

Accruals Quality, Information Risk and Cost of Capital: Evidence from Australia

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Abstract: Recent theoretical work argues that information risk is a non-diversifiable risk factor that is priced in the capital market. Using accruals quality to proxy for information risk, Francis et al. (2005) provide empirical support for this argument using a sample of US firms. This paper re-examines the interplay of accruals quality, information risk and cost of capital in Australia, where a number of important institutional and regulatory differences are hypothesized to affect the relation between accruals quality and cost of capital. The results suggest that, while accruals quality impacts on the cost of capital for Australian firms, some salient differences exist. In contrast to findings for US firms, the costs of debt and equity for Australian firms are largely influenced by accruals quality arising from economic fundamentals (i.e., innate accrual quality) but not discretionary reporting choices (i.e., discretionary accrual quality). This finding is consistent with our predictions based on the Australian institutional and regulatory environment. In addition, using both the asset pricing tests in Francis et al. (2005) and Core et al. (2008), we provide evidence consistent with accruals quality being a priced risk factor.

Keywords: accruals quality, information risk, cost of capital, information asymmetry, information precision, discretionary, innate

1. INTRODUCTION

Recent theoretical work posits that information risk is a non-diversifiable risk factor that is priced by the capital market. Several explanations have been given for the cause of information risk. In a multi-asset rational expectations framework, Easley and O'Hara (2004, hereafter EOH) analyze the role of information asymmetry among investors in the determination of cost of capital. In their model, less-informed investors recognize their informational disadvantage to more-informed investors. Accordingly, they demand a return premium for firms with a higher degree of information asymmetry (i.e., a higher level of information risk). In contrast, Lambert et al. (2007)

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(hereafter LLV) argue that, in models with perfect competition, it is the precision of information (rather than information asymmetry per se) that is the key determinant of information risk which affects cost of capital. LLV (2007) define information precision as the average quality of information that investors have on the expected cash flows of the firm, and information asymmetry as the difference in precision across investors. As such, LLV argue that it is not the distribution of information across investors that matters, but more so how precise that information is.¹

Despite their different perspectives, the models of EOH (2004) and LLV (2007) share some common ground. EOH (2004) recognize an important role for information precision in reducing the cost of capital by mitigating the systematic risk arising from information asymmetry across informed and uninformed investors. LLV (2007) acknowledge a cost of capital role for information asymmetry when competition is imperfect. Irrespective of the source of information risk, the theoretical models of both EOH (2004) and LLV (2007) predict that equilibrium asset prices are influenced by information risk and that information risk may contribute to cross-sectional differences in firms' required return.²

Francis et al. (2005) (hereafter FLOS) provide empirical support for the association between information risk and the cost of capital. Using the quality of accruals as a proxy for information risk, FLOS (2005) report that US firms with poorer accruals quality (AQ) exhibit higher costs of debt and equity capital than firms with better AQ. Further, FLOS (2005) partition AQ into components that reflect economic fundamentals (innate AQ) and managerial reporting choices (discretionary AQ) and show that, while both components are associated with cost of capital, the former has the greater effect.

From an asset-pricing perspective, FLOS (2005) argue that information risk is a priced risk factor. Using the time-series regression approach of Fama and French (1993), they document a significant positive loading on an AQ-based factor-mimicking portfolio. Recently, however, Core et al. (2008) (hereafter CGV) suggest that a better-specified test of whether a proposed risk factor is priced requires a two-stage cross-sectional regression (2SCSR) method. Using the 2SCSR approach, CGV find no evidence that AQ is a priced risk factor. Nonetheless, subsequent empirical work by Kim and Qi (2008) and Ogenva (2008) attribute CGV's lack of findings to the impact of stocks with low share price and negative cash flow shocks respectively.³ After controlling

1 In a recent study, Bhattacharya et al. (2008) provide evidence consistent with both a direct path from earnings quality (which proxies for information precision) to the cost of capital, and an indirect path that is mediated by information asymmetry, with the direct path as the dominant factor. Their evidence is thus consistent with both the EOH's (2004) and LLV's (2007) theoretical models on the association between information risk and cost of capital.

2 Unlike EOH (2004) and LLV (2007), Hughes et al. (2007) do not find theoretical support for cross-sectional effect of information asymmetry on cost of capital. Hughes et al. (2007) demonstrate analytically that in large economies characterized by large number of risky assets and related private signals, private information can affect market-wide factor risk premium but does not affect cost of capital in the cross-section. However, Hughes et al. (2007) point out that their analytical model is silent on the choice of systematic risk factors. Therefore, their theory is not inconsistent with studies that presume an existence of a systematic 'information risk' factor (e.g., FLOS, 2005; and Aboody et al., 2005).

3 Kim and Qi (2008) find the AQ risk factor is statistically and economically priced in the US after the exclusion of low-priced stocks. Further, they show that the AQ risk premium is associated with fundamental risks related to macroeconomic conditions and firms' economic activities. Ogenva (2008) hypothesizes that CGV's (2008) result arises because poor accrual quality firms experience negative cash flow shocks in the future, which results in negative returns that offset the higher expected returns for such firms. Using the 2SCSR approach, she finds AQ is priced after controlling for the effects of cash flow shocks.

for these characteristics, Kim and Qi's (2008) and Ogneva's (2008) results support the notion of FLOS (2005) that AQ is priced risk factor.

The theoretical work of EOH (2004) and LLV (2007), along with the empirical findings of FLOS (2005), CGV (2008) and others, motivate the current study which examines the interplay of AQ, information risk and cost of capital for Australian firms. The Australian regulatory and institutional environment provides an interesting setting in which to further explore the market pricing of AQ for two important reasons. First, relative to US firms, Australian firms are significantly more reliant on private debt compared to public debt.⁴ Private lenders typically have more privileged access to the financial and business information of the borrowing firm than do public debt lenders. Hence, the level of information asymmetry across debt holders is likely to be lower in Australia compared to the US. In addition, private lenders are more likely than public lenders to perform a monitoring role through their close relations with borrowing firms, thereby mitigating managerial opportunism in financial reporting. The more privileged access to information and closer monitoring by private lenders increase information precision and mitigate information asymmetry. Therefore, the information risk associated with discretionary reporting by managers is likely to be reduced, thus diminishing the effect of discretionary AQ on the cost of debt. Accordingly, while FLOS (2005) report that both innate and discretionary AQ significantly affect the cost of debt, we expect that innate AQ is likely to dominate the AQ effect on Australian firms' cost of debt while discretionary AQ is likely to have a negligible effect.

Second, over the period examined in this paper, Australian firms are subject to a continuous disclosure regime (CDR) designed to increase the quality and timeliness of corporate disclosure to the public. The CDR requires listed firms to immediately disclose price-sensitive information to the public via the Australian Stock Exchange when firms become aware of the information. Selective disclosure to third parties such as analysts is strictly prohibited.⁵ A regulatory environment characterized by non-selective disclosure of timely and high-quality information to capital markets is likely to mitigate the opportunities and incentives for managerial discretion in financial reporting. This serves to reduce information asymmetry across investors and increase the average precision of information on the firms' expected cash flows. This reduction in overall information risk provides a second reason why the relation between discretionary AQ and cost of capital for Australian firms may differ from that previously reported for US firms by FLOS (2005).

The results of this paper suggest that AQ is priced by both the debt and equity markets in Australia. While, in general, we document similar relations to those reported by FLOS (2005) for US firms, salient differences exist. Specifically, we report that total AQ is not associated with cost of debt. However, when total AQ is partitioned into innate and discretionary AQ components, cost of debt is significantly influenced

4 Between 1998 and 2006, public debt ranges between 7.8% and 12.3% of total non-government debt securities issued in Australia (statistics computed based on data available on Reserve Bank of Australia's website, <http://www.rba.gov.au/Statistics/Bulletin/index.html>). Similarly, Cotter (1998) documents that public debt issues make up less than 5% of total debt outstanding in her sample of Australian firms.

5 In the US, Regulation Fair Disclosure (RegFD) focuses on the fair access to information by market participants. That is, if a firm chooses to disclose pertinent information, such information must be disclosed to all parties and selective disclosure is not allowed. Unlike RegFD, the Australian CDR not only prohibits selective disclosure, but also requires firms to disclose *all* price-sensitive information once such information is known – non-disclosure of such information is not an option.

by a firm's innate AQ. This lack of relation between discretionary AQ and cost of debt is consistent with our argument that the heavy reliance of Australian firms on private debt reduces information risk related to discretionary AQ, and consequently the influence of discretionary AQ on cost of debt. Similar to US findings with respect to cost of equity, we provide evidence that total AQ is significantly related to the cost of equity. However, unlike FLOS (2005), we find that this association is driven solely by the innate component of AQ, with no evidence that discretionary AQ impacts cost of equity.

