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Accruals Quality, Stock Return Seasonality, and the Cost of Equity Capital: International Evidence*

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Accruals Quality, Stock Return Seasonality, and the Cost of Equity Capital: International Evidence

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ABSTRACT

Mashruwala and Mashruwala (2011) argue that inconsistent earlier findings regarding whether accruals quality (AQ) is priced in equity markets (Core et al. 2008; Kim and Qi 2010) may be explained by seasonality in returns deriving from tax-loss selling. Finding no evidence of annual AQ premia for U.S. firms, Mashruwala and Mashruwala report that significant monthly premia concentrate in January, with the remainder of the year demonstrating negative or insignificant returns to AQ, and attribute this strong seasonality to tax-loss selling by investors, rather than information risk. However, the end of the tax year for U.S. investors coincides with the calendar year and the financial year for the majority of firms, which may suggest alternative explanations for seasonal variation in returns. We extend Mashruwala and Mashruwala's study, using an international sample including countries where incentives for tax-loss selling exist, but in which the standard tax and financial years differ (Japan and the UK), and where the tax and financial years conclude in a month other than December (Australia), as well as employing a longer U.S. sample. We find some evidence of an AQ premium in the United States, which although dominated by January returns, remains significant annually. However, these findings are sensitive to the inclusion of low price stocks and the choice of asset pricing test. In Japan, the UK and Australia we document consistent evidence that an AQ premium exists on average throughout the year, and in samples excluding the first month of the tax year. The sensitivity of our U.S. results to the January period may reflect the conflation of numerous seasonal influences on returns, not all of which necessarily reflect mispricing.

Keywords: accruals quality; stock return seasonality; low-priced returns; cost of equity capital

JEL Classification: M41; G14; G12

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Accruals Quality, Stock Return Seasonality and the Cost of Equity Capital: International Evidence

1. Introduction

Accruals quality (AQ) measures the extent to which a firm's reported accruals map to cash flows, and is employed in the literature as a proxy for information risk (e.g., Bharath et al. 2008; Kravet and Shevlin 2009). There has been considerable debate regarding whether AQ is priced in stock markets. Recently, Kim and Qi (2010) analyze annual abnormal returns to AQ-defined strategies, and find that AQ is a priced risk factor if low-priced stocks are excluded or controlled. However, Mashruwala and Mashruwala (2011; hereafter "MM") find no evidence that AQ is priced on average across the year, and that abnormal positive returns to AQ-based strategies occur exclusively in January (the first month of the tax year for individual investors), regardless of whether low-priced stocks are controlled. This leads MM to argue that the absence of both an annual AQ premium, and of returns to AQ outside the turn of the tax year, suggests that the apparent pricing of AQ reflects tax-loss selling of poor-AQ stocks in December, rather than compensation for information risk. MM suggest that contrary findings in prior research (e.g., Aboody et al. 2005; Kim and Qi 2010) may reflect time variation in the extent to which January premia reverse across the rest of the year, and the failure to account for this seasonality in asset pricing tests. Consequently, MM question the meaningfulness of research that employs AQ as an information risk proxy.

To further investigate whether the pricing of AQ is purely a seasonal effect that reverses across the remainder of the year, and whether any seasonality observed relates to tax incentives, we study the annual and intra-year pricing of AQ in the United States and in three other large market economies for which sufficient data exists to estimate AQ robustly (Australia, Japan and the UK). Importantly, there are incentives for tax-loss selling in each jurisdiction, but these countries differ both in their tax-year end dates and the alignment of the tax and standard financial reporting years. By studying jurisdictions with a range of tax year end dates, we are able to separate tax year effects from end of financial year or end of calendar year effects that may confound tests that rely purely on U.S. data (where the end of

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the tax year for individual investors coincides with the end of financial year for the majority of firms, the holiday period and generally lower trading volume).

While the primary focus of asset pricing tests concerns whether an AQ premium exists across the year, understanding the strength and nature of any seasonality observed helps us to better understand whether an observed premium is likely to reflect mispricing or risk. We examine samples comprising all calendar months and several subsamples that exclude returns earned in the first month of the tax year and / or financial reporting year to attempt to identify the source and nature of any observed AQ premium. By excluding the first tax month, we provide evidence regarding whether any observed AQ premium is singularly driven by tax-loss selling late in the prior tax year. Excluding returns earned in the first month of the financial reporting year controls for the possible impact of window-dressing by institutional investors seeking to remove "loser stocks" from their portfolios at balance date, and other factors varying around the turn of the financial year, such as the realization of information uncertainty or the likelihood of director trading.[1]

In the two sample countries where the tax year for individuals and the most common reporting year are distinct (Japan and the UK), we find consistent evidence of an annual AQ premium that persists outside the first months of tax and financial years. In Japan, we also find some evidence of abnormally high returns to AQ in the first month of the financial reporting year, consistent with non-tax related explanations, such as window-dressing by institutional investors. Taken together, results for these two countries suggest that an AQ

[1] The UK Companies Act 1985, for instance, prohibits director trades during the two month period preceding annual earnings announcements which, for the typical firm, would begin a few weeks before the end of the financial year (similar provisions are included in the Model Code of the London Stock Exchange (1977)). The Australian Stock Exchange (ASX) requires listed firms to have at least one such "blackout" or "closed" period, but does not mandate the timing of this no-trade window. However, the ASX states that it would generally expect an entity to include within its closed period "the period from, or just prior to, the close of books….until a reasonable period after the release of their financial results" (ASX Guidance Note 27, p.8).

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premium exists (in those markets) and that this premium is not attributable to tax-loss selling. We find similar evidence in Australia, where the standard tax and financial years end in June.

For the United States, we find evidence of a significant annual AQ premium in two of three asset pricing tests. However, we find no premium if either January returns (the first tax month and typically the first financial month), or returns to stocks with a low beginning-of-month price, are excluded from the sample. The fact we find less consistent evidence that AQ is priced in the United States outside the first tax month, however, aligns with findings from previous studies that other proposed risk factors (e.g., size and book-to-market) exhibit a strong January seasonality (Keim 1983; Loughran 1997), and is plausibly driven by the fact that January returns reflect multiple factors that may induce seasonality, some of which are not risk-related (e.g., tax-loss selling and window-dressing). January premia could also reflect greater demand from uninformed investors for the stock of firms that impose trading blackouts on their directors in the period between balance date and reporting date, or the gradual dissipation of information uncertainty early in the reporting year. Consistent with this possibility, our additional tests find evidence of an AQ premium in the first month of the financial reporting year of U.S. firms with June 30 and May 31 balance dates.

We conclude our analysis by investigating the impact of alternative controls for transaction costs. Kim and Qi (2010) find evidence of an annual AQ premium if stocks with both opening and closing prices below $5 are excluded or controlled. However, in a footnote, MM report that they find no evidence of an annual AQ premium if they apply a traditional ex ante price filter. As noted above, our U.S. price-restricted results are similar to those of MM. Because the Kim and Qi (2010) filter has the potential to severely bias returns upwards, we investigate the extent to which this filter rule explains the difference between their results and those of MM, and also examine the impact of alternative means of identifying stocks for which mispricing is most likely. We show that the filter used in Kim and Qi (2010) generates significant annual and January-specific AQ premia in all sample countries and all asset pricing tests, but that average returns are severely biased upwards. We also show that alternative filter rules, based on proxies for transaction costs applying throughout the month

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in which returns are estimated, but with less severe biases, also detect significant AQ premia in almost all countries and tests. We caution, however, that two of the three alternate filter rules reduce, rather than remove, the apparent bias in returns generated by the Kim and Qi (2010) filter.

Our study contributes to the extant literature by providing evidence that the documented AQ premium is not singularly driven by the tax-loss selling effect. Rather, the evidence is consistent with AQ being a priced risk factor, particularly in the case of our non-U.S. samples. We also contribute to the literature interested in the role of tax-loss selling in explaining seasonality more broadly, by presenting cross-jurisdictional evidence that suggests that the tax-loss selling explanation for seasonality in returns may be overemphasized. Finally, our additional tests identify the likely source of disagreement between the findings of MM and Kim and Qi (2010), viz. the latter paper's use of a low-price stock filter that severely biases returns upwards.

2. Prior research and theory

AQ is a construct introduced by Dechow and Dichev (2002), measured by the within-firm standard deviation of residuals from a regression of short-term accruals against past, current and prior realized cash flows. Firms whose accruals have a weak relationship with realized cash flows generate a high AQ score, and are described as having poor AQ. The AQ score has been used extensively in the accounting literature as a proxy for information risk, most commonly as a measure of information precision (e.g., Bharath et al. 2008; Ashbaugh-Skaife et al. 2009; Kravet and Shevlin 2009). Recent analytical literature predicts a positive relation between information risk and the cost of equity capital (Easley and O'Hara 2004; Yee 2006; Lambert et al. 2007; Lambert and Verrechia 2015). Easley and O'Hara (2004) develop a rational expectations model in which firms for which there is a greater proportion of private information experience greater information risk and higher required return, while Yee (2006) argues that poor earnings quality increases the cost of equity capital by magnifying the uncertainty about future dividend payments. Lambert et al. (2007) find that accounting quality influences firms' cost of capital by affecting investors' assessments of the

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distributions of firms' future cash flows and real production decisions. While each of the analytical studies discussed identify reasons why information precision should be priced, Lambert et al. (2007) and later studies (Lambert et al. 2012; Lambert and Verrechia 2015) also allow a role for information asymmetry to be priced, conditional on the existence of market imperfections. Centrally, Lambert et al. (2012) show that under perfect competition, only the average precision of investors information directly affects the cost of capital, as less-informed investors are able to infer private information from the traded price. Under imperfect competition, informed investors' demand is attenuated due to illiquidity, and private information is thus less fully impounded in price. This, in turn, reduces the average precision of uninformed investors' information and thus the cost of capital may also be affected by the distribution of information across investors. Related empirical research has identified the conditional pricing of information asymmetry on realized returns (Armstrong et al. 2011) and the implied cost of equity (Bhattacharya et al. 2012).

