***	(3.63)***	(20.66)***	(6.45)***	(-0.39)		
RP^{RIM}	0.0211	2.0472	0.1070	0.3733	0.0260	0.0017	28.62	3.73
t-value	(-0.81)	(8.37)***	(6.77)***	(18.90)***	(3.70)***	(-0.05)		
RP^{AEGM}	0.0284	1.6239	0.0564	0.4444	0.0555	0.0048	29.00	4.10
t-value	(-1.10)	(8.60)***	(3.47)***	(21.04)***	(5.51)***	(-0.14)		
Panel B: Results for individual CNEWS according to Easton and Monahan (2005)								
RP metric	Intercept	RP	$CNEWS^{EM}$-components			DNEWS-ß	Adj. R^2	max. VIF

(*continued on next page*)

[34] We take the natural logarithm of one plus market beta at t+1 less the market beta at t. To expect a positive sign on the coefficient, we multiply the result by minus one. We determine market beta using the market model and require at least 12 and up to 60 months of lagged monthly returns. Due to this additional data constraint, our sample size decreases slightly, to a total of 41,305 observations.

Table B.4 (*continued*)

Panel B: Results for individual CNEWS according to Easton and Monahan (2005)

RP metric	Intercept	RP	$CNEWS^{EM}$-components						DNEWS-ß	Adj. R^2	max. VIF
			$(ROE_{i,t}$	$FROE_{i,t,t})$	$(FROE_{i,t+1,t+1}$	$FROE_{i,t,t+1})$	$(FROE_{i,t+1,t+2}$	$FROE_{i,t,t+2})$			
			$(ROE_{i,t}$	$FROE_{i,t,t})$	$(FROE_{i,t+1,t+1}$	$FROE_{i,t,t+1})$	$(FROE_{i,t+1,t+2}$	$FROE_{i,t,t+2})$			
$RP^{E/P}$	0.0164	1.0475	0.03155		0.3966		2.0253		0.0025	22.54	10.79
t-value	(0.69)	(4.41)***	(-1.57)		(1.85)*		(11.13)***		(0.07)		
RP^{PEG}	0.0209	0.6825	0.0332		0.2515		2.1535		0.0090	21.26	11.62
t-value	(0.82)	(3.41)***	(-1.58)		(1.05)		(10.39)***		(-0.23)		
RP^{RIM}	0.0031	1.1542	0.0323		0.3946		2.0422		0.0004	21.36	10.79
t-value	(0.11)	(4.62)***	(-1.68)*		(1.78)*		(10.54)***		(-0.01)		
RP^{AEGM}	0.0004	0.8213	0.0409		0.2105		2.1945		0.0022	21.16	11.60
t-value	(-0.02)	(4.68)***	(-1.68)*		(0.90)		(10.82)***		(-0.06)		

Panel C: Results for combined CNEWS as developed in this paper

RP metric	Intercept	RP	CNEWS	DNEWS-β	Adj. R^2	max. VIF
$RP^{E/P}$	0.0188	1.0284	0.0082	0.0084	23.48	1.13
t-value	(0.71)	(4.20)***	(12.32)***	(-0.25)		
RP^{PEG}	0.0077	0.9719	0.0060	0.1708	10.96	1.39
t-value	(-0.29)	(4.38)***	(6.59)***	(-0.40)		
RP^{RIM}	0.0133	1.3042	0.0087	0.0103	16.63	1.19
t-value	(-0.50)	(4.84)***	(10.61)***	(-0.26)		
RP^{AEGM}	0.0299	1.0850	0.0060	0.0107	10.82	1.32
t-value	(-1.09)	(5.70)***	(6.36)***	(-0.25)		

Panel D: Results for combined CNEWS according to Easton and Monahan (2005)

RP metric	Intercept	RP	$CNEWS^{EM}$	DNEWS-ß	Adj. R^2	max. VIF
$RP^{E/P}$	0.0136	0.6940	0.2578	0.0101	10.43	1.09
t-value	(0.49)	(2.74)***	(8.54)***	(-0.25)		
RP^{PEG}	0.0292	0.0904	0.2494	0.0167	9.07	1.09
t-value	(1.15)	(0.44)	(8.52)***	(-0.39)		
RP^{RIM}	0.0138	0.4864	0.2539	0.0112	9.20	1.11
t-value	(0.52)	(1.82)*	(8.50)***	(-0.27)		
RP^{AEGM}	0.0200	0.2576	0.2516	0.0136	8.84	1.08
t-value	(0.77)	(1.52)	(8.47)***	(-0.32)		