Overall, this study contributes to extant accounting and finance literature in several ways. First, our results support the theoretical argument that equilibrium asset prices are influenced by information risk and that information risk contributes to cross-sectional differences in firms' costs of capital (e.g., EOH, 2004; O'Hara, 2003; and LLV, 2007). Consistent with FLOS (2005), our results suggest that information risk proxied by AQ is positively related to costs of debt and equity; specifically, poorer accruals quality is associated with higher costs of capital.

Second, our study is amongst the first to examine the effects of accruals quality on the cost of capital for non-US firms. In light of the Australian regulatory and institutional environment, our study demonstrates that different settings can lead to differences in the pricing of reporting quality. Unlike FLOS (2005), who find that both innate and discretionary AQ are significantly associated with cost of capital in the US, we find that the association between accruals quality and cost of capital for Australian firms is driven solely by innate AQ. We attribute this difference in findings to the greater reliance on private debt (as opposed to public debt) and the requirements of the continuous disclosure regime in Australia.

Finally, the paper sheds further light on the issue of whether AQ is a priced risk factor. In contrast to the findings in CGV (2008), our results using the 2SCSR approach indicate that AQ is a priced risk factor for Australian firms. This conclusion is consistent with concurrent studies by Kim and Qi (2008) and Ogneva (2008) using US data. Our findings, therefore, contribute to the growing body of empirical work that suggests that AQ is a priced risk factor.

The remainder of the paper is structured as follows. In Section 2, we develop our hypotheses on the pricing of accruals quality based on the theoretical models of EOH (2004) and LLV (2007). We also note how the Australian institutional and regulatory environment is likely to impact on the relation between AQ, information risk and the cost of capital. Section 3 outlines the empirical methodology and sample selection. In particular, it describes the procedures to measure AQ and partition total AQ into innate and discretionary components. Furthermore, Section 3 also introduces the models used to assess the relation between AQ and the costs of debt and equity capital. Section 4 reports the empirical findings of the study. Section 5 discusses the findings of sensitivity analysis and Section 6 concludes the paper.

2. HYPOTHESIS DEVELOPMENT

The link between the quality of financial reporting and information risk has been the focus of several recent empirical papers (e.g., FLOS, 2005; Aboody et al., 2005; and Chen et al., 2007). These studies rely on theoretical models which suggest that information risk is non-diversifiable and may be priced by the market. For example, EOH (2004) and LLV (2007) argue that accounting information pertaining to a firm's expected

cash flows, amongst other things, affects the information environment surrounding that firm's equilibrium stock returns.

EOH (2004) investigate the behavior of investors (both informed and uninformed) in response to the proportion and precision of private and public information. They argue that the information asymmetry arising from higher levels of private information increases the risk faced by less-informed investors compared to more-informed investors. This information risk is non-diversifiable prompting less-informed investors to demand higher returns on stocks with greater private information. In addition, EOH (2004) note the role of the precision of accounting information in reducing the cost of capital by mitigating the information risk faced by uninformed investors arising from the information asymmetry across investors.

LLV (2007) analytically demonstrate that information precision directly affects equilibrium prices when capital markets are characterized by perfect competition among investors (a condition which EOH's model assumes). LLV (2007) define information precision as the quality of information on a firm's expected cash flows available to investors. In their model, the investor's average information precision is a key determinant of the firm's expected return, and therefore its cost of capital. In summary, the theoretical models of EOH (2004) and LLV (2007) both predict that information risk is non-diversifiable and firms with higher information risk have higher cost of capital.

FLOS (2005) empirically examine whether information risk, proxied by the precision of public information, is priced by the capital markets. While cash flow is regarded as the primitive element of earnings relevant to valuation and risk assessment, earnings are widely believed to convey information about future cash flows. Further, the accrual component of earnings is subject to greater uncertainty than the cash flow component.⁶ Accordingly, FLOS (2005) argue that, the higher the quality of accruals, the better earnings map into cash flows and hence, the lower the information risk and consequently the cost of capital. Consistent with this argument, FLOS (2005) report evidence that US firms with poorer AQ have higher costs of debt and equity capital.

Our first hypothesis, therefore, examines whether there is an association between AQ and the cost of capital. Specifically, we test the following hypothesis:

H₁: The cost of capital for firms with poorer accruals quality is higher than for firms with better accruals quality.

We examine the relation of AQ with both the cost of debt and equity capital. To the extent that AQ captures information risk, and that information risk is priced by the market, we expect to find evidence consistent with H₁.

While the theoretical models in EOH (2004) and LLV (2007) do not differentiate between possible alternate sources of information risk, FLOS (2005) follow the lead from earnings management literature whereby the financial reporting outcome can be partitioned into innate and discretionary components (see, for example, Guay et al., 1996; and Subramanyam, 1996). FLOS (2005) argue that accruals quality can be influenced by economic fundamentals (i.e., innate AQ) and management reporting choices on accounting policies and estimates (i.e., discretionary AQ). It is plausible

⁶ FLOS (2005, p. 301) note that, while cash flows are actually realized, accruals are the product of judgments, estimations and allocations.

that these components of accrual quality have differential effects on information risk and consequently the cost of capital.

FLOS (2005) discuss three possible subcomponents of discretionary AQ and their respective effects on information risk. The performance subcomponent arises when management uses accruals to enhance the ability of earnings to faithfully reflect firm performance. Hence, the performance subcomponent is expected to reduce information risk. In contrast, the opportunistic and noise subcomponents of discretionary AQ are likely to exacerbate information risk. While the net effect of these three subcomponents of discretionary AQ is uncertain, FLOS (2005) predict that innate and discretionary AQ will have differential relations with information risk. Their empirical results suggest that the cost of capital effect of discretionary AQ is smaller in both magnitude and statistical significance than that for innate AQ.

Along these lines, our second hypothesis examines the relation between the cost of capital and the innate and discretionary components of total AQ:

H₂: Innate accruals quality has a greater effect on cost of capital than discretionary accruals quality.

As with H₁, H₂ is examined separately for both the cost of debt and equity capital. While H₂ is identical to the hypothesis tested by FLOS (2005), differences between the US and Australian regulatory and institutional environment motivate its re-examination in our study. As noted in the introduction, Australian firms are more reliant on private debt. Private lenders are likely to have more privileged access to business information compared to public debt lenders. Relative to public debt lenders, private lenders are also likely to monitor firms more closely because of their closer relations with the borrowing firms. To the extent that both effects mitigate opportunistic managerial behavior in relation to discretionary reporting choices, information risk will be reduced as a result of the higher information precision and lower information asymmetry, as well as enhancement in the performance subcomponent of discretionary reporting choices. Further, the non-selective disclosure of high-quality and timely information to the public under the continuous disclosure regime (CDR) in Australia may also mitigate opportunistic managerial discretionary reporting and enhance the performance measurement aspect of financial reporting.⁷ Both the greater dependence on private debt and the CDR in Australia are likely to reduce the information risk associated with discretionary AQ arising from managerial reporting choices. Therefore, we expect that the costs of capital effects arising from information risk in Australia are much more likely to be observed for the innate rather than the discretionary component of AQ.

3. METHODOLOGY AND SAMPLE SELECTION

(i) *Estimating Accruals Quality and its Components*

The empirical analysis requires a metric of AQ and its partition into innate and discretionary components. We adopt the approach developed in Dechow and Dichev (2002) to capture the precision of financial statement information. The Dechow-Dichev

⁷ Using a sample of UK firms in the information technology industry, Gietzmann and Ireland (2005) provide empirical evidence consistent with the prediction that *timelier* disclosure is associated with lower cost of capital.

model recognizes that the timing of a firm's economic achievements and sacrifices often differs from the timing of the related cash flows. The role of accruals, therefore, is to adjust for these cash flow timing problems. Accordingly, the Dechow and Dichev (2002) model captures these timing differences by including leading and lagging cash flows in their accruals model (in addition to contemporaneous cash flows). We adopt McNichols' (2002) modification of Dechow-Dichev model as follows:

$$TCA_{j,t} = \alpha_0 + \alpha_1 CFO_{j,t-1} + \alpha_2 CFO_{j,t} + \alpha_3 CFO_{j,t+1} + \alpha_4 \Delta REV_{j,t} + \alpha_5 PPE_{j,t} + \varepsilon_{j,t}, \quad (1)$$

where, for firm j , $TCA_{j,t}$ is total current accruals in year t measured as income before depreciation and amortization minus operating cash flow, $CFO_{j,t}$ is cash flow from operations in year t , $\Delta REV_{j,t}$ is change operating revenue between year $t - 1$ and year t , $PPE_{j,t}$ is gross property, plant and equipment in year t . All variables are scaled by average total assets measured over year $t - 1$ and t .