Empirical studies of the pricing of AQ report inconsistent findings. Francis et al. (2005) conduct a range of tests to investigate whether investors price AQ, including cross-sectional analyses of earning-to-price ratios, and market betas, from which they infer that AQ is priced. They also construct an AQ factor-mimicking portfolio and estimate time-series regressions of firms' realized abnormal returns on the AQ factor, and find a significant positive regression coefficient for this factor, which they interpret as indicating that AQ "plays an economically meaningful role in determining the cost of equity capital" (Francis et al. 2005, 315). Applying a similar approach with Australian data, Gray et al. (2009) report a significant positive coefficient for the AQ factor, driven by the innate rather than discretionary component of AQ.[2]

However, Core et al. (2008) argue that the method used by Francis et al. (2005) does not test whether AQ is priced. Instead, they use a two-stage cross-sectional regression

[2] The innate component of AQ is the fitted value from a regression of AQ against various firm fundamentals including size, book-to-market, operating volatility, growth and negative profits.

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technique (2SCSR), in which they first estimate time-series regressions of excess returns for each firm against the contemporaneous returns to the Fama-French three factors and the AQ factor, and then estimate a cross-sectional regression of firms' excess returns on the factor loadings estimated in the first stage. Core et al. report an insignificant coefficient for the AQ factor loading in the second stage, and thus conclude that AQ is not a priced risk factor. However, when Gray et al. (2009) apply this method to Australian data, they find that the AQ factor *is* significant in the second stage regression.

Aboody et al. (2005) report tests of the association between future excess returns and a hedge portfolio long in firms with high AQ betas and short in low AQ beta firms. Although the reported hedge returns are insignificantly different from zero, the authors interpret significant positive returns to the high AQ exposure leg of the portfolio as indicative that AQ is priced, attributing the insignificant hedge portfolio results to statistical noise contained in the low AQ exposure portfolio. Core et al. (2008) re-estimate these tests and, in fact, find a significant excess return to the AQ hedge over Aboody et al.'s sample period (1985-2003), but not if a longer sample period (1971-2003) is used.

Kim and Qi (2010) attempt to reconcile the conflicting findings in Francis et al. (2005) and Core et al. (2008) by controlling for the influence of low-priced stocks. Kim and Qi emphasize that asset pricing tests use realized returns as a proxy for *ex-ante* expected returns, because the latter cannot be observed directly. However, realized returns of low-priced stocks are biased due to liquidity effects and other transaction costs (Ball et al. 1995), which may distort the true relation between AQ and the cost of equity capital. Kim and Qi thus re-examine the relation between AQ and the cost of equity capital by replicating the method in Core et al. (2008, Tables 3 and 4), controlling for or excluding low-priced returns, which they define as the returns earned by stocks with consecutive month-end prices below $5. They find that AQ is priced in stock markets and that this pricing effect is correlated with proxies for fundamental risk. We note, however, that the method by which Kim and Qi identify low-priced stocks has the potential to bias subsequent asset pricing tests, because

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positive monthly returns are systematically more likely to be retained in their sample than are zero or negative returns. We explain this bias in detail in our method section.

Finally, MM study the pricing of AQ across 1971-2008, and the intra-year seasonality of returns to AQ therein. They argue that the stock of low AQ firms may be systematically more likely to exhibit mispricing due to tax-loss selling late in the year, and the attenuation of this selling pressure in January. MM find that there is, on average, no annual AQ premium (other than for the smallest quintile of firms); an intra-year AQ premium exists *only* in January, and not across the rest of the year; the observed January AQ premium concentrates in the first four trading days of that month; and the AQ premium increases with proxies for the likelihood of tax-loss selling around the turn of the year. The absence of an annual AQ premium, combined with the strong and apparently tax-related seasonality, lead MM to conclude that AQ is not a priced risk factor. While the January AQ premium occurs in almost all sample years, MM report that the extent of the reversal of this premium across other months varies significantly over time. Consequently, MM suggest the presence of strong but time-varying seasonality renders estimates of the annual AQ premium sensitive to the study period employed. In fact, MM report (MM, 1351) that their asset pricing tests *do* generate significant annual AQ premia if they limit their sample period to that studied by Aboody et al. (2005), and attribute the difference between this finding and the longer sample results as reflecting the long-term variation in the strength of seasonal effects. MM thus emphasize the importance of controlling for the impact of seasonality when testing whether AQ is priced, as it is "the outsized January premium that occasionally results in an annual premium being observed" (MM, 1351).

As MM note, the existence of strong January seasonal returns to a candidate risk factor does not necessarily imply that the candidate risk factor is not priced. From prior research we know that there is a strong January effect in total returns in the United States, and this return is associated with various proposed asset pricing factors including the market factor, the size factor, the book-to-market factor and the debt-related factor (Keim 1983; Blume and Stambaugh 1983; Tinic and West 1984; Fama and French 1993; Davis 1994;

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Loughran 1997).[3] If the observed seasonality of returns to a candidate risk factor is so strong that the average annual premium to a proposed factor is insignificant, and there is no theoretical reason why the proposed risk should concentrate in January, it is difficult to conceive of a risk-based explanation. However, other candidate risk factors have both strong January seasonal returns and a significant annual average return (e.g., exposure to market risk premium in Fama and French 1993; book-to-market factor in Loughran 1997). Such a combination of results is consistent with the proposed risk factor being priced, although a risk explanation is more convincing if seasonal variation in underlying risk can be established. As MM note, however, the strong seasonality observed may explain the difference in annualized premia reported in studies that do not control for seasonality, if the extent to which January returns reverse over the rest of the year varies across study periods.

Thus, like MM, we consider three core questions. First, is there evidence that AQ is priced, on average, throughout the calendar year? Second, is there significant seasonality in the returns to AQ? And finally, to what extent does any observed seasonality correlate with incentives for tax-loss selling, other indicators of mispricing, or seasonal variation in information risk? MM's ability to address these questions was, however, constrained by the U.S. institutional environment. The turn of the U.S. tax year coincides with both the turn of the calendar year (a time of generally quiet trade and low liquidity) and the turn of the financial year for most U.S. firms. Several papers argue that information risk is higher around the turn of the financial year (Rozeff and Kinney 1976; Keim 1983; Barry and Brown 1984;

[3] Keim (1983) and Blume and Stambaugh (1983) find that January returns account for more than half of the annual size premium, and more than half of this size premium concentrates in the first trading week in January. Davis (1994) finds that the book-to-market premium occurs only in January for the 1940-1963 period, while Loughran (1997) provides evidence that for all but the two smallest size quintiles, the book-to-market premium occurs only in the month of January for the 1963-1995 period. Fama and French (1993) provide evidence that market factor, the size factor, the book-to-market factor, and the debt-related factor are significant only in January. For non-January returns, these factors are not significant.

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Chan 1985; Kim 2006), as information asymmetry increases prior to the release of annual financial results. A more nuanced variant of the information risk argument is that, as the release of financial reports approaches, uninformed investors perceive an increased probability that the other side to any trade will have superior information (Ritter 1988), and consequently the uninformed investors price-protect by adjusting their bids until either the source of information uncertainty is resolved (Glosten and Milgrom 1985), or the perceived probability of facing an informed trader is otherwise reduced (as may be the case where firms apply "close periods" on director trading). The end of the standard financial year also coincides with the end of the reporting year for mutual funds, and has been argued to induce window-dressing by fund managers looking to clear loser stocks from their year-end portfolios (Lakonishok et al. 1991). This conflation of potential explanations for seasonality has the potential to obscure the nature of seasonal returns *and* increase the strength of any seasonality observed. By studying the pricing of AQ in countries where potential seasonality-inducing factors occur at different times of the year, we are able to shed greater light on the sources of seasonality, and the extent to which any apparent pricing of AQ is influenced by seasonality. In Japan and the UK, the turns of tax and financial years fall in different months, making it possible to disentangle the turn of tax year effect on AQ from that of the turn of financial year. In Australia, although the tax year aligns with the financial reporting year for most firms, this occurs on 30 June rather than 31 December, creating an opportunity to parse out turn of the calendar year related seasonality.

3. Taxation of stock transactions and incentives for tax-loss selling

MM posit that seasonality in the pricing of AQ is largely attributable to incentives for tax-loss selling that concentrate late in the tax year, and which disappear in the first month of the new tax year. The variation in tax law and tax year-end date across our international sample provides a setting suitable to investigate more thoroughly the source of seasonality in the AQ premium. Below we briefly describe the tax law that may provide incentives for tax-loss selling in each of the sample countries. As alternate explanations for seasonality may align with the financial reporting year, we also identify the standard financial reporting year in each country.

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In Australia, individuals pay tax on realized capital gains as part of their personal income tax. Capital losses can be used to offset capital gains in the same tax year, and can be carried forward and used to offset capital gains in future years. For shares held for more than 12 months, only 50 percent of the capital gain is taxable.[4] The Australian tax year ends on 30 June, so the tax-loss selling hypothesis implies that a July AQ premium should be observed. For the majority of Australian-listed corporations, the financial reporting year also ends on 30 June.

Prior to 2003, Japanese investors could choose between two options for paying tax on realized capital gains. The withholding tax option levied tax at 1.05 percent of gross proceeds, regardless of profit or loss on disposal. Alternately, individuals could elect to pay 26 percent of realized profit on transactions. In 2003, Japan implemented a new capital gains tax requiring a flat 10 percent tax on realized capital gains. Capital losses can be carried forward for three years and, from 2009, losses can be deducted from dividend income. As the tax year ends on 31 December, the tax-loss selling hypothesis suggests a January AQ premium. The financial reporting year for most listed Japanese firms ends on 31 March.