This table provides results of regressions of realized returns on expected returns as well as cash flow and return news. Panels A and B provide results with individual cash-flow news proxies and Panels C and D provide results with the summarized cash-flow news proxies (see Appendix A.2 for more details). The coefficient and the adjusted R^2 are the mean of annual values and the statistic is according to Fama and MacBeth (1973). The four implied risk premia estimates, i.e., implied cost of capital less the prevailing risk-free rate, $RP^{E/P}$, RP^{PEG}, RP^{RIM}, and RP^{AEGM}, are estimated as outlined in Appendix A. RET_1 is the one-year-ahead buy-and-hold return following the calculation of the implied risk premia, less the risk-free rate. $CNEWS^{EM}$ is the cash-flow news as outlined in Equation (2) (according to Easton & Monahan, 2005). $CNEWS^{E/P}_{i,t+1}$ ($CNEWS^{PEG,AEGM}_{i,t+1}$) is the approach-specific cash-flow news proxy for the $RP^{E/P}$ (RP^{PEG} and RP^{AEGM}). $CNEWS^{RIM}_{i,t+1}$ is the approach-specific cash-flow news proxy for the RP^{RIM}. ST, LT-level, and LT-growth stand for short-term, long-term, and the long-term growth components of the cash-flow news proxy, respectively. The developed CNEWS are calculated as outlined in Appendix A.2. *DNEWS*- β stands for the firm-specific discount-rate news according to Botosan et al. (2011). In particular, we use the change in market beta as the firm-specific return news proxy. We determine market beta by using the market model and requiring at least 12 and up to 60 months of lagged monthly returns. All variables are measured as the natural log of one plus the respective variable, according to Easton and Monahan (2005). Max. VIF is the maximum value of variance inflation factor across the annual regressions. T-value is the test of difference from zero. ***, **, and * indicate significance at the 1%, 5%, and 10% level, respectively. The sample size is 41,305 observations, i.e., 43,465 observations less observations for which data to calculate the discount-rate news according to Botosan et al. (2011) is not available.

Panel A of this Table shows the results using the individual cash-flow news proxies and Panel C those with the summed cash-flow news proxies. These results reveal that the return news proxies show a negative sign, insignificant across all assessed ICC estimates. This is consistent with Botosan et al. (2011) (Table 6 in their paper), who also report a contradictory, insignificant sign for the discount-rate news based on market beta. Thus, it is of no surprise that the results regarding the relation between the expected return proxies (ICC approaches) and the cash-flow news proxies are almost identical to those in Panels A and C of the previous Table in this Appendix.

Moreover, Panels B and D show the corresponding results using the EM cash-flow news proxy (as also performed by Botosan et al., 2011). Again, we report an insignificant sign for the discount-rate news of Botosan et al. (2011) and the relation between the expected return proxies and the cash-flow news proxies is almost identical to those reported in Panels B and D of the previous Table in this Appendix (and in Table 10 in Botosan et al., 2011).

In addition, the adjusted R^2 in this Table are much smaller than those in Table 5, when we apply EM's proxy for discount-rate news.

As suggested by EM (2016), the return news proxy employed by EM outperforms that of Botosan et al. (2011). Moreover, while the inferences drawn from the Botosan et al. (2011) approach and the approach used in Table 5 (our consistent cash-flow and return news) are almost identical, this might not be the case in other settings/for other samples (countries, time periods, etc.), given the data constraints imposed by using ValueLine target prices.[35]

Appendix C. Removing predictable forecast errors

To adjust for analysts' optimism, we rely on Mohanram and Gode (2013) and Larocque (2013) and use the following variables to predict forecast error: long-term growth, lagged returns, and revision in analysts' forecasts.[36] To ensure that the models' estimated coefficients are not biased by measurement error, in line with Larocque (2013),[37] we also control for abnormal stock return between the forecast date and the year t+1 (t+2) end date. For predicting forecast errors, the abnormal stock return is not included in the model because it is not available at the time the forecasts are collected.