In each year, model (1) is estimated cross-sectionally within industry groupings given by four-digit GICS codes. A minimum of 10 firm observations within an industry are required. Using the estimated industry-year coefficients, firm-year specific residuals ($\varepsilon_{j,t}$) are calculated. Our proxy for AQ is the standard deviation of a given firm's residuals over the past five years; that is, $AQ_{j,t} = \text{stdev}(\varepsilon_{j,t})$, $t = t - 5, \dots, t - 1$.⁸ Intuitively, since $\varepsilon_{j,t}$ captures the estimation error in the mapping of accruals to cash flows, larger (smaller) values of AQ indicate poorer (better) accruals quality.⁹

To partition total AQ into innate and discretionary components, we follow Dechow and Dichev (2002) and FLOS (2005) by adopting five innate variables that affect accruals quality (firm size, standard deviation of cash flows from operations, standard deviation of sales revenues, length of operating cycle and incidence of negative earnings realization). Dechow and Dichev (2002) suggest that these five innate variables capture economic fundamentals, as opposed to managerial discretion, that drive accruals quality. The innate and discretionary components of AQ follow from annual cross-sectional estimation of the following model:

$$\begin{aligned} AQ_{j,t} = & \phi_0 + \phi_1 SIZE_{j,t} + \phi_2 \sigma(CFO)_{j,t} + \phi_3 \sigma(Sales)_{j,t} \\ & + \phi_4 OpCycle_{j,t} + \phi_5 NegEarn_{j,t} + v_{j,t}, \end{aligned} \quad (2)$$

where firm size ($SIZE$) is measured as the natural log of total assets, standard deviation of cash flows from operations ($\sigma(CFO)$) is measured over the previous five years, standard deviation of sales revenues ($\sigma(Sales)$) is measured over the previous five years, length of operating cycle ($OpCycle$) measured as log of the sum of days accounts receivable and days inventory, and the incidence of negative earnings realization ($NegEarn$) is measured by the number of years out of the last five with negative reported income before extraordinary items. The predicted values from model (2) proxy for the innate portion of accruals quality ($InnateAQ$), while the residuals proxy for discretionary accruals quality ($DiscAQ$).¹⁰

8 Calculating the AQ measure in year t using firm-specific residuals in year $t - 5$ to year $t - 1$ eliminates the need for future CFO measures in the estimation process.

9 To avoid confusion over this terminology, we will use the terms better (poorer) accruals quality rather than lower (higher) accruals quality.

10 An alternate approach to estimating $DiscAQ$ is explored in Section 5.

(ii) AQ and Cost of Debt

Our proxy for the cost of debt (*CostDebt*) is interest expense divided by the average total debt. The relation between AQ and cost of debt is examined using a regression model that controls for other factors known to affect the cost of debt. Specifically, we control for financial leverage, firm size, return on assets, interest coverage and earnings volatility (Kaplan and Urwitz, 1979; and Palepu et al., 2000). The regression is as follows:

$$\begin{aligned} CostDebt_{j,t} = & \gamma_0 + \gamma_1 Leverage_{j,t} + \gamma_2 Size_{j,t} + \gamma_3 ROA_{j,t} + \gamma_4 IntCov_{j,t} \\ & + \gamma_5 \sigma(NIBE)_{j,t} + \gamma_6 AQrank_{j,t} + \mu_{j,t}, \end{aligned} \quad (3)$$

where *Leverage* is the ratio of total debt to total assets, *Size* is the natural log of total assets, *ROA* is return on assets, *IntCov* is the ratio of operating income to interest expense and $\sigma(NIBE)$ is the standard deviation of net income before extraordinary items, scaled by average assets, over the past five years. Consistent with FLOS (2005), we use firm *j*'s decile rank of AQ rather than the raw AQ score. Using decile rank controls for outliers and non-linearity, and facilitates interpretation of the economic magnitude of the cost of capital effect. The estimated coefficient on *AQrank* captures the accruals quality effect on cost of debt that is incremental to the control factors. All else equal, if lenders view firms with poorer AQ as riskier than firms with better AQ, there will be a positive association between *CostDebt* and *AQrank*.

(iii) AQ and Cost of Equity

Following FLOS (2005), we use an industry-adjusted earnings-to-price ratio (*IndEP*) to proxy for cost of equity. Specifically, *IndEP* is the firm's earnings-price ratio for a particular year less the median earnings-price ratio of all firms with the same 4-digit GICS industry code. We use the following model to examine the relation between AQ and cost of equity:

$$\begin{aligned} IndEP_{j,t} = & \delta_0 + \delta_1 Growth_{j,t} + \delta_2 Leverage_{j,t} + \delta_3 Beta_{j,t} \\ & + \delta_4 Size_{j,t} + \delta_5 AQrank_{j,t} + \zeta_{j,t}, \end{aligned} \quad (4)$$

where *Growth_{j,t}* is the natural log of one plus the firm's growth in book value of equity over the past five years, *Beta_{j,t}* is the CAPM firm-specific beta estimated using rolling five-year regressions (for firms with at least 18 monthly returns). All other variables are as previously defined. The estimated coefficient on *AQrank* captures the accruals quality effect on cost of equity that is incremental to the control factors. All else equal, if investors view firms with poorer AQ as riskier than firms with better AQ, there will be a positive association between *IndEP* and AQ.

(iv) Asset Pricing Tests

We use two approaches to examine whether AQ is a priced risk factor. The first approach is based on FLOS (2005), who document statistically significant loadings on a factor-mimicking portfolio constructed to capture accruals quality. Consistent with FLOS (2005) and in the spirit of the Fama and French (1993) analysis of size and book-to-market factors, we form an AQ factor-mimicking portfolio equal to the difference

between the monthly returns to the best two AQ quintiles and the worst two AQ quintiles. Portfolios are re-formed on a monthly basis using the most-recent AQ measure to accommodate inter-temporal changes in accruals quality and differences in firms' financial year ends. The following asset-pricing model is estimated:

$$R_{j,t} - R_{f,t} = \alpha_j + \beta_{j,MRP} (R_{m,t} - R_{f,t}) + \beta_{j,SMB} SMB_t + \beta_{j,HML} HML_t + \beta_{j,AQ} AQfactor_t + \varepsilon_{j,t}, \quad (5)$$

where $R_{j,t}$, $R_{f,t}$ and $R_{m,t}$ are the time t returns on stock j , the risk-free asset and the market portfolio respectively, SMB and HML are the Fama-French size and book-to-market factors respectively, and $AQfactor$ is the accrual quality factor. Model (5) represents the common three-factor asset-pricing model augmented by the AQ factor. FLOS (2005) interpret a significant estimate of $\beta_{j,AQ}$ as support for the notion that AQ is priced.

Our second approach relies on CGV (2008) who argue that documenting a contemporaneous relation between the returns on assets and the AQ factor-mimicking portfolio (i.e., by the significance of $\beta_{j,AQ}$ in model 5) does *not* explicitly test the hypothesis that AQ is a priced risk factor. Rather, the positive coefficient from such a time-series regression merely indicates that, on average, firms have positive exposure to the AQ mimicking factor.¹¹

Following Cochrane (2005), CGV suggest that a better-specified test of whether a proposed factor is priced requires a two-stage cross-sectional regression (2SCSR) approach.¹² Rather than implementing the 2SCSR approach on individual stock returns, we follow the majority of the empirical asset-pricing literature by using the returns on 25 portfolios cross-sorted on size and book-to-market ratio. The 25 portfolios are formed using a procedure identical to that of Fama and French (1993).¹³ In stage 1, factor betas are estimated using time-series regressions of excess portfolio returns on the contemporaneous returns of proposed risk factors as shown in model (5). Stage 2 then estimates a cross-sectional regression of the mean excess portfolio returns on the factor betas estimated in model (5):

$$\bar{R}_{p,t} - \bar{R}_{f,t} = \lambda_0 + \lambda_1 \hat{\beta}_{p,MRP} + \lambda_2 \hat{\beta}_{p,SMB} + \lambda_3 \hat{\beta}_{p,HML} + \lambda_4 \hat{\beta}_{p,AQ} + u_{p,t} \quad (6)$$

where $\bar{R}_{p,t} - \bar{R}_{f,t}$ is the time-series mean excess return for portfolio p , $\hat{\beta}_{p,MRP}$ is the market factor beta, $\hat{\beta}_{p,SMB}$ is the size factor beta, $\hat{\beta}_{p,HML}$ is the book-to-market factor beta, and $\hat{\beta}_{p,AQ}$ is the AQ factor beta derived from estimating model (5) in stage 1 using portfolio returns. This approach examines whether AQ is a priced risk factor, after controlling for the three Fama-French (1993) factors. Specifically, λ_4 will have a significant positive coefficient if AQ is priced. Standard errors are calculated using Shanken's (1992) correction to reflect the fact that the independent variables in the cross-sectional model (6) are estimated in Stage 1 (i.e., they are 'generated regressors').