From 1992 to 2008, the UK's Taxation of Chargeable Gains Act (1992) required individuals to pay capital gains tax based on their marginal rate of income tax, which ranged from 0 percent to 40 percent. Individuals were entitled to an annual capital gains tax allowance and other forms of relief, including inflation indexation and taper relief, which provided for a reduction in taxable gains for assets that had been held for a certain length of time. Capital losses could be offset against capital gains in current and future periods. From 6 April 2008, a flat capital gains tax rate of 18 percent was introduced and taper relief was abolished.[5] As the tax year in the UK starts on 6 April for individuals and 1 April for corporations, the tax-loss selling hypothesis implies that an April AQ premium should be

[4] Prior to 1999, the original cost of shares could be inflation indexed, but no 50 percent discount was available.

[5] The tax rate has recently been increased to 28 percent for individuals subject to a marginal tax rate above the lowest threshold.

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observed. While approximately 20 percent of UK-listed firms have a financial reporting year coinciding with their tax year, the most common reporting-year end date is 31 December.

In the United States, the applicable tax rate depends on the period for which securities have been held. For short-term holdings (held less than 12 months), the full amount of any capital loss can be deducted from an investor's gross income,[6] implying a tax benefit equal to the investor's marginal tax rate (up to 35 percent). Concessional tax rates apply to investments held for longer than 12 months, but deductions are subject to annual loss limits. Losses beyond the deductible limit can be carried forward for use in future years. As the U.S. tax year for individuals ends on 31 December, the tax-loss selling hypothesis implies that a January AQ premium should be observed. For most listed U.S. firms, the tax year coincides with the reporting year.[7]

4. Data and methodology

Our U.S. sample spans 1971 to 2014, and includes all companies listed on the NYSE, AMEX and NASDAQ with available data. We use CRSP to obtain the monthly stock returns and price data, and COMPUSTAT for the annual accounting and shareholding data. The five Fama and French (2015) factors for U.S. firms and the risk-free rate are obtained from Kenneth French's website. Accounting data for other countries is drawn from Morningstar Equity Feed (Australia), COMPUSTAT Global (Japan) and Worldscope (UK). We use Datastream to obtain the stock return and price data and risk-free rates for these countries. For our non-U.S. samples, we first construct the size, book-to-market, profitability, and investment portfolios following closely the methods used by Fama and French (2015). Detailed portfolios construction procedures for each country are presented in the Appendix. We then estimate our size, book-to-market, profitability and investment factors based on the constructed size, book-to-market, profitability and investment portfolios. Our non-U.S.

[6] The immediate deductibility is subject to "wash sale" rules, which require that the security sold is not re-purchased in the next 30 days.

[7] In our sample, approximately 63 percent of U.S. observations have financial years ending in December. The next most common year-end for U.S. firms is June (8 percent).

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sample periods differ due to the availability of data. For Australia, direct cash flow from operations data is available from 1993, allowing the measurement of AQ from 1998 onwards. For Japan and the UK, we estimate accruals and cash flows indirectly because there is insufficient direct cash flow data available until very recent years. Our final samples, for which we have both stock return data and sufficient lagged accounting data, span October 1998 to September 2014 for Australia, January 2003 to June 2015 for Japan, and July 1993 to March 2014 for the UK.

Full vs. price-restricted samples

We test the impact of low-priced stocks on our asset pricing tests because the realized returns of these typically illiquid securities are biased proxies for investors' *ex-ante* expected returns. Ball et al. (1995) study the returns to contrarian portfolios formed at the end of December, and show that future positive returns to long positions in "loser" stock concentrate among low-priced stock, for which a price increase within the likely spread is sufficient to reduce future returns by 25 percent. Other research argues that the relatively high transaction costs and resulting biased returns of low-priced stocks contribute heavily to the January effect observed in U.S. stock returns (Bhardwaj and Brooks 1992). Since low-priced stocks have a greater than random probability of exhibiting negative annual returns, they may be systematically associated with the tax-loss selling that MM argue induces the January effect in AQ pricing.

To investigate the impact of low-priced returns on our main results, we estimate our realized stock returns tests using two samples: (1) the *full sample*, which includes all firms with available data during the sample period; and (2) the *price-restricted sample*, which excludes stocks with a low stock price at the beginning of the month (Ball et al. 1995; Jegadeesh and Titman 2001; Chan et al. 2006). A potential weakness of this traditional filter rule is that, where seasonality due to month-specific selling pressure is hypothesized, the filter may tend to retain negative returns in the "selling" month, and filter positive returns for the same firms in the following month. In our additional tests, we analyze the impact of using alternate filters for the impact of transactions costs/illiquidity.

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We follow prior studies and choose \$5 as the threshold for identifying low-priced U.S. stocks (Jegadeesh and Titman 2001; Chan et al. 2006; Kim and Qi 2010). For non-U.S. stocks, we identify low-priced stocks using price thresholds that cause similar levels of sample attrition to that induced by applying the \$5 threshold to U.S. stocks (which filters approximately 20 percent of firm-months in our price-restricted sample). To generate similar sample attrition for the other countries, we set threshold stock prices at AUD0.19 (Australia), ¥180 (Japan) and STG0.47 (UK).

Measurement of AQ

Following prior studies (Francis et al. 2005; Core et al. 2008; Kim and Qi 2010; Mashruwala and Mashruwala 2011), we use the modified Dechow and Dichev model (McNichols 2002) to estimate AQ. AQ captures the mapping of total current accruals into operating cash flows after controlling for changes in revenues and the level of fixed assets (*PPE*). All variables are deflated by average total assets. The model is described in equation (1):

$$TCA_{j,t} = \Phi_{0,j} + \Phi_{1,j}\,CFO_{j,t-1} + \Phi_{2,j}\,CFO_{j,t} + \Phi_{3,j}\,CFO_{j,t+1} + \Phi_{4,j}\,\Delta Rev_{j,t} + \Phi_{5,j}\,PPE_{j,t} + \varepsilon_{j,t}\,, \tag{1}$$

Where

$TCA_{j,t}$ = total current accruals of firm *j* in year *t*, estimated as: $\Delta CA_{j,t} - \Delta CL_{j,t} - \Delta Cash_{j,t} + \Delta STDEBT_{j,t}$ (for United States, Japan and the UK) or Net Operating Profit – Cash from Operating Activities (for Australia)

$\Delta CA_{j,t}$ = change in current assets (ACT)[8] of firm *j* between year *t*–1 and *t*;

$\Delta CL_{j,t}$ = change in current liabilities (LCT) of firm *j* between year *t*–1 and *t*;

$\Delta Cash_{j,t}$ = change in cash (CHE) of firm *j* between year *t*–1 and *t*;

$\Delta STDEBT_{j,t}$ = change in debt in current liabilities (DLC) of firm *j* between year *t*–1 and *t*;

$CFO_{j,t}$ = $NIBE_{j,t} - TA_{j,t}$ = cash flow from operations of firm *j* in year *t*;

$NIBE_{j,t}$ = net income before extraordinary items (IB) of firm *j* in year *t*;

$TA_{j,t}$ = $\Delta CA_{j,t} - \Delta CL_{j,t} - \Delta Cash_{j,t} + \Delta STDEBT_{j,t} - DEPN_{j,t}$ = total accruals of

[8] COMPUSTAT variable names are reported in parentheses.

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firm j in year t;

$DEPN_{j,t}$ = depreciation and amortization (DP) of firm j in year t;

$\Delta Rev_{j,,t}$ = change in revenues (SALE) of firm j between year t–1 and t; and

$PPE_{j,t}$ = gross value of property, plant and equipment of firm j in year t.

Consistent with prior studies, we winsorize current accruals and each regressor at the first and ninety-ninth percentiles (Francis et al. 2005; Kim and Qi 2010). Equation (1) is estimated within country-industry-year-specific samples because the meaningfulness of the DD model relies on the relative homogeneity of firms' operations and operating environments.[9] We then use the standard deviation of the firm-year specific residuals $\varepsilon_{j,t}$ from year t-4 through year t as our AQ score. A lower standard deviation of the residuals implies that the mapping of accruals to cash flow is stronger, indicating superior AQ. Thus, AQ is inversely related to the measured AQ score.

Portfolio construction

In each month, we form ten AQ-sorted portfolios based on the most recently measured AQ score, which are based on information assumed to be available to the public three months after the firms' fiscal year end. For example, for a firm whose financial year ends in December, monthly returns from April of year t+1 to March of year t+2 are assigned into one of ten decile portfolios according to the AQ score calculated in year t.

Asset pricing test

To test whether AQ is a priced risk factor, we use three approaches employed in relevant earlier literature: (1) estimation of the abnormal returns to an AQ hedge portfolio, (2) estimation of Fama-Macbeth cross-sectional regressions of returns against the AQ variable

[9] To estimate AQ firms must have at least 7 years accounting data. We further require at least 15 firms per industry-year. We use Fama and French's (1997) 48-industry classification for U.S. firms. Australian firms' industry classifications are obtained from data in the CRIF file, which are based on the former 2-digit ASX Industry codes. For Japan and the U.K., industries are defined using the Industry Classification Benchmark (ICB). We use the first two digits of the ICB to divide firms into 19 industries for Japanese data, but only the first digit of the ICB to divide firms into 10 industries for the U.K. data, due to the lower sample size.

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and other risk factors to test whether the AQ variable is able to explain the cross-section of average returns, and (3) regressions of estimated factor betas on the cross-section of returns (two-stage cross-sectional regressions). These tests are described below.