We run the following two regressions for one- and two-year ahead forecast errors:

$$ERROR_{i,t+1} = \alpha_1 + \alpha_2 LTG_{i,t} + \alpha_3 REV_{i,t} + \alpha_4 RET_LAG_{i,t} + \alpha_5 RET_EZ_{i,t+1} + \varepsilon_t. \quad (C.1)$$

$$ERROR_{i,t+2} = \alpha_1 + \alpha_2 LTG_{i,t} + \alpha_3 REV_{i,t} + \alpha_4 RET_LAG_{i,t} + \alpha_5 RET_EZ_{i,t+2} + \varepsilon_t. \quad (C.2)$$

Where:

$ERROR_{i,t+1} = \frac{F_EPS_{i,t+1,t} - F_EPS_{i,t+1,t+1}}{Price_t}$, $F_EPS_{i,t+1,t}$ is forecasted earnings per share for year t+1 estimated at time t, and $F_EPS_{i,t+1,t+1}$ is forecasted earnings per share for year t+1 at time t+1. We calculate forecast error based on two forecasts to avoid look-ahead bias. Because we collect our forecast data on the last Thursday in December, the realized earnings for year t+1 are not observable at that time. Accordingly, we use the forecasts of earnings for year t+1 (just a few days before fiscal year-end) (Easton, 2004). We repeat the analysis with realized earnings per share (which imposes a look-ahead bias) and the results are virtually identical.

$ERROR_{i,t+2} = \frac{F_EPS_{i,t+2,t} - F_EPS_{i,t+2,t+2}}{Price_t}$ and $LTG_{i,t}$ is the median long-term growth rate from I/B/E/S at time t.

$REV_{i,t}$ is the revision in analysts' forecasts between the six months prior to the collection of analysts' forecasts and the date of the collection of analysts' forecasts scaled by price. We also use a three-month period with virtually identical results.

$RET_LAG_{i,t}$ is the total return to shareholders minus the value-weighted return on a market portfolio from the 12 months prior to the collection of forecasts at time t.

$RET_EZ_{i,t+1}$ is the total return to shareholders minus the value-weighted return on a market portfolio from the date of the collection of the forecasts at time t (last Thursday of December in t) to the date of the collection of forecasts at time t+1 (last Thursday of December in t+1).

$RET_EZ_{i,t+2}$ is the total return to shareholders minus the value-weighted return on a market portfolio from the date of the collection of the forecasts at time t (last Thursday of December in t) to the date of the collection of forecasts at time t+2 (last Thursday of December in t+2).

We estimate the following coefficients from C1 and C2. The table below presents the average coefficients and the corresponding t-statistic:

Variable	Intercept	$LTG_{i,t}$	$REV_{i,t}$	$RET_LAG_{i,t}$	$RET_EZ_{i,t+1}$	Adj. R^2
$ERROR_{i,t+1}$	0.0046	0.0191	0.4631	0.0174	0.0481	25.32
(t-value)	(1.58)	(2.86)***	(-13.19)***	(-6.09)***	(-11.55)***	
	Intercept	$LTG_{i,t}$	$REV_{i,t}$	$RET_LAG_{i,t}$	$RET_EZ_{i,t+2}$	
$ERROR_{i,t+2}$	0.0120	0.0436	0.4570	0.0225	0.0526	29.56
(t-value)	(3.35)***	(4.17)***	(-8.68)***	(-6.64)***	(-14.44)***	

The signs and the significance of the coefficients of LTG, REV, and RET_LAG are identical to those reported by Mohanram and Gode (2013). The coefficients in our analysis are slightly smaller, which is due to the inclusion of RET_EZ in our regressions. This also explains why our adjusted R^2 are considerably higher than 19.8% and 16.1% reported by Mohanram and Gode (2013), respectively.

[35] Botosan et al. (2011) report a sample size of 14,521 observations between 1984 and 2004. Our I/B/E/S based sample size, if we were also focusing on observations up to 2004, is 27,332 i.e., an increase almost 90%.

[36] Mohanram and Gode (2013) also include total accruals scaled by total assets, sales growth, growth in PPE, and growth in other long-term assets in the model. However, the coefficients for these variables are insignificant in their empirical analysis. Based on this, we exclude these variables. In robustness checks, we include these variables and also observe insignificant coefficients for these variables, although, due to the additional data needed, the sample size decreases.

[37] Larocque (2013) also uses lagged forecasts error in their model. We repeat the regressions with lagged forecast error. Our results show that lagged error is insignificant. Only when we remove REV (which is not used in Larocque, 2013) does lagged error become highly significant in the expected way.

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