11 Using a familiar analogy, CGV (2008, p. 3) note that a positive coefficient in a contemporaneous regression of stock returns on the market portfolio does not imply that beta is priced, but simply confirms that the average beta in a random sample of firms is positive and mechanically close to one.

12 The 2SCSR approach has been widely used to test asset pricing models. For example, in testing the CAPM (Fama and MacBeth, 1973), the conditional CAPM (Brennan et al., 2004; and Petkova, 2006), and the two-beta model (Campbell and Vuolteenaho, 2004)

13 We sincerely thank Michael O'Brien for supplying these SIZE/BM portfolio returns. Details of the portfolio formation procedure are outlined in O'Brien, Brailsford and Gaunt (2008).

(v) Sample Selection

As AASB1026 *Statement of Cash Flows* came into effect in 1992, we restrict our initial sample to post-1992 so that the cash flows from operations can be measured directly from the Statement of Cash Flows.¹⁴ Accounting data for the study are drawn from the Aspect FinAnalysis database for the period 1992–2005. To be included in the sample for a given year, a firm must have at least seven years of financial data to compute the AQ measure (model 1), in addition to the financial data to compute the five innate variables that proxy for economic fundamentals driving AQ (model 2). These data requirements restrict our final sample period from 1998 to 2005. In order to construct the AQ_{factor} used in the asset-pricing tests (model 5), we also require returns data which results in a base sample of 736 firms from 1998–2006. The estimation of models (3) and (4) imposes further data requirements for the control variables. The final sample is 509 firms (2,057 firm-year observations) for the cost of debt model (3) and 346 firms (1,362 firm-year observations) for the cost of equity model (4).

Table 1 provides descriptive statistics on the variables used in the empirical analysis. The distribution of AQ has a mean of 0.081, but exhibits significant dispersion with a standard deviation of 0.061. FLOS (2005) report lower mean AQ of 0.044, which may be attributable to their larger sample. Recall that model (1) is fitted each year on a per industry basis, with AQ measured as the standard deviation of the fitted residuals. With significantly fewer sample firms per year per industry than in FLOS (2005), it is difficult to achieve a tight fit, which manifests itself in the AQ metric.

4. RESULTS AND DISCUSSION

(i) Relation Between AQ and Economic Fundamentals

The partition of total AQ into innate and discretionary components is based on model (2) which models total AQ as a function of five variables reflecting economic fundamentals. Table 2 reports the time-series mean coefficient estimates and the associated Fama and MacBeth (1973) *t*-statistics from the annual cross-sectional regressions over the period 1998–2005.

The estimated coefficient on each of the five variables has the predicted sign and is statistically significant at the 5% level or better. Specifically, firm size is negatively related with AQ, while the remaining variables are positively related. It is important to recall that AQ is measured as the standard deviation of the error term from model (1); hence, a 'high' AQ score reflects poor quality accruals and vice versa. The average goodness-of-fit of the cross-sectional regressions exceeds 30%. The current results, which are highly consistent with prior findings reported by Dechow and Dichev (2002) and FLOS (2005), suggest that the five innate variables are reasonable proxies for economic fundamentals that drive accruals quality.

For the subsequent analysis, the innate and discretionary components of AQ are derived from the fitted and error elements of model (2) respectively. Table 1 reports summary statistics for *InnateAQ* and *DiscAQ*. Since by construction *DiscAQ* is mean zero

14 Hribar and Collins (2002) suggest that significant measurement errors arise when accruals and cash flows from operations are measured from changes in balance sheet items, as opposed to accruals and cash flows from operations measured using the cash flows statement.

Table 1
Summary Statistics

| Variables | Mean | Std Dev | 10% | 25% | Median | 75% | 90% | N |
|----------------------------------|----------|----------|--------|--------|--------|---------|----------|------|
| AQ | 0.081 | 0.061 | 0.025 | 0.037 | 0.064 | 0.104 | 0.164 | 2057 |
| Inmate AQ | 0.081 | 0.036 | 0.042 | 0.053 | 0.073 | 0.102 | 0.128 | 2057 |
| Disc AQ | 0.000 | 0.046 | -0.048 | -0.029 | -0.009 | 0.020 | 0.058 | 2057 |
| Financial variables | | | | | | | | |
| Market value of equity (\$m) | 807.437 | 2397.137 | 4.533 | 13.085 | 47.080 | 292.138 | 2031.160 | 2057 |
| Total Assets (\$m) | 1034.905 | 3739.479 | 9.121 | 22.192 | 72.384 | 389.231 | 2162.530 | 2057 |
| Sales (\$m) | 695.503 | 1951.648 | 3.572 | 15.297 | 73.827 | 395.336 | 1510.420 | 2057 |
| ROA | -0.002 | 0.196 | -0.183 | -0.001 | 0.050 | 0.081 | 0.130 | 2057 |
| Book-to-market | 1.007 | 0.954 | 0.246 | 0.437 | 0.735 | 1.245 | 1.975 | 2057 |
| CostDebt | 8.719 | 10.785 | 0.000 | 5.156 | 7.054 | 9.275 | 13.742 | 2057 |
| Leverage | 0.211 | 0.168 | 0.000 | 0.058 | 0.200 | 0.321 | 0.432 | 2057 |
| σ (NIBE) | 0.119 | 0.160 | 0.013 | 0.023 | 0.056 | 0.150 | 0.309 | 2057 |
| EP ratio | 0.089 | 0.067 | 0.029 | 0.051 | 0.074 | 0.107 | 0.160 | 1362 |
| IndEP | 0.014 | 0.064 | -0.042 | -0.021 | 0.000 | 0.028 | 0.076 | 1362 |
| Sales growth | 0.715 | 3.874 | -0.250 | -0.053 | 0.081 | 0.268 | 0.761 | 2024 |
| Growth (in book value of equity) | 0.375 | 0.929 | -0.607 | -0.031 | 0.331 | 0.784 | 1.440 | 2033 |
| Inmate variables | | | | | | | | |
| Size (log of total assets) | 11.529 | 2.074 | 9.118 | 10.007 | 11.190 | 12.872 | 14.587 | 2057 |
| σ (CFO) | 0.090 | 0.090 | 0.021 | 0.034 | 0.058 | 0.112 | 0.197 | 2057 |
| σ (Sales) | 0.229 | 0.218 | 0.045 | 0.083 | 0.155 | 0.297 | 0.495 | 2057 |
| OpCycle (in days) | 121.951 | 135.013 | 36.040 | 59.140 | 89.380 | 129.180 | 208.760 | 2057 |
| OpCycle | 4.481 | 0.772 | 3.585 | 4.080 | 4.493 | 4.861 | 5.341 | 2057 |
| NegEarn | 0.324 | 0.350 | 0.000 | 0.000 | 0.200 | 0.600 | 1.000 | 2057 |

Notes:

AQ = standard deviation of firms' residuals, from years $t - 5$ to $t - 1$ from annual cross-sectional estimations of the modified Dechow-Dichev (2002) model as per model (1). Inmate AQ = predicted value obtained from the annual parameter estimates from model (2) and firms' reported value of the innate factors (see model (3)). Disc AQ = is the residual from model (2). ROA = return on assets. CostDebt = interest expense divided by the average total debt. Leverage = ratio of total debt to total assets. σ (NIBE) = standard deviation of net income before extraordinary items, scaled by average assets, over the past five years. EP ratio = earnings-price ratio. IndEP = earnings-price ratio less the median earnings-price ratio of all firms in the same 4-digit SIC industry in which the firm operates during that year. Sales growth = year-to-year percentage change in sales. Growth (in book value of equity) = log of one plus the firm's growth in book value of equity over the past five years. Size = log of total assets. σ (CFO) = standard deviation of CFO over the past five years. σ (Sales) = standard deviation of sales over the past five years. OpCycle = log of operating cycle. NegEarn = number of years, out of the past five, where reported income before extraordinary items is less than zero.

Table 2

Regression Estimates of the Relation Between AQ and Economic Fundamentals

 $(N = 2,057; 509 \text{ firms})$

$$AQ_{j,t} = \phi_0 + \phi_1 SIZE_{j,t} + \phi_2 \sigma(CFO)_{j,t} + \phi_3 \sigma(Sales)_{j,t} + \phi_4 OpCycle_{j,t} + \phi_5 NegEarn_{j,t} + v_{j,t}$$

| | <i>Pred. Sign</i> | <i>Estimate</i> | <i>t-stat</i> |
|-----------------|-------------------|-----------------|---------------|
| <i>Size</i> | – | –0.0023 | –5.24*** |
| $\sigma(CFO)$ | + | 0.1926 | 7.17*** |
| $\sigma(Sales)$ | + | 0.0425 | 6.57*** |
| <i>OpCycle</i> | + | 0.0030 | 2.69* |
| <i>NegEarn</i> | + | 0.0082 | 16.09*** |
| Adj. R^2 | | 0.3123 | |

Notes:

This table reports time-series mean coefficients and Fama-MacBeth *t*-statistics from annual cross-sectional regressions of model (2). ***, ** and * denote statistical significance at the 0.1%, 1% and 5% levels (one-tailed) respectively.