Abnormal returns to an AQ hedge portfolio

Following MM, to estimate the AQ hedge portfolio abnormal returns, we first calculate the monthly returns to an AQ hedge portfolio ($R_{p,t}$), which is long in the highest AQ decile portfolio (poorest AQ firms) and short in the lowest AQ decile portfolio (strongest AQ firms). Following prior studies (Francis et al. 2005; Core et al. 2008; Kim and Qi 2010; Armstrong et al. 2011), we weight firms equally in calculating the monthly AQ portfolio returns. We do not compute the value-weighted portfolio returns since returns calculated under this method can be dominated by a few large stocks, and our purpose is to test the expected returns for an average stock.

We then regress the monthly AQ hedge portfolio returns against the five factors used in Fama and French (2015): the market, size, book-to-market, profitability and investment factors. The intercept (Jensen's α) from this time-series regression is used to test whether AQ is priced, e.g., a significant positive α indicates that AQ is priced. The model is described in equation (2):

$$R_{p,t} = a_{0,p} + B_p(R_{m,t} - R_{f,t}) + s_p SMB_t + h_p HML_t + r_p RMW_t + c_p CMA_t + \varepsilon_{p,t} \quad (2)$$

Where

$R_{p,t}$ = return on the AQ hedge portfolio for month t;
$R_{m,t}$-$R_{f,t}$ = value-weighted market return less risk-free rate for month t;
SMB_t = the return on the size factor-mimicking portfolio in month t;
HML_t = the return on the book-to-market factor-mimicking portfolio in month t;
RMW_t = the return on the operating profitability factor-mimicking portfolio in month t; and
CMA_t = the return on the investment factor-mimicking portfolio in month t.

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Cross-sectional regressions of returns against risk factors

We next employ the cross-sectional approach of Fama and MacBeth (1973), regressing each stock's monthly returns on the estimated AQ value, market beta, firm size, book-to-market ratio, operating profitability and cash flow volatility. In addition to the potential risk factors traditionally identified, we control for a firm's operating profitability because Ball et al. (2015) provide evidence that operating profitability predicts expected returns, and this measure is likely correlated with AQ. As AQ is also likely to be correlated with a firm's operating volatility, which may affect the expected returns, we also include a firm's cash flow volatility (CV) as an additional control variable. Our cross-sectional regression model is as per equation (3):

$$R_{it} = \alpha_0 + \alpha_1 AQ_{it-1} + \alpha_2 BETA_{it-1} + \alpha_3 ln(SIZE)_{it-1} + \alpha_4 ln(BM)_{it-1} + \alpha_5\, OP_{it-1} + \alpha_6 CV_{it-1} + \varepsilon_{it} \quad (3)$$

Where

R_{it}	=	firms' monthly stock return;
AQ	=	firms' accruals quality, estimated using equation (1);
$BETA$	=	firms' market beta derived from the CAPM (using value-weighted market return. For U.S. firms, we use the 60-month rolling beta; for other countries, we use the 36-month rolling beta, each for windows ending in the month prior to the current month);
$SIZE$	=	firms' market value of common equity;
BM	=	firms' book-to-market ratio. Book value is calculated at the previous fiscal year-end and market value is calculated on the last trading day in the previous fiscal year;
OP	=	firms' operating profit, calculated as revenue minus cost of goods sold and selling, general and administrative expense (excluding expenditures on research and development) deflated by the book value of total assets; and

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CV = Firms' cash flow volatility, estimated as the standard deviation of the firm's rolling seven-year cash flow from operations deflated by book value of total assets.

Two-stage cross-sectional regressions

In addition to the above methods employed by MM, we follow Core et al. (2008) and Kim and Qi (2010) and use a two-stage cross-sectional regression technique (2SCSR) to test whether AQ is priced. The 2SCSR estimates portfolio betas in the first stage, and assigns the estimated betas to individual stocks in the second stage. We first rank all firms into 10 size portfolios according to their market value at the end of the sixth month after the standard financial reporting year in each country. Within each size decile portfolio, firms are further ranked into 10 portfolios based on their market-to-book ratio at the end of year t–1 (December) for the United States and the UK, and at the end of financial year *t* for Australia (June) and Japan (March). Thus, we obtain 100 portfolios for which equal-weighted portfolio returns are calculated. Portfolios are re-balanced every year. We then estimate betas by conducting a single time-series regression for each portfolio using the model described by equation (4):

$$R_{p,t} - R_{F,t} = \alpha_0 + \beta_{p,\,RM\text{-}RF}\,(R_{M,t} - R_{F,t}) + \beta_{p,\,SMB}SMB_t + \beta_{i,HML}HML_t + \beta_{i,AQ\,factor}\,AQfactor_t + \beta_{p,RMW}RMW_t + \beta_{i,CMA}CMA_t + \varepsilon_t, \tag{4}$$

where AQ$factor_t$ is the monthly return to the accruals quality factor–mimicking portfolio, calculated as the difference between the monthly equal-weighted portfolio returns of firms in the poorest two AQ quintiles and firms in the best two AQ quintiles and $R_{p,t}$ is the monthly equal-weighted portfolio returns. All other variables are defined as in equation (2).

In the second stage, we assign the estimated betas from the first stage to the individual stocks in each portfolio to increase the test power (Fam a and French 1992; Kim and Qi 2010). We then conduct a cross-sectional regression as per equation (5):

$$R_{pi,t} - R_{F,t} = \lambda_0 + \lambda_1\hat{\beta}_{pi,RM-RF} + \lambda_2\hat{\beta}_{pi,SMB} + \lambda_3\widehat{\beta}_{pi,HML} + \lambda_4\widehat{\beta}_{pi,AQfactor} + \lambda_5\widehat{\beta}_{pi,RMW}$$

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$$+\lambda_6 \widehat{\beta}_{pi,CMA} + \lambda_7 Low\_priced + \mu_{pi,t}\,, \tag{5}$$

where $\hat{\beta}_{pi}$ is firm *i*'s beta for a specific risk factor estimated from the first-stage time-series regression, and *Low_priced* is a dichotomous variable indicating the presence of a firm's returns with opening stock price less than the country-specific thresholds detailed earlier. The time-series average of the estimated λ_i from equation (5) is the estimated risk premium of the corresponding risk factor. If the time-series average of the estimated λ_4 is statistically significantly positive, it implies that AQ is a priced risk factor.

While the two-stage cross-sectional regressions are standard tests used in the finance and accounting literatures to examine whether a proposed factor is priced, previous studies have showed that the resulting estimated coefficients in the second-stage often (and sometimes strongly) reject standard asset pricing factors. For example, Jagannathan and Wang (1996), Petkova (2006), Core et al. (2008), and Kim and Qi (2010) report a (significant) negative regression coefficient for the factor loading on market beta.[10] Consequently, we caution against placing excessive emphasis on results generated from this test, but include it in our analysis for consistency with related prior research (Core et al. 2008; Kim and Qi 2010).

5. Results

Summary statistics

Table 1 reports descriptive statistics for the main variables used in our study. For Japan, we obtain 17,070 firm-year observations (2,396 firms) of AQ, which has a mean (median) value of AQ of 0.024 (0.019). For the UK, we obtain 11,170 firm-year observations (1,237 firms) and a mean (median) value of AQ of 0.054 (0.041). For Australia, we obtain 4,888 firm-year observations (751 discrete firms) of AQ over the period 1998 to 2013. By comparison, Gray et al. report a sample of 2,057 firm-year observations (509 firms) over the period 1998 to

[10] The regression coefficients on the *SMB* and *HML* factor beta are also inconclusive. For example, Kim and Qi (2010) document a significant negative regression coefficient on *SMB*, while Petkova (2006), Core et al. (2008) and Cremers et al. (2009) present evidence that the regression coefficient for *SMB* is insignificant. Cremers et al. (2009) show that the regression coefficient for *HML* is insignificantly different from zero.

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2006. Our mean (median) value of AQ is 0.091 (0.072), marginally larger than those provided by Gray et al. (2009), who report a mean (median) AQ value of 0.081 (0.064). For the United States, we obtain 147,829 firm-year observations (12,726 discrete firms). The mean (median) AQ score is 0.062 (0.083), which is similar to those reported by other U.S. studies (Francis et al. 2005; Core et al. 2008; Kim and Qi 2010; Mashruwala and Mashruwala 2011).

[Insert Table 1 Here]

AQ hedge portfolio abnormal returns

We now report results of tests of whether firms ranked in the poorest AQ decile outperform those in the best AQ decile by regressing the monthly AQ hedge portfolio return (return of AQ10–return of AQ1) on the five Fama-French factors. A significant positive intercept ($\alpha_{0,p}$) would indicate that firms with poorer AQ earn higher risk-adjusted returns than do firms with superior AQ, consistent with AQ being a priced risk factor. To examine the extent to which any observed AQ premium is potentially affected by tax-loss selling and other seasonal effects, we estimate the regressions using alternate samples: (1) all calendar months, (2) all calendar months except the first month of the tax year, and (3) all calendar months except the first month of the financial year. For Australia and the United States, samples (2) and (3) are identical, as the standard tax and financial years coincide. For comparison with MM, we also tabulate January-specific results for the U.S. sample.

Table 2 shows the average monthly abnormal returns to an AQ hedge portfolio based on AQ deciles (AQ10–AQ1). To calculate the portfolio returns, we require each portfolio to include at least 20 stocks.