AQ = standard deviation of firms' residuals, from years $t - 5$ to $t - 1$ from annual cross-sectional estimations of the modified Dechow-Dichev (2002) model as per model (1). *Size* = log of total assets. $\sigma(CFO)$ = standard deviation of *CFO* over the past five years. $\sigma(Sales)$ = standard deviation of sales over the past five years. *OpCycle* = log of operating cycle. *NegEarn* = number of years, out of the past five, where reported income before extraordinary items is less than zero.

(it is the error term from an OLS regression), the mean *InnateAQ* is identical to the mean total AQ (0.081). Nonetheless, there is considerable variation around *DiscAQ* (std dev = 0.046) and this uncertainty has a non-trivial impact on total AQ. For example, the 10th and 90th percentiles for the distribution of total AQ (0.025 and 0.164) are significantly wider than those for *InnateAQ* (0.042 to 0.128), reflecting the additional variation induced into total AQ by *DiscAQ*. This finding further motivates our analysis of the separate effects of *InnateAQ* and *DiscAQ* on costs of capital, as opposed to simply examining total AQ.

(ii) AQ and Cost of Debt

Table 1 reports a mean (median) cost of debt over the sample period of 8.7% (7.1%). These statistics are marginally lower than those reported by FLOS (2005) for US firms (mean and median of 9.9% and 9.2% respectively) which is likely attributable to the low interest-rate environment in Australia over the sample period.¹⁵ Nonetheless, the sample exhibits considerable dispersion in *CostDebt* with the 10th and 90th percentiles being 0% and 13.74% respectively.

Table 3 Panel A presents estimates for model (3). In order to address concerns over cross-sectional and time-series dependence, we use pooled regressions with fixed year

¹⁵ During FLOS's sample period (1970–2001), the mean and median effective overnight Federal fund rates were 7.26% and 6.39% respectively (data sourced from www.federalreserve.gov/Releases/). In contrast, during our sample period of 1998–2005, the mean and median cash rates in Australia were 5.10% and 5.00% respectively (data available from www.rba.gov.au/Statistics/Bulletin/index.html).

Table 3
Cost of Capital Effects of Accruals Quality

| Indep. Var. | Total AQ | | Clustered <i>t</i> -stats | AQ Components | |
|--|---------------|---------|---------------------------|---------------|---------------------------|
| | Pred. Sign | Coef. | | Coef. | Clustered <i>t</i> -stats |
| Panel A: Relation Between Accruals Quality and Cost of Debt (N = 2,057; 509 firms) | | | | | |
| $CostDebt_{j,t} = \gamma_0 + \gamma_1 Leverage_{j,t} + \gamma_2 Size_{j,t} + \gamma_3 ROA_{j,t} + \gamma_4 IntCov_{j,t} + \gamma_5 \sigma(NIBE)_{j,t} + \gamma_6 AQrank_{j,t} + \mu_{j,t}$ | | | | | |
| Leverage | + | -5.4101 | -2.53 | -5.6921 | -2.62 |
| Size | - | -0.0730 | -0.55 | 0.2972 | 1.48 |
| ROA | - | -0.0702 | -0.03 | 0.3056 | 0.12 |
| IntCov | - | -0.0002 | -0.97 | -0.0002 | -0.99 |
| $\sigma(NIBE)$ | + | 6.5077 | 1.94* | 4.0544 | 1.13 |
| AQrank | + | -0.0392 | -0.35 | | |
| AQrank | + | | | 0.4392 | 2.43** |
| AQrank | ? | | | -0.1677 | -1.25 |
| F test | Innate > Disc | | | | 5.87** |
| Panel B: Relation Between Accruals Quality and Cost of Equity (N = 1,362; 346 firms) | | | | | |
| $IndEP_{j,t} = \delta_0 + \delta_1 Growth_{j,t} + \delta_2 Leverage_{j,t} + \delta_3 Beta_{j,t} + \delta_4 Size_{j,t} + \delta_5 AQrank_{j,t} + \zeta_{j,t}$ | | | | | |
| Growth | - | -0.0107 | -3.22*** | -0.0109 | -3.37*** |
| Leverage | + | 0.0270 | 1.75* | 0.0242 | 1.59 |
| Beta | + | 0.0159 | 1.87* | 0.0124 | 1.38 |
| Size | - | -0.0053 | -4.58*** | -0.0030 | -2.20* |
| AQrank | + | 0.0021 | 2.50** | | |
| AQrank | + | | | 0.0038 | 3.78*** |
| AQrank | ? | | | 0.0010 | 0.98 |
| F test | Innate > Disc | | | | 4.34* |

Notes:

***, **, and * denote statistical significance at the 0.1%, 1% and 5% levels respectively (one-tailed if signs are predicted, two-tailed otherwise). We only note significance if signs are in the predicted direction. The clustered *t*-statistics are based on pooled regressions with fixed year effects and adjusted for firm-level clustering. $CostDebt$ = interest expense divided by the average total debt. $Leverage$ = ratio of total debt to total assets. $Size$ = log of total assets. ROA = return on assets. $IntCov$ = ratio of operating income to interest expense. $\sigma(NIBE)$ = standard deviation of net income before extraordinary items, scaled by average assets, over the past five years. $IndEP$ = earnings-price ratio less the median earnings-price ratio of all firms in the same 4-digit SIC industry in which the firm operates during that year. $Growth$ = log of one plus the firm's growth in book value of equity over the past five years. $Beta$ = five-year rolling beta estimated from firm-specific CAPM estimation for firms with at least 18 monthly returns. $AQrank$ is an integer representing the decile rank of a firm based on accruals quality.

effects and standard errors clustered by firm.¹⁶ Table 3 Panel A shows that earnings volatility (σ (*NIBE*)) is positively related to *CostDebt*, but like FLOS (2005), the negative sign on financial leverage (*Leverage*) is opposite to that predicted. Most importantly, the coefficient on *AQrank* is insignificant. As such, there is no evidence to support the conjecture of H_1 that firms with poor AQ have higher cost of debt. This finding differs from that of FLOS (2005) who report a modest, albeit statistically significant, positive relation between total AQ and cost of debt.

To test H_2 , we extend the analysis to examine whether the innate and discretionary components of total AQ have differential effects on cost of debt. Our approach is to replace the *AQrank* for total AQ in model (3) with similarly-constructed decile rankings for both *InnateAQ* and *DiscAQ* measures. The modified model (3), which retains the five control variables, estimates the unique impact of innate and discretionary AQ on cost of debt.

Table 3 Panel A reports that *InnateAQ* has a significant positive association with *CostDebt* – the poorer the innate accruals quality, the higher the cost of debt. The estimated coefficient of 0.4392 implies that the cost of debt increases by about 395 basis points as we move from firms in the best decile of *InnateAQ* to those in the worst decile of *InnateAQ*. In sharp contrast, the relation between *DiscAQ* and cost of debt is insignificant. An *F*-test of the hypothesis that the true coefficients on *InnateAQ* and *DiscAQ* are equal is rejected.¹⁷

Consistent with H_2 , these findings support the notion that innate AQ has a greater effect on Australian firms' cost of debt than discretionary AQ. While this finding is consistent with FLOS (2005), our results display one important difference. FLOS report that both the innate and discretionary components of total AQ have a significant impact on the cost of debt, the former being larger in magnitude. Our results suggest that the cost of debt for Australian firms is only affected by *InnateAQ*.¹⁸ This finding is consistent with the earlier argument that, given their greater reliance on private debt, Australian firms operate in an environment of higher information precision, lower information asymmetry among and closer monitoring by private lenders. Thus, information risk associated with managerial reporting discretion is reduced resulting in the diminished effect of discretionary AQ on the cost of debt.

(iii) AQ and Cost of Equity

Our analysis of the effect of AQ on cost of equity (as proxied by *IndEP*) is based on model (4). We only include firms with positive earnings to obtain a meaningful

16 Petersen (2008) suggests that, in the presence of cross-sectional and time-series dependence, one dependence effect can be addressed parametrically (e.g., including time dummies for cross-sectional dependence) and then estimate standard errors clustered on the other dependence effect (e.g., clustering by firms for time-series dependence). As we have more firm than year observations, we use year dummies and cluster by firms because a larger number of clusters leads to standard errors that are less biased.

17 Our full sample contains a number of firms that do not report any interest expense, resulting in an estimated cost of debt of zero. As a robustness check, we repeat the analysis in Table 3 Panel A after excluding firms with reported interest expense equal to zero. Untabulated results are qualitatively similar and do not change our inferences.

18 Rather than using a pooled regression with clustered *t*-statistics, FLOS (2005) simply report *t*-statistics based on the mean coefficients of annual regressions of model (3). As a robustness check, we also estimate model (3) on an annual basis, with the coefficients being the time-series means of the annual estimates. The untabulated results are qualitatively similar to those using the pooled regression. That is, the coefficients on total AQ and *DiscAQ* remain insignificant while the coefficient on *InnateAQ* is significantly positive.

earnings-to-price ratio, which reduces the sample to 346 firms ($N = 1,362$). Table 1 reports a mean (median) earnings-price ratio over the sample period of 0.089 (0.074), with 80% of earnings-price ratios ranging between 0.029 and 0.160. These statistics are very similar to those reported by FLOS (2005) for their US sample where the mean (median) of earnings-price ratio is 0.089 (0.073) and 80% of their sample's earnings-price ratios range between 0.026 and 0.166. These statistics suggest the cost of equity is similar for US and Australian firms.

Table 3 Panel B presents regression estimates for model (4). Each of the four control variables is statistically significant with the predicted sign. The coefficient on *AQrank* is positive and significantly related to *IndEP*. This finding suggests that firms with poorer AQ have higher cost of equity, thus supporting H_1 . Regarding economic significance, the coefficient estimate (0.0021) implies that the industry-adjusted cost of equity increases by about 189 basis points as we move from firms in the best decile of *AQrank* to those in the worst decile of *AQrank*.

In terms of H_2 , Table 3 Panel B shows that the coefficient on *InnateAQ* is significantly positive, while the coefficient on *DiscAQ* is insignificant.¹⁹ An *F*-test of the hypothesis that *InnateAQ* and *DiscAQ* have equal impact on *IndEP* is rejected. The economic significance of *InnateAQ* for cost of equity is unambiguous – the differential in industry-adjusted cost of equity between firms with the poorest and best *InnateAQ* is around 342 basis points. The finding that discretionary AQ is not significantly associated with Australian firms' cost of equity may in part be a function of the continuous disclosure regime in place over our sample period. The non-selective disclosure of high quality and timely information to the public domain increases the precision of information and mitigates information asymmetry on expected cash flows. Thus, the information risk associated with discretionary AQ arising from managerial opportunistic reporting is reduced.

(iv) *AQ and Asset Pricing Implications*

Our first analysis of the asset-pricing implications of accruals quality replicates the approach of FLOS (2005) by estimating the loadings on *AQfactor* when it is added to the three-factor asset-pricing model (5). Using a pooled regression, model (5) is estimated for firms with sufficient data to calculate their AQ measure and at least 18 monthly returns between September 1998 and August 2006, resulting in a sample of 736 firms and 65,110 firm-month observations.²⁰ Table 4 reports the mean and standard deviation of parameter estimates with *t*-statistics based on clustering by month, along with regression goodness-of-fit statistics.²¹

19 Note that, as in model (3), FLOS (2005) report *t*-statistics based on the mean coefficients of annual regressions of model (4). Replicating this approach, the results (not explicitly reported) are qualitatively similar to those using the pooled regression corrected for cross-sectional and time-series dependencies reported in Table 3 Panel B.

20 As a sensitivity check, we relax the data requirement that sample firms must have available data for AQ to calculate the *AQfactor*, in addition to returns. Specifically, we repeat the analysis reported in Table 4 for all firms that have at least 18 months of returns over our sample period. This increases the sample from 736 firms to 1,828 firms. The untabulated results remain qualitatively similar and do not affect our inferences.

21 The econometric concern here is cross-sectional dependence because, over long time windows, common market shocks induce high cross-sectional correlation. As noted by Petersen (2008), clustering by time (e.g., by month or by year) will correct the standard errors for cross-sectional dependence. We do not address time-series dependence econometrically in the model because the dependent variable is returns, which is expected to be serially independent in an efficient market.

Table 4
 Asset-pricing Tests with AQ Factor-mimicking Portfolio

$$R_{j,t} - R_{f,t} = \alpha_j + \beta_{j,MRP}(R_{m,t} - R_{f,t}) + \beta_{j,SMB}SMB_t + \beta_{j,HML}HML_t + \beta_{j,AQ}AQfactor_t + \varepsilon_{j,t}$$

| Independent Variable | Predicted Sign | Column 1 Three-factor Base Case | | Column 2 Three-factor with AQ factor | | Column 3 Three-factor with AQ Components | |
|----------------------|-------------------------|------------------------------------|-------------------------------|---|-------------------------------|---|-------------------------------|
| | | Coefficient | Clustered <i>t</i> -statistic | Coefficient | Clustered <i>t</i> -statistic | Coefficient | Clustered <i>t</i> -statistic |
| Intercept | | 0.0002 | 0.13 | 0.0007 | 0.50 | 0.0019 | 1.43 |
| $R_m - R_f$ | + | 1.2066 | 21.66*** | 1.1488 | 19.21*** | 1.1508 | 21.77*** |
| SMB | + | 0.8372 | 38.93*** | 0.7334 | 17.16*** | 0.6517 | 14.92*** |
| HML | + | 0.3159 | 5.28*** | 0.2947 | 4.97*** | 0.2634 | 4.55*** |
| AQ factor | + | | | 0.0812 | 2.97** | | |
| AQ factor | + | | | | | 0.1293 | 5.29*** |
| AQ factor | ? | | | | | -0.0211 | -0.88 |
| <i>F</i> test | <i>Inmate > Disc</i> | | | | | | 27.27*** |
| Adj. R^2 | | 0.1079 | | 0.1126 | | 0.1123 | |

Notes:

***, **, * denote statistically significant at the 0.1%, 1% and 5% levels (one-tailed if signs are predicted, two-tailed otherwise) respectively. The 'clustered *t*-statistics' are pooled regressions *t*-statistics based on clustering the standard errors by month.

$(R_j - R_f)$ = firm specific excess returns. $(R_m - R_f)$ = excess return on the market portfolio. *SMB* = return to size factor-mimicking portfolio. *HML* = return to market-to-book factor-mimicking portfolio. *AQ factor* = return to the accruals quality factor-mimicking portfolio for AQ.

As a base case, a three-factor model is estimated initially (i.e., without the *AQfactor*). Table 4 Column 1 shows that market risk premium, *SMB* and *HML* factors are statistically significant, with an adjusted R^2 for the base-case model of 10.79%. When this model is augmented with the factor-mimicking portfolio designed to capture accruals quality (*AQfactor*), Table 4 Column 2 shows a marginal increase in goodness of fit (11.26%). Most importantly, the loading on *AQfactor* is significantly positive. The decline in reported loading on *SMB* suggests that it may be correlated with *AQfactor*.²²

Table 4 Column 3 shows that the significance of accruals quality is solely attributable to *InnateAQ*. When the single *AQfactor* in model (5) is replaced with separate factor-mimicking portfolios constructed for *InnateAQ* and *DiscAQ*, only the innate component of *AQ* is associated with asset returns. An *F*-test suggests that the difference between the loadings on *InnateAQ* and *DiscAQ* is statistically significant at the 1% level. This finding differs from FLOS (2005), who report that both *InnateAQ* and *DiscAQ* are significantly related with stock returns. We attribute this difference to the CDR regulatory environment in existence in Australia.

Our second asset-pricing analysis adopts the 2SCSR approach commonly employed to test proposed risk factors using 25 portfolios cross-sorted on size and book-to-market ratio. Table 5 Panel A reports the mean parameter estimates from the stage 1 time-series regression (5) estimated for each of the 25 size/BM portfolios. Similar to Table 4, we report a base case using the three-factor model before including *AQ* as an additional factor. The parameter estimates for the portfolios are similar to those reported for individual stock regressions in Table 4 for all three models (i.e., base case model, three-factor model with *AQ* and three-factor model with *AQ* components). This is not surprising given that the sample of individual stocks represents a broad cross-section of the population. Further, and as expected, the average adjusted R^2 (69%) is significantly higher in Table 5 Panel A – the portfolio-formation procedure attenuates much of the idiosyncratic noise surrounding individual stocks.²³

The results reported in Table 5 are interesting on a number of dimensions. First, they are generally consistent with the conclusion from individual-stock regressions in Table 4 (namely, that the *AQfactor* is positively related with asset returns, and that this association is driven by the innate component of *AQ*). Second, the findings are in contrast to CGV (2008) who find no association between the *AQfactor* and portfolio returns for US stocks.²⁴

According to CGV (2008), a well-specified test of whether a proposed risk factor is priced is judged in the stage 2 cross-sectional regression of average portfolio excess returns on the factor betas estimated in stage 1. Table 5 Panel B reports the estimated factor loadings (i.e., the λ s) from estimating model (6). While CGV (2008) find little

22 Note that FLOS (2005) only report the mean coefficients from *firm-specific* regressions of model (5) instead of pooled regressions with clustering by time (i.e., month). As noted in Barth et al. (2006), an econometric concern with averaging the coefficients from these firm-specific regressions is that the standard errors (*t*-statistics) are not adjusted for cross-sectional correlation and are likely to be biased downwards (upwards). We also estimate model (5) using firm-specific regressions as in FLOS (2005) for comparison. The results based on the mean coefficients are qualitatively similar to those using pooled regression with clustering by month.