[Insert Table 2 Here]

Table 2, panel A reports the monthly abnormal returns to an AQ hedge portfolio, using the full samples. For both Japan and the UK, the annual AQ hedge portfolio abnormal returns are significantly positive at the 1 percent level (Japan: α = 1.0 percent, t-stat = 2.80; UK: a = 1.5 percent, t-stat = 3.74), indicating that firms in the poorest AQ decile earn

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significantly higher average risk-adjusted returns across the year compared to firms in the best AQ decile. For these countries, the documented AQ premium persists outside the first month of the tax year (Japan: α = 1.0 percent, t-stat = 3.01; UK: a = 1.5 percent, t = 3.76) and the first month of the financial year (Japan: α = 0.9 percent, t-stat = 2.70; UK: α =1.8 percent, t-stat = *3.54*), suggesting that the AQ premia does not singularly reflect mispricing due to tax-loss selling or window dressing by fund managers. In fact, in these two countries in which the tax and financial years are separate, the monthly AQ hedge portfolio abnormal returns are marginally greater outside the turn of the tax year than when measured across all calendar months. For Japan, the monthly abnormal returns increase from 0.96 percent to 1.02 percent (t-stat increases from 2.80 to 3.01), for the UK, the monthly abnormal returns increase from 1.47 percent to 1.54 percent (t-stat increases from 3.74 to 3.76). For Australia, there is also a significant premium to the AQ hedge portfolio, both across the year, and outside the first tax (financial) month (all months: α = 0.9 percent, t-stat = 3.23; non first month of tax (financial) year: α = 1.0 percent, t-stat = 3.20). Thus, in all non-U.S. countries we find evidence of a significant AQ premium across the year, and that this premium persists outside the first tax month.

For the United States, we observe a significant average AQ premium across all calendar months, but no AQ premium is documented outside the first month of the tax (financial) year (all months: α = 0.5 percent, t-stat = 2.52; non first tax (financial) month: α = –0.1 percent, t-stat = –0.45). There is a significant AQ premium in January (α = 5.3 percent, t-stat = 6.78) as well as in February (untabulated: α = 0.95 percent, t-stat = 1.97) and May (untabulated: α = 0.69 percent, t-stat = 1.71). While the strong January seasonal returns are suggestive of a tax-loss selling or other seasonal effects, the existence of a significant average annual premium is consistent with AQ being priced.

Table 2, panel B presents the monthly abnormal returns to an AQ hedge portfolio based on AQ deciles (AQ10–AQ1) using the price-restricted sample. Results show that when low-priced stocks are filtered, the AQ hedge portfolio abnormal monthly returns are still significantly positive in Japan and the UK, for returns observed in all months (Japan: α = 0.4

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percent, t-stat = 2.01; UK: α = 0.7 percent , t-stat = 1.83), and the documented AQ premium persists outside the first month of the tax year (Japan: α = 0.5 percent, t-stat = 2.43; UK: a = 0.8 percent, t = 1.9), and outside the first month of the financial year (Japan: α = 0.4 percent, t-stat = 1.89; UK: α = 0.9 percent, t-stat = 1.81). For Australia, there remains a significant AQ premium observed across all months (α = 0.5 percent, t-stat = 1.68) and when the first tax (financial) month is excluded (α = 0.6 percent, t-stat = 1.99).

For the United States, however, no AQ premium is documented, either annually or when the first tax (financial) month is excluded, if stocks with beginning-of-month price below $5 are removed from the sample. We do, however, observe a significant AQ premium in January (α = 1.5 percent, t-stat = 3.35). The contrast between these results and those in the full sample tests is consistent with theory arguing that the January effect is heavily influenced by mispricing due to high transaction costs and illiquidity (e.g., Bharwaj and Brooks 1992), and that this mispricing may contribute to the observed AQ premium. However, the exclusion of low priced stocks also greatly reduces the variation in AQ (more than 50 percent of firms in the AQ10 portfolio are filtered), and improves the average AQ of the retained sample, potentially reducing the likelihood of detecting a pricing effect. Further, the use of an ex ante price filter in cases where month-to-month variation in selling pressure is predicted may tend to retain within the sample firms with negative returns in the 'selling' month but exclude the same firms in the following month when prices recover. We discuss this possibility more fully in our additional tests.

In summary, our AQ hedge portfolio tests find evidence that AQ is priced on average across the year in Japan, the UK and Australia in both the full and price-restricted samples, and that these results hold outside the turn of the tax year and the turn of the financial year. Thus, in these countries an annual premium exists and does not appear to be significantly influenced by tax-loss selling behavior. For the United States, the full sample results show a significant AQ premium across the year. While this premium is strongly concentrated in January, the significant annual premium is consistent with AQ being priced on average. This juxtaposition of results is similar to that reported for other risk factors proposed in prior

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research (e.g., the size and book-to-market factors, see Fama and French 1993, for example). However, no AQ premium is observed in the United States for the price-restricted sample, raising the possibility that the observed U.S. AQ premium reflects mispricing rather than risk.

Cross-sectional regressions of returns against risk factors

We now report results of tests based on the single-stage cross-sectional approach of Fama and MacBeth (1973) and regress each stock's monthly returns on the estimated AQ score, market beta (*BETA*), firm size (*SIZE*), book-to-market ratio (*BM*), operating profitability (*OP*), and cash flow volatility (*CV*) as per equation (3).

Table 3, panel A shows the full sample results for all months, and all months other than the first tax (financial) month. In Japan and the UK, the regression coefficient for AQ is significantly positive across all calendar months (Japan: β = 0.263, *t*-stat = 3.87; UK: β = 0.109, *t*-stat = 1.75), when the first tax month is excluded (Japan: β = 0.242, *t*-stat = 3.53; UK: β = 0.118, *t*-stat = 1.74), and when the first financial month is excluded (Japan: β = 0.259, *t*-stat = 3.52; UK: β = 0.115, *t*-stat = 1.67). For Australia, a significant positive AQ premium is observed across all calendar months (β = 0.056, *t*-stat = 2.72), and when the first tax (financial) month is excluded (β = 0.064, *t*-stat = 2.92).

For the United States, a positive AQ premium is observed across all calendar months (β = 0.031, *t*-stat = 2.29), but no AQ premium is observed when January is excluded from the sample (β = 0.006, *t*-stat = 0.47). We note that, for the United States, once January returns are excluded from the sample, the size premium also disappears (all months: β = –0.0013, *t*-stat = –3.81; non-first tax (financial) month: β = –0.0005; *t*-stat = –1.39), and the book-to-market premium decreases from 0.004 to 0.003 (*t*-stat decreases from 5.63 to 4.52), consistent with findings from previous studies that there is a strong January size and book-to-market seasonal effect (Keim 1983; Blume and Stambaugh 1983; Davis 1994; Loughran 1997). Saliently, the annual AQ premium in the United States remains significant.

Consistent with findings from previous studies (Fama and French 1992; Chan et al. 1996; Lewellen and Nagel 2006; MM 2011; Asness 2012; Frazzini 2014), the regression

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coefficient for market beta (*BETA*) is insignificantly different from zero across all calendar months for all of these four countries.[11] The regression coefficient for firms' size (*ln(SIZE)*) across all months has the predicted sign and is significant at the 1 percent level for Japan, Australia, and the United States, but insignificant for the UK, consistent with previous findings (Chan et al. 1996; Fletcher 1997). Our regression coefficient for book-to–market (*ln(BM)*) has the predicted sign for all four countries and is significantly positive at the 1 percent level for Japan, the UK, and the United States, and at the 10 percent level for Australia. The regression coefficient for operating profit (*OP*) across all months is significantly positive for Japan and the United States at the 1 percent level (Japan: β = 0.091, t = 5.18; US: β = 0.022, t = 6.68), but insignificantly different from zero for Australia and the UK (Australia: β = 0.004, t = 0.45; UK: β = 0.003, t = 0.18). The regression coefficient for cash flow volatility (*CV*) across all months is insignificantly different from zero for all countries.

[Insert Table 3 Here]

Table 3, panel B reports results for Fama-Macbeth cross-sectional regressions estimated on price-restricted samples. For Japan, the UK and Australia the exclusion of low-priced stocks generates substantively similar AQ premia to those based on the full sample for returns across all calendar months, all calendar months except the first tax month, and all calendar months except the first financial month. For the United States, the coefficients for AQ are not significantly different from zero across all calendar months, but a significant January AQ premium (β = 0.140, t = 2.65) is observed. Once more, the absence of a U.S. AQ

[11] It is argued that borrowing constraints may lead to the overweighting of high-beta asset and underweighting of low-beta stocks, which in turn may explain the documented insignificant coefficient for market beta (Black 1972; Frazzini et al. 2014). Recently, Hong and Sraer (2016) provide an analytical model showing that high-beta assets can be overpriced relative to low-beta assets if investors' disagreement about the stock market's prospects is high. This overpricing of high-beta asset leads to the negative relation between stock returns and their market beta.

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premium for the price-restricted sample could reflect a reduction in the impact of mispricing, reduction in the variation in AQ in this sample, and / or the tendency of the filter rule to retain negative returns in months where selling pressure is hypothesized, while filtering subsequent positive returns.

Two-stage cross-sectional regressions

We now report the results of tests using two-stage cross-sectional regressions. Table 4, panel A presents the average estimated coefficients from the first stage. For all calendar months, the regression coefficient of AQ factor is significantly positive at the 1 percent level for all four countries (Japan: Coef. = 0.237, *t*-stat = 3.2; UK: Coef. = 0.101, *t*-stat = 3.14; Australia: Coef. = 0.072, *t*-stat = 2.37; US: Coef. = 0.641, *t*-stat = 8.92). Similar results are observed when the first tax and/or financial month are excluded from the sample.

The regression coefficients of *RM-RF*, *SMB* and *HML* have the normal signs and all are significantly different from zero at the 1 percent level, consistent with findings from previous studies (Francis et al. 2005; Core et al. 2008; Gray et al. 2009). The regression coefficients of *RMW* is significantly negative for Japan, the UK and Australia, but significantly positive for the United States. A deep investigation of why *RMW* behaves differently in our international sample relative to the United States is beyond the scope of this paper. However, we note from Fama and French (2015) that the directional impact of *RMW* on returns is strongly conditional on *HML* (see Fama and French 2015 Table 7, panel B). Strong correlations between these (and other) factors may give rise to erratic behavior in estimated coefficients.[12] Finally, we find that the coefficients on *CMA* are not significantly different from zero for Japan and the UK, but are significantly negative for both the United States and Australia.