23 Indeed, this is the original motivation for Black, Jensen and Scholes (1972) and Fama and MacBeth (1973) to use portfolios rather than individual stocks in their asset-pricing tests.

24 CGV note that the appropriate test of whether model (5) constitutes a well-specified asset-pricing model requires a test of whether the intercepts (α_j) are jointly zero across the 25 SIZE/BM portfolios. Consistent with the findings in CGV, the Gibbons et al. (1989) test of this hypothesis is unambiguously rejected for both the three-factor and augmented models in Table 5.

Table 5
2-stage Cross-sectional Regression

| Independent Variable | Predicted Sign | Three-factor Base Case | | Three-factor with AQfactor | | Three-factor with AQ Components | |
|--|----------------|------------------------|-------------|----------------------------|-------------|---------------------------------|-------------|
| | | Coefficient | t-statistic | Coefficient | t-statistic | Coefficient | t-statistic |
| Panel A: Stage 1 Time-series Regressions of Portfolio Returns on Risk Factors^a | | | | | | | |
| $R_{p,t} - R_{f,t} = \alpha_j + \beta_{p,MRP}(R_{m,t} - R_{f,t}) + \beta_{p,SMB}SMB_t + \beta_{p,HML}HML_t + \beta_{p,AQ}AQ_{factor,t} + \varepsilon_{p,t}$ | | | | | | | |
| Intercept | | -0.0019 | -0.67 | -0.0008 | -0.41 | -0.0001 | -0.02 |
| $R_m - R_f$ | + | 1.0976 | 22.33*** | 1.0524 | 23.85*** | 1.0602 | 23.38*** |
| SMB | + | 0.7143 | 6.15*** | 0.6336 | 6.63*** | 0.5925 | 5.96*** |
| HML | + | 0.2833 | 3.69** | 0.2645 | 3.58** | 0.2482 | 3.28** |
| AQ_factor | + | | | 0.0628 | 2.40* | | |
| AQ_factor | + | | | | | 0.0844 | 5.18*** |
| AQ_factor | ? | | | | | -0.0112 | -0.68 |
| Adj. R ² | | 0.6885 | | 0.6922 | | 0.6900 | |
| Panel B: Stage 2 Cross-sectional Regression of Returns on Factor Betas^b | | | | | | | |
| $R_{p,t} - R_{f,t} = \lambda_0 + \lambda_1\hat{\beta}_{p,MRP} + \lambda_2\hat{\beta}_{p,SMB} + \lambda_3\hat{\beta}_{p,HML} + \lambda_4\hat{\beta}_{p,AQ} + u_{p,t}$ | | | | | | | |
| Intercept | | 0.0112 | 2.43* | 0.0179 | 3.06** | 0.0141 | 2.65* |
| $\hat{\beta}_{p,MRP}$ | + | -0.0109 | -2.11* | -0.0149 | -3.34* | -0.0134 | -2.16* |
| $\hat{\beta}_{p,SMB}$ | + | 0.0152 | 2.15** | 0.0096 | 1.40 | 0.0121 | 1.75 |
| $\hat{\beta}_{p,HML}$ | + | 0.0131 | 4.00*** | 0.0095 | 2.69* | 0.0122 | 3.43** |
| $\hat{\beta}_{p,AQ}$ | + | | | 0.0477 | 2.43* | | |
| $\hat{\beta}_{p,AQ}$ | + | | | | | 0.0459 | 2.19* |
| $\hat{\beta}_{p,AQ}$ | ? | | | | | 0.0104 | 0.59 |
| Adj. R ² | | 0.6245 | | 0.6545 | | 0.5946 | |

Notes:

^a Panel A reports estimates from time-series regressions of monthly portfolio returns (in excess of the risk-free rate) on the three Fama-French factors and the AQfactor. The reported coefficients are the average estimate across the 25 size/BM portfolios. ***, ** and * reflect statistical significance at the 0.1%, 1% and 5% levels (one-tailed if signs are predicted, two-tailed otherwise) respectively. Standard errors are calculated from the regression estimates across 25 portfolios.

^b Panel B reports estimated coefficients from a cross-sectional regression of the mean excess portfolio return on 25 size/BM portfolios on full-period factor betas. ***, ** and * reflect statistical significance at the 0.1%, 1% and 5% levels (one-tailed if signs are predicted, two-tailed otherwise) respectively. Standard errors are calculated using the Shanken (1992) correction to reflect that the factor betas are estimated in the Stage 1 time-series regressions.

$(R_p - R_f)$ = portfolio specific excess returns. $\bar{R}_{p,t} - \bar{R}_{f,t}$ = mean excess portfolio returns over 96 months (September 1998 and August 2006). $(R_{m,t} - R_{f,t})$ = excess return on the market portfolio. SMB = return to size factor-mimicking portfolio. HML = return to market-to-book factor-mimicking portfolio. AQ_{factor} = return to the accruals quality factor-mimicking portfolio for AQ. $\hat{\beta}_{p,MRP}$ = market factor beta. $\hat{\beta}_{p,SMB}$ = size factor beta. $\hat{\beta}_{p,HML}$ = market-to-book factor beta. $\hat{\beta}_{p,AQ}$ = accruals quality factor beta.

evidence that AQ loads as a significant factor, our evidence suggests that the AQ factor is significantly related to average portfolio returns, in particular the innate component of AQ. To summarize, using the 2SCSR methodology advocated by CGV (2008), we find evidence supporting the conclusion in FLOS (2005) that AQ is a priced risk factor.

5. SENSITIVITY ANALYSIS

The primary approach adopted in this paper to partition total AQ into innate and discretionary components is described in Section 3(i) (see model 2). Specifically, five variables representing economic fundamentals are used to estimate *InnateAQ*, while the regression residual estimates *DiscAQ*. This approach is useful in that it facilitates statistical inference for H_2 on the differential impact of the two components of AQ.

However, this procedure has potential drawbacks.²⁵ In model (2), non-systematic noise (i.e., noise that is unrelated to the five innate variables) is captured in the regression residual, which is our proxy for *DiscAQ*. Further, the residual may be affected by model mis-specification caused by any omitted innate variables. Both of these problems impact on our estimate of *DiscAQ*, potentially hindering our ability to statistically detect a relation between cost of capital and *DiscAQ*. Noting these drawbacks, FLOS (2005) suggest that this approach provides a lower bound for the effects of *DiscAQ* on the cost of capital.

To ensure that our lack of significant findings on the effects of discretionary AQ on the cost of capital is not driven by measurement errors, we conduct two alternative estimation procedures as robustness checks. First, we control for the five innate factors (*SIZE*, $\sigma(CFO)$, $\sigma(Sales)$, *OpCycle* and *NegEarn*) by including them directly in models (3) and (4) (cost of debt and equity respectively), together with total AQ. Under this specification, the coefficient on *AQrank* captures the cost of capital effect attributable to the portion of AQ that is incremental to the effects captured by the innate factors (i.e., attributable to *DiscAQ*). FLOS (2005) refer to this as 'Method 2' and note that it provides an upper bound for the relation between discretionary AQ and costs of capital. While this approach does not allow statistical comparison on the differences between *InnateAQ* and *DiscAQ*, it does partially mitigate the econometric concerns noted above. The untabulated results under Method 2 do not change our inferences, and we continue to find that discretionary component of AQ is not significantly associated with the cost of debt and equity.

Our second sensitivity check involves estimating models (3) and (4) (cost of debt and equity respectively) using fixed- and random-effects models. The main advantage of using fixed- and random-effects models is that these two models control for all unobservable firm-specific characteristics that potentially contribute to innate AQ. In the spirit of FLOS (2005) Method 2 discussed above, we interpret the estimated coefficient on total AQ (*AQrank*), after considering fixed- and random-effects, as the effects of discretionary AQ on the costs of capital incremental to unobservable firm-specific innate factors that drive accruals quality. Untabulated results continue to show insignificant coefficients on total AQ, suggesting that the effects of discretionary AQ on the cost of capital are insignificant for Australian firms.

In summary, our conclusion that the effects of discretionary AQ on the cost of capital is negligible for Australia firms continue to hold after using both FLOS (2005)

²⁵ We thank an anonymous referee for highlighting these issues.

Method 2 and the fixed- and random-effects models. Although we cannot completely rule out concerns over measurement errors in the estimation of discretionary AQ, the additional tests above suggest our findings on the relation between discretionary AQ and cost of capital are robust.