[Insert Table 4 Here]

[12] In Australia, the U.K. and the U.S., *RMW* has correlations of above 20 percent with at least 3 other factors. In Japan, *RMW* has correlations with *SMB* and *CMA* of more than 50 percent.

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We next conduct cross-sectional regressions of individual firms' excess returns on factor betas estimated in the first-stage regression. Table 4, panel B presents the results of the second stage regression. To test the sensitivity of the results to low-priced returns, we estimate the second stage regressions both with and without a dichotomous control indicating that the subject firm has a stock price less than ¥180 (Japan), £0.47 (UK), $0.19 (Australia), or $5 (U.S.) at the beginning of month. The average estimated coefficients of the AQ factor beta are significantly positive for Japan, the UK and Australia, regardless of whether low-priced returns are controlled, and these returns persist outside both the first tax month, and first financial month of the year (Full Sample – Outside First Month of Tax Year: Japan: Coef. = 0.007, *t*-stat = 2.29; UK: Coef. = 0.010, *t*-stat = 2.80; Australia: Coef. = 0.013, *t*-stat = 3.77). The estimated coefficient for the U.S. AQ factor beta, however, is not significantly different from zero across the year for the full sample, and even becomes significantly negative at the 10 percent level when low-priced returns are controlled (Full sample: Coef. = –0.001, *t*-stat = –0.72; Price-restricted sample: Coef. = –0.002, *t*-stat = –1.81). This result is in stark contrast to that reported in the equivalent test in Kim and Qi (2010), who find a significant AQ premium (Coef. = 1.95, *t*-stat = 9.25, Kim and Qi 2010 Table 5) if low-priced stocks are defined as those with an opening *and* closing price less than $5. When January is excluded from the sample, the estimated coefficient for the AQ factor beta is significantly negative for the full sample and price-restricted sample (Full sample: Coef. = –0.005, *t*-stat = –3.55; Price-restricted sample: Coef. = –0.004, *t*-stat = –3.48). The estimated coefficient for the U.S. AQ factor beta is significantly positive in January for both the full sample and the price-restricted sample (Full sample: Coef. = 0.037, *t*-stat = 5.49; Price-restricted sample: Coef. = 0.030, *t*-stat = 5.49).

Table 4, panel B also shows the estimated regression coefficients for other risk factor betas. As noted in Section 3, this method frequently rejects commonly accepted risk factors included in the regressions. While our regression coefficients for the *HML* factor beta are positive and significant in all models across all countries, the evidence in favor of other control risk factors is less consistent with theory. The estimated coefficients of the market

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factor beta are insignificant in all regressions based on the Japanese and UK samples. For Australia, coefficients for the market factor beta are significantly negative across all calendar months, and this coefficient is significantly negative in all U.S.-based regressions. The average annual regression coefficient for the *SMB* factor is only positive and significant in Australia. For the United States, the regression coefficients for *SMB* are insignificantly different from zero across the year for the full and price-restricted sample, and become significant (but in the wrong direction) once January is excluded from the sample. The regression coefficients for the *RMW* factor beta are insignificant for Japan, the UK, and the United States under all models, but are significantly negative in Australia. While the regression coefficients of *CMA* factor beta are significantly positive for the UK, Australia, and the United States, the first stage coefficients are either insignificant or negative.

In summary, the results from the two-stage cross-sectional regressions show evidence consistent with AQ being a priced risk factor in Japan, the UK and Australia. There is no evidence on an AQ premium in the United States. Given that the regression coefficients on most other factor betas are also either insignificant or significant but in the opposite direction to that hypothesized (e.g., coefficients on the market factor beta, *SMB* factor beta and *RMW* factor beta), we urge caution in evaluating evidence drawn from this, albeit standard, asset pricing test.

We summarize the results of our three asset pricing tests in Table 5. Across the three tests, we find evidence of the pricing of AQ across the year, outside of the first tax months, and outside of the first financial months in Japan, the UK, and Australia regardless of whether low-priced stocks are controlled. In the United States, two of three tests detect significant annual AQ premia for the full sample, but no AQ premium is observed in the 2SCSR test. When low-priced stocks are excluded (controlled), we detect no annual AQ premium for U.S. firms.[13] We do, however, observe a significant January AQ premium in all of the three tests.

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[13] Our results are not sensitive to modest variations in the price thresholds selected (from AUD0 to AUD 0.35 for Australian firms, from ¥0 to ¥220 for Japanese firms and from STG0 to STG0.48 for UK firms). For U.S.

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[Insert Table 5 Here]

Alternate filters for transactions costs

While the results for our pricing tests in Japan, the UK and Australia are robust to the exclusion of stocks based on an ex ante price filter, the annual AQ premium detected in two of three U.S. tests disappears if such a filter is applied. Kim and Qi (2010) detect an annual AQ premium in the United States if stocks with *both* an opening and closing price below $5 are excluded. However, this filter mechanically biases returns upwards, because firms with opening prices below $5 which thereafter experience a positive monthly return have a chance of being retained in the sample, but those with zero or negative monthly returns are excluded from the sample. If low AQ firms are over-represented in the observations affected by this bias, then the measured returns to an AQ-based test variable will be biased upwards. Notwithstanding this, attempting to control for transaction costs in terms of their likely effect on measured returns *across the month* appears to be sensible, because filters based purely on ex ante prices may tend to retain negatively biased returns in cases where there are seasonal changes in selling pressure. Consider Bhardwaj and Brooks' (1992) explication of the positive bias in January returns for low-priced stocks, whereby excess selling pressure affecting stocks with high transactions costs causes late December trades to be more likely to be settled at or near the bid price, and this imbalance is corrected in early January when the selling pressure dissipates. Using data on the average relative spread (6.09 percent) for a sample of stocks with January opening prices below $5 and the proportion of daily closing prices recorded at or below (above) the bid (ask) price, Bhardwaj and Brooks' (1992) show that measured returns for windows extending from the last day of December to the each of the first five trading days in January are biased upwards by 1.01 percent, 1.18 percent, 1.42

firms, the application of a $1 low price stock filter does not generate a significant premium across our full sample period, but does detect an annual AQ premium if the sample is restricted to that of Aboody et al. (2005), which spanned 1985-2003.

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percent, 1.23 percent and 0.8 percent respectively.[14] However, the same transaction cost related factors that bias January returns upwards, may also contribute to the price decline in December. If end of November closing prices reflect the bid and ask price with equal probability, but late December trades tend to reflect the bid price (as per the 27 percent excess probability of a bid transaction in Bhardwaj and Brooks 1992, Table 4), the estimated bias for December returns is approximately -0.7 percent. For any ex ante price filter, there will be a subset of firms that commence December with a price above the $5 threshold, but finish the month below that threshold due to selling pressure that causes a price decline "within the spread". These firms will be retained in the sample in the month in which returns are negatively biased, but excluded in the following month in which the positive bias is expected.

To assess the impact of the filter employed, in Table 6 we report summary results for tests using alternate filters for the impact of transactions costs: the Kim and Qi (2010) filter – excluding stocks with both opening and closing prices below $5, a filter that excludes stocks with closing prices below $5, a modification of the Kim and Qi (2010) filter that excludes firms if *either* the opening or closing price below $5, and a filter based on the dollar volume of stock traded.

[Insert Table 6 Here]

Table 6 reports test coefficient and *t*-stats for each of our three tests, using the alternate filter rules. To help assess the impact of each filter rule on average returns, in the subheading for each panel, we report the difference in mean returns between observations filtered and those retained. All test coefficients were significant when using a traditional ex ante price filter in the non-U.S. countries (see Table 5, panel B), and remain significant under

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[14] The bid-ask bias in measured returns is estimated as $[s^2(1 - pq) / (4 - s^2) + 2s(p - q) / (4 - s^2)]$, where s is the relative spread, p is the excess likelihood of bid (relative to ask) transactions embedded in the price at the end of the previous return period, and q is the excess likelihood of bid transactions on the last day of return window (Bhardwaj and Brooks 1992, 572).

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all of the alternative filters reported in this table. Thus, we focus our discussion on the U.S. results. Whereas tests using the U.S. ex ante price filter sample failed to detect an AQ premium, similar tests using Kim and Qi (2010) filter rule (Table 6, panel A, 2nd and 3rd last rows) report strongly positive abnormal returns to AQ both across the year, and outside January. However, the mechanical bias in the Kim and Qi (2010) filter is clear, as the average monthly returns of the observations retained in the sample are 0.76 percent more positive than the observations excluded from the sample. For the ex ante price filter, firms retained in the sample have returns 1.21 percent *lower* than firms filtered from the sample (Table 5, panel B). It is thus not surprising that the application of the two rules generates such starkly different results. Panel B reports results for observations with end of month price above $5. The impact of this filter rule on average returns (+2.14 percent) is even greater than that of the Kim and Qi (2010) approach, and this effect seems to be reflected in test results, with test statistics more positive than in panel A. The extent of bias implied by these two filter rules, however, renders any results generated by them highly dubious.

Panels C and D report results for alternative filters for which the impact on measured returns is likely to be more modest. Panel C reports results for a filter that excludes firms with *either* an opening or closing price below $5. Unlike the Kim and Qi (2010) filter, this rule treats firms with positive and negative returns during the month similarly, but may still include a bias caused by the fact that equal-weighed returns are not symmetrical around a given opening price.[15] The observations retained by this filter rule are just 0.15 percent more positive than those excluded (about 1/5th of the bias suggested by the Kim and Qi rule). All tests using this filter rule find evidence of an annual AQ premium (Table 6, panel C, 3rd last row) and premia outside January for two of three tests. Our final filter uses the dollar volume of stock traded in the previous month as a proxy for transaction costs. Dollar volume has been shown to affect the relative spread (Stoll 1989), is a major determinant of liquidity (Brennan and Subrahmanyam 1995), has been used in tests of the impact of illiquidity and

[15] That is, if a stock falls in price from $5 to $4.90 and then rises again to $5.00 the equal weighted returns are -2 percent and +2.04 percent.

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transaction costs on required returns (Brennan et al. 1998; Liu et al. 2016), or as a means of controlling for firm-specific liquidity (Brockman and Chung 2003) and is available on a monthly basis for all of our sample countries. The dollar volume filter may also assist in controlling for transaction cost effects, while having a less severe impact on variation in AQ. In the United States, stock price-based filters exclude between 55-60 percent of firms in the AQ10 portfolio, whereas the lagged dollar volume filter excludes approximately 34 percent of AQ10 firms. Further, this filter rule is independent of price movements occurring in the month in which returns are measured. Panel D reports the results of tests in which observations in the bottom 20 percent of cases ranked by the dollar volume of stock traded in the previous month are excluded from the sample. Our lagged dollar volume filter does not suggest a positive bias to average returns, with observations retained in the sample having 0.72 percent *lower* returns than those filtered. The U.S. dollar volume restricted sample generates positive AQ premia in all three asset pricing tests, regardless of whether January returns are included. The use of current month dollar volume (untabulated) generates similar results.

Other sources of seasonality

As noted above, we find some evidence of modest seasonality in returns around the turn of the financial year in Japan, where the tax and financial years are distinct. There are small reductions in the extent to which AQ is priced when the first month of the financial year is removed from the Japanese sample. We conducted a series of untabulated tests to shed further light on this issue. Untabulated tests of returns in the first month of the financial year generate significant AQ premia in Japan under all three testing methods. For the United States, we examined (untabulated) the first financial month returns for firms whose financial year is distinct from the tax year. While collectively there is no evidence of an AQ premium in this sample, we find significant positive abnormal returns to the AQ hedge portfolio when this sample is further restricted to first financial months occurring in July (the second most common start of the financial year for U.S. firms) and May.

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Why might returns in the first month of the reporting year be abnormally positive in some cases? One possibility is that fund managers' desire to report end-of-year portfolios that are under-weighted in "loser" stock induces selling pressure in the final month of the reporting year (i.e. window-dressing), which reverses swiftly in the first month of the new-year (Lakonishok et al. 1991). While this explanation appears plausible, we find no evidence of negative returns to AQ in the final month of the financial year in Japan (untabulated). Further, an examination of the daily returns in the first month of the financial year in these countries reveals no evidence of a sudden reversal of mispricing (Japan AQ hedge return for April = 2.72 percent, *t*-stat = 2.29; first trading day return = 0.08 percent, *t*-stat = 0.73; first five trading days return = 0.7 percent, *t*-stat = 1.4: UK AQ hedge return for January = 2.54 percent, *t*-stat = 2.37; first trading day return = –0.01 percent, *t*-stat = –0.08; first five trading days return = 0.5 percent, *t*-stat = 1.53). Another possible explanation for abnormally positive returns in the first financial year is that, late in the previous financial year, there is an increase in uncertainty regarding the realization of future earnings, which is gradually resolved early in the new reporting year, and that this uncertainty is priced (Rozeff and Kinney 1976; Kim 2006). While the concentration of January returns in the first few trading days reported by MM suggests that this "information hypothesis" is unlikely to singularly explain the January AQ premium in the United States, the fact that no similar concentration of returns is observed in the first few trading days of the new financial year in Japan and the UK suggests that the hypothesis should not be discounted. An alternate information risk-based explanation may relate to jurisdictional or company specific restrictions on directors' share trading. Many firms impose "close" periods during which directors are banned from trading, and these periods often include the first month of the reporting year. Uninformed investors demand for a firm' stock early in the reporting year may be greater if they believe that trades executed are less likely to involve an informed counterparty. This hypothesis, however, implicitly assumes a differential impact of insider trading restriction on uninformed investors' buying and selling behavior.

6. Conclusion

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Prior U.S. studies report mixed results regarding whether AQ affects the cost of equity capital. In particular, MM show that AQ is priced in the United States only in January and does not attract a significant average annual premium. MM argue that this seasonal pattern likely reflects a mispricing, rather than information risk effect: tax-loss selling of low-AQ firms in December depresses stock prices which subsequently recover in the new year. MM also speculate that the failure to control for seasonality may partly explain the inconsistency in earlier research into the pricing of AQ, as the extent to which the January premium unwinds through the rest of the year varies across study periods. Extending MM's analysis to include other market economies in which tax and reporting year end dates differ (both with respect to each other and relative to the United States), and a longer U.S. sample, we find evidence that an AQ premium exists both across the full year and outside the turn of the tax year in each of the non-U.S. countries studied, and thus that the documented AQ premium is unlikely to be driven by investors' tax-loss selling behavior or other indicators of seasonal mispricing. For the United States, we find a significant AQ premium across the year in two out of three asset pricing tests, but no AQ premium is documented once January (the first tax and financial month) is excluded from the tests. We note, however, that for samples that exclude stocks with an opening price below $5, a proxy for transaction costs, we find no annual AQ premium for U.S. firms. For this reason, we cannot discount the possibility that the premium detected in U.S. full sample tests at least partly reflect mispricing rather than risk. Further analysis using alternative proxies for transactions costs generate results consistent with our full sample tests, but several of these alternative filter rules have the potential to bias recorded returns upwards.

Notwithstanding the fact that we observe AQ premia in non-U.S. countries outside the turn of the tax year, we do observe some evidence of seasonality in returns. For instance, significant AQ premia in April (the first month of the reporting year) in Japan. While it is plausible that a stronger pricing effect around the turn of the financial year reflects mispricing caused by window-dressing, there is no evidence of a concentration of daily returns in the first few trading days of the new financial year. Further, there are plausible risk-related

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explanation for these phenomena, including the coincidence of the start of the new year with director trading "close" periods. Whatever the relative importance of the myriad reasons for seasonality, the fact that the seasonal bias in returns is much stronger in the United States may reflect the conflation of multiple risk and mispricing factors in the December-January period in that country, rather than a simple tax-loss selling explanation.

Our study contributes to the literature that uses earnings quality as a proxy for information risk, presenting evidence that MM's findings that the pricing of AQ likely reflects the mispricing resulting from tax-loss selling is an incomplete explanation of the AQ premia, and supporting earlier findings that AQ is a suitable proxy for this risk. We also contribute to the broader literature on causes of seasonality in returns, utilizing data from countries with distinct tax and financial years to show that seasonal variation in some risk-related premia are more prominent at the turn of the financial year.

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Appendix

Factor-mimicking portfolio formation

Australian, Japanese, and UK data

Most Australian companies have June fiscal year-ends, consequently, we use positive ordinary shareholder's equity on or preceding June 30 in fiscal year *t*, deflated by the total market capitalization on 30 June, to divide stocks into three book-to-market value of equity groups. Firms in the bottom 30 percent, middle 40 percent and top 30 percent of book-to-market ratio are defined as Value, Neutral and Growth book-to-market groups, respectively. In December of each year, all stocks are allocated to one of two size groups, small (S) or big (B) that are split on the median value of total market capitalization on 31 December of each year. We then construct 6 portfolios from the intersections of the two size and three book-to-market groups. For example, the Small Growth portfolio contains the stocks in the small size group that are also in the low book-to-market group. Monthly value-weighted returns on the six portfolios are calculated from January of year t+1 to December of t+1, and the portfolios are reformed in December of year t+1. We calculate returns beginning in January of year t+1 to be sure that book equity for year *t* is known to the market. HML is the difference, each month, between the simple average of the returns on the two value portfolios (Small Value and Big Value) and the average of the returns on the two growth portfolios (Small Growth and Big Growth).That is: *HML* =1/2(Small Value + Big Value) –1/2(Small growth + Big Growth).

The construction of *RMW* (Robust Minus Weak) is similar to that of *HML*, except that it is calculated as the difference between the simple average of returns on the two robust operating

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profitability portfolios minus the average return on the two weak operating profitability portfolios (RMW=1/2(Small Robust + Big Robust)– 1/2(Small Weak + Big Weak)). Operating profit is calculated as revenue minus cost of goods sold, interest expense, and selling, general and administrative expense deflated by book equity at the end of fiscal year *t*–1. Firms in the bottom 30 percent, middle 40 percent and top 30 percent of operating profit are defined as Weak, Neutral and Robust profitability groups, respectively

The construction of *CMA* (Conservative Minus Aggressive) is similar to that of *HML*, except that it is calculated as the difference between the simple average of the monthly returns on the two conservative investment portfolios minus the average return on the two aggressive investment portfolio (*CMA*=1/2 (Small Conservative + Big Conservative) – 1/2(Small Aggressive + Big Aggressive)). Investment is calculated as the change in total assets from the fiscal year ending in year *t*–2 to the fiscal year ending in *t*–1, divided by total asset at the end of year *t*–2. Firms in the bottom 30 percent, middle 40 percent and top 30 percent of investment are defined as Conservative, Neutral and Aggressive Investment groups, respectively

SMB is calculated as the difference, each month, between average return on the nine small stock portfolios and the average return on the nine big stock portfolios.

$SMB = 1/3(SMB_{(B/M)} + SMB_{(OP)} + SMB_{(INV)})$

Where:

$SMB_{(B/M)}$ = 1/3 (Small Value + Small Neutral + Small Growth)

– 1/3 (Big Value + Big Neutral + Big Growth).

$SMB_{(OP)}$ = 1/3 (Small Robust + Small Neutral + Small Weak)

– 1/3 (Big Robust + Big Neutral + Big Weak).

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$SMB_{(INV)}$ = 1/3 (Small Conservative + Small Neutral + Small Aggressive)

– 1/3 (Big Conservative + Big Neutral + Big Aggressive).