6. CONCLUSIONS

Recent theoretical work argues that information risk is a non-diversifiable risk factor that is priced in the capital market (e.g., EOH, 2004; and LLV, 2007). Using accruals quality as a proxy for information risk, FLOS (2005) provide empirical support for this hypothesis. They show that US firms with poorer accruals quality exhibit higher costs of debt and equity capital. This study extends the analysis of FLOS by examining the interplay of accruals quality, information risk and cost of capital for a sample of Australian firms. While the work may be viewed as an out of (US) sample test of the effects of information risk on cost of capital, the Australian institutional and regulatory environment provides a number of subtle differences which potentially impact on the relation between accruals quality and cost of capital.

Australian firms rely heavily on private debt which is likely to both increase the precision of accounting information and lower the information asymmetry across lenders. To the extent that this mitigates information risk associated with opportunistic managerial reporting choices, the effect of discretionary reporting on cost of debt will be reduced. In addition, over the duration of our sample period, Australian firms are subject to a continuous disclosure regime that promotes the non-selective disclosure of timely and high quality information to the capital markets. This reduces information risk associated with opportunistic managerial reporting and disclosure choices and consequently the influence of discretionary reporting on the cost of equity. In such an environment, the cost of equity is more likely to be influenced by accruals quality arising from economic fundamentals (i.e., *InnateAQ*) rather than accruals quality arising from managerial reporting discretion (i.e., *DiscAQ*).

Our empirical results are largely consistent with these predictions. In Australian debt markets, the innate component of AQ exerts an economically significant influence on cost of debt. There is no association between discretionary AQ and cost of debt, consistent with the argument that the heavy reliance of Australian firms on private debt reduces information risk associated with managerial reporting choices, and consequently the influence of discretionary AQ on cost of debt. In the equity market, there is evidence that total AQ affects the cost of equity. However, this relation is driven mainly by the innate component rather than the discretionary component of AQ. This relation is consistent with the argument that the continuous disclosure regime mitigates the information risk resulting from opportunistic financial reporting choice and hence, the influence of discretionary AQ on the cost of equity.

The paper also contributes to the debate over whether information risk is a priced risk factor. CGV (2008) argue that the econometric approach of FLOS (2005) does not explicitly test the hypothesis that AQ is priced. Instead, they advocate the 2SCSR approach commonly-employed in the asset-pricing literature. Using the 2SCSR approach, our asset-pricing analysis provides support for the conjecture of FLOS (2005) that AQ (and more so for the innate component of AQ) is a priced risk factor. Our findings also complement those of Kim and Qi (2008) and Ogneva (2008), who show that AQ is a priced risk factor for US firms using the 2SCSR approach.

Overall, the empirical results in this paper support the recent theoretical arguments that information risk is a non-diversifiable risk factor which is priced by the capital markets. Reporting quality seemingly matters to capital-market participants and has important economic consequences for firms' cost of capital. These findings are likely to be of interest to regulators, standard setters, managers and investors who are interested in the quality of financial reporting in general.

REFERENCES

- Aboody, D., J. Hughes and J. Liu (2005), 'Earnings Quality, Insider Trading, and Cost of Capital', *Journal of Accounting Research*, Vol. 43, No. 5, pp. 651–73.
- Barth, M., Y. Konchitchki and W. Landsman (2006), 'Cost of Capital and Financial Statement Transparency', Working Paper (Stanford University).
- Bhattacharya, N., F. Ecker, P. Olsson and K. Schipper (2008), 'Direct and Mediated Association Among Earnings Quality, Information Asymmetry and the Cost of Equity', Working Paper (Southern Methodist University and Duke University).
- Black, F., M. Jensen and M. Scholes (1972), 'The Capital Asset Pricing Model: Some Empirical Findings', in M. Jensen (ed.), *Studies in the Theory of Capital Markets* (Praeger, New York).
- Brennan, M.J., A. Wang and Y. Xia (2004), 'Estimation and Test of a Simple Model of Intertemporal Capital Asset Pricing', *Journal of Finance*, Vol. 59, No. 4, pp. 1743–76.
- Campbell, J.Y. and T. Vuoleenaho (2004), 'Bad Beta, Good Beta', *American Economic Review*, Vol. 94, No. 5, pp. 1249–75.
- Chen, S., T. Shevlin and Y.H. Tong (2007), 'Does the Pricing of Financial Reporting Quality Change Around Dividend Changes?', *Journal of Accounting Research*, Vol. 45, No. 1, pp. 1–40.
- Cochrane, J.H. (2005), *Asset Pricing* (Princeton: Princeton University Press).
- Core, J., W. Guay and R. Verdi (2008), 'Is Accrual Quality a Priced Risk Factor?', *Journal of Accounting and Economics*, Vol. 46, No. 1, pp. 2–22.
- Cotter, J. (1998), 'Utilisation and Restrictiveness of Covenants in Australian Private Debt Contracts', *Accounting and Finance*, Vol. 38, No. 2, pp. 181–96.
- Dechow, P. and I. Dichev (2002), 'The Quality of Accruals and Earnings: The Role of Accrual Estimation Errors', *The Accounting Review*, Vol. 77 (Supplement), pp. 35–59.
- Easley, D. and M. O'Hara (2004), 'Information and the Cost of Capital', *Journal of Finance*, Vol. 59, No. 4, pp. 1553–83.
- Fama, E.F. and K.R. French (1993), 'Common Risk Factors in the Returns on Stocks and Bonds', *Journal of Financial Economics*, Vol. 33, No. 1, pp. 3–56.
- and J. MacBeth (1973), 'Risk, Return and Equilibrium: Empirical Tests', *Journal of Political Economy*, Vol. 81, No. 3, pp. 607–36.
- Francis, J., R. LaFond, P. Olsson and K. Schipper (2005), 'The Market Pricing of Accruals Quality', *Journal of Accounting and Economics*, Vol. 39, No. 2, pp. 295–327.
- Gibbons, M.R., S.A. Ross and J. Shanken (1989), 'A Test of the Efficiency of a Given Portfolio', *Econometrica*, Vol. 57, No. 5, pp. 1121–52.
- Gietzmann, M. and J. Ireland (2005), 'Cost of Capital, Strategic Disclosures and Accounting Choice', *Journal of Business, Finance & Accounting*, Vol. 32, Nos. 3&4 (April/May), pp. 599–634.
- Guay, W., S.P. Kothari and R. Watts (1996), 'A Market-Based Evaluation of Discretionary Accruals Models', *Journal of Accounting Research*, Vol. 34 (supplement), pp. 83–105.
- Hribar, P. and D. Collins (2002), 'Errors in Estimating Accruals: Implications for Empirical Research', *Journal of Accounting Research*, Vol. 40, No. 1, pp. 105–34.
- Hughes, J., J. Liu and J. Liu (2007), 'Information Asymmetry, Diversification, and Cost of Capital', *The Accounting Review*, Vol. 82, No. 3, pp. 705–29.
- Kaplan, R. and G. Urwitz (1979), 'Statistical Models of Bond Ratings: A Methodological Inquiry', *Journal of Business*, Vol. 52, No. 2 (April), pp. 231–61.
- Kim, D. and Y. Qi (2008), 'Accounting Information Quality, Stock Returns, and Macroeconomic Conditions', Working Paper (Korea University and Concordia University).

- Lambert, R.A., C. Leuz and R. Verrecchia (2007), 'Information Asymmetry, Information Precision, and the Cost of Capital', Working Paper (University of Pennsylvania and University of Chicago).
- McNichols, M. (2002), 'Discussion of The Quality of Accruals and Earnings: The Role of Accrual Estimation Errors', *The Accounting Review*, Vol. 77 (Supplement), pp. 61–9.
- O'Brien, M.A., T. Brailsford and C. Gaunt (2008), 'Size and Book-to-Market Factors in Australia', Working Paper (University of Queensland).
- Ogneva, M. (2008), 'Accrual Quality and Expected Returns: The Importance of Controlling for Cash Flow Shocks', Ph.D. Dissertation (University of Southern California).
- O'Hara, M. (2003), 'Presidential Address: Liquidity and Price Discovery', *Journal of Finance*, Vol. 58, No. 4, pp. 1335–54.
- Palepu, K., P. Healy and V. Bernard (2000), *Business Analysis and Valuation* (2nd ed., Southwestern Publishing).
- Petersen, M. (2008), 'Estimating Standard Errors in Finance Panel Data Sets: Comparing Approaches', *Review of Financial Studies* (forthcoming).
- Petkova, R. (2006), 'Do the Fama-French Factors Proxy for Innovations in Predictive Variables?', *Journal of Finance*, Vol. 61, No. 2, pp. 581–612.
- Shanken, J. (1992), 'On the Estimation of Beta-Pricing Models', *Review of Financial Studies*, Vol. 5, No. 1, pp. 1–33.
- Subramanyam, K.R. (1996), 'The Pricing of Discretionary Accruals', *Journal of Accounting and Economics*, Vol. 22, No. 1, pp. 249–81.