The Japanese and UK size, book-to-market, profitability and investment factor-mimicking portfolio are constructed using similar methods as those used in the Australian data, with an adjustment to portfolio formation dates to reflect the standard Japanese and UK financial year-ends (March 31 and December 31 respectively).

U.S. data

The U.S. size, book-to-market, profitability and investment factor-mimicking portfolio are obtained from Kenneth French's website (http://mba.tuck.darthmouth.edu/pages/faculty/Ken.french/data_library.html). See the website and Fama and French (2015) for details of his calculations.

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TABLE 1
Descriptive statistics

	Japan					United Kingdom				
	Mean	Std. Dev.	Q1	Median	Q3	Mean	Std. Dev.	Q1	Median	Q3
AQ	0.024	0.019	0.012	0.019	0.030	0.054	0.043	0.024	0.041	0.069
BETA	0.836	0.539	0.453	0.770	1.147	0.884	0.714	0.423	0.824	1.262
BM	1.437	0.929	0.794	1.233	1.836	0.798	0.766	0.331	0.592	0.990
SIZE	56483	139389	5016	12913	38959	862	3196	16	79	376
PRICE	2598	26808	263	511	1040	4.02	41.93	0.45	1.27	3.15
RM-RF	0.007	0.050	-0.022	0.008	0.038	0.002	0.044	-0.018	0.009	0.031
SMB	0.007	0.037	-0.013	0.003	0.024	-0.002	0.039	-0.025	-0.003	0.018
HML	0.007	0.022	-0.005	0.007	0.021	0.006	0.030	-0.010	0.005	0.021
RMW	0.009	0.037	-0.004	0.006	0.018	0.004	0.032	-0.015	0.007	0.021
CMA	0.003	0.038	-0.011	0.001	0.010	0.005	0.027	-0.011	0.005	0.019
CV	0.048	0.031	0.027	0.040	0.059	0.082	0.073	0.036	0.060	0.103
OP	0.090	0.058	0.051	0.082	0.123	0.104	0.138	0.065	0.116	0.170
	Australia					**United States**				
	Mean	Std. Dev.	Q1	Median	Q3	Mean	Std. Dev.	Q1	Median	Q3
AQ	0.091	0.066	0.045	0.072	0.114	0.062	0.083	0.021	0.038	0.071
BETA	0.789	0.846	0.287	0.702	1.219	1.128	0.668	0.666	1.063	1.493
BM	1.078	1.157	0.414	0.760	1.308	1.271	3.911	0.298	0.564	1.024
SIZE	935	3164	16	71	416	1931	6789	35	169	921
PRICE	4.10	20.87	0.24	0.90	3.29	21.46	26.20	6.00	14.96	29
RM-RF	0.001	0.037	-0.021	0.006	0.027	0.005	0.046	-0.020	0.009	0.035
SMB	0.008	0.055	-0.030	0.004	0.041	0.002	0.031	-0.015	0.001	0.020
HML	0.005	0.027	-0.012	0.008	0.021	0.004	0.030	-0.013	0.003	0.017
RMW	0.004	0.031	-0.013	0.003	0.022	0.003	0.022	-0.009	0.002	0.014
CMA	0.006	0.031	-0.012	0.003	0.025	0.003	0.020	-0.009	0.002	0.015
CV	0.105	0.093	0.041	0.071	0.139	0.089	0.075	0.040	0.067	0.110
OP	0.026	0.235	-0.001	0.071	0.121	0.143	0.121	0.089	0.144	0.208

Notes: AQ is firms' accruals quality, estimated using equation (1); *BETA* is a firm's rolling market beta from the CAPM model, estimated over the 36 months prior to the current month for Australia, Japan and the UK and 60 months for the US; *SIZE* is firms' market value of common equity (in millions) calculated as price times shares outstanding; *BM* is firms' book-to-market ratio; *RM-RF* is Datastream value-weighted market return less risk-free rate for Japan, the UK, and Australia. For the United States, *RM-RF* is obtained from Kenneth French's website; *SMB* is the return on the size factor-mimicking portfolio; *HML* is the return on the book-to-market factor-mimicking portfolio; *RMW* is the return on the operating profitability factor-mimicking portfolio; *CMA* is the return on the investment factor-mimicking portfolio; *CV* is a firm's cash flow volatility, estimated as the standard deviation of the firm's rolling seven-year cash flow from operations deflated by average assets, e.g., all the cash flow from operations used in estimating that firm' AQ; *OP* is firms' operating profit, calculated as revenue minus cost of goods sold and selling, general and administrative expense (excluding expenditures on research and development) deflated by the book value of total assets. *AQ, BETA, BM, SIZE, OP,* and *CV* are

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winsorized at the first and ninety-ninth percentiles. We obtain 17,070 firm-year observations (2,396 firms) of AQ with fiscal years ending between 2002 and 2014 for Japan (stock returns lag our AQ observations by up to one year), 11,170 firm-year observations (1, 237 firms) of AQ with fiscal years ending between 1993 and 2013 for the UK, 4,888 firm-year observations (751 firms) with fiscal years ending between 1998 and 2013 for Australia, and 147,829 firm-year observations (12,726 firms) with fiscal years ending between 1970 to 2014 for the United States.

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TABLE 2

AQ hedge portfolio abnormal returns

$$R_{p,t} = a_{0,p} + B_p(R_{m,t} - R_{f,t}) + s_pSMB_t + h_pHML_t + r_pRMW_t + c_pCMA_t + \varepsilon_{p,t}$$

Panel A: Using the full sample

Japan

	All months			Excluding first tax month			Excluding first financial month		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.010	***	2.80	0.010	***	3.01	0.009	**	2.70
RM-RF	0.421	***	7.37	0.431	***	7.35	0.391	***	7.27
SMB	0.670	***	10.67	0.756	***	10.6	0.686	***	9.41
HML	-0.517	***	-5.22	-0.497	***	-4.73	-0.443	***	-4.40
RMW	-0.485	***	-4.06	-0.483	***	-3.71	-0.602	***	-5.50
CMA	0.017		0.15	-0.052		-0.42	0.111		1.03
Adj. R^2	0.457			0.464			0.444		
N(Month)	150			137			137		

United Kingdom

	All months			Excluding first tax month			Excluding first financial month		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.015	***	3.74	0.015	***	3.76	0.018	***	3.54
RM-RF	0.145	***	2.26	0.117	*	1.84	0.132	**	2.13
SMB	0.627	***	5.53	0.630	***	5.49	0.648	***	4.84
HML	-0.567	***	-4.08	-0.611	***	-4.45	-0.623	***	-4.01
RMW	-0.397	***	-3.25	-0.464	***	-3.68	-0.489	***	-3.46
CMA	-0.108		-0.81	-0.034		-0.30	-0.090		-0.66
Adj. R^2	0.504			0.537			0.459		
N(Month)	209			192			191		

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TABLE 2, Panel A (continued)

Australia

	All months			Excluding first tax (financial) month		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.009	***	3.23	0.010	***	3.20
RM-RF	-0.013		-0.19	-0.047		-0.64
SMB	0.362	***	5.16	0.389	***	4.94
HML	-0.077		-0.77	-0.083		-0.82
RMW	0.008		0.08	0.002		0.01
CMA	-0.261	***	-3.33	-0.212	***	-2.67
Adj. R^2	0.331			0.300		
N(Month)	192			176		

United States

	All months			Excluding first tax (financial) month			January		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.005	**	2.52	-0.001		-0.45	0.053	***	6.78
RM-RF	0.124	**	2.36	0.151	***	3.62	0.030		0.21
SMB	0.917	***	11.13	0.822	***	11.7	1.264	***	5.02
HML	-0.439	***	-3.98	-0.539	***	-5.71	-0.883	***	-3.6
RMW	-0.715	***	-6.18	-0.540	***	-6.00	-1.103	**	-2.4
CMA	-0.206		-0.89	-0.009		-0.07	-0.514		-0.97
Adj. R^2	0.586			0.625			0.568		
N(Month)	528			484			44		

Panel B: Using the price-restricted sample

Japan

	All months			Excluding first tax month			Excluding first financial month		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.004	**	2.01	0.005	**	2.43	0.004	*	1.89
RM-RF	0.348	***	9.15	0.355	***	9.48	0.348	***	9.06
SMB	0.510	***	8.91	0.582	***	9.35	0.489	***	7.64
HML	-0.401	***	-6.02	-0.399	***	-5.61	-0.387	***	-5.02
RMW	-0.361	***	-5.86	-0.364	***	-5.67	-0.384	***	-5.60
CMA	-0.001		-0.02	-0.058		-0.74	0.032		0.39
Adj. R^2	0.587			0.613			0.571		
N(Month)	150			137			137		

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TABLE 2, Panel B (continued)

United Kingdom

	All months			Excluding first tax month			Excluding first financial month		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.007	*	1.83	0.008	*	1.90	0.009	*	1.81
RM-RF	0.216	***	3.49	0.216	***	3.13	0.219	***	3.66
SMB	0.543	***	5.46	0.556	***	5.11	0.535	***	4.50
HML	-0.446	***	-3.39	-0.422	***	-3.03	-0.485	***	-3.28
RMW	-0.413	***	-3.54	-0.448	***	-3.58	-0.488	***	-3.69
CMA	-0.245	**	-2.29	-0.236	**	-2.25	-0.212	**	-2.10
Adj. R^2	0.540			0.554			0.413		
N(Month)	208			191			190		

Australia

	All months			Excluding first tax (financial) month		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.005	*	1.68	0.006	**	1.99
RM-RF	-0.069		-1.06	-0.077		-1.18
SMB	0.231	***	3.35	0.288	***	3.65
HML	-0.195	*	-1.84	-0.135		-1.22
RMW	-0.086		-0.86	-0.047		-0.45
CMA	-0.259	***	-3.64	-0.225	***	-3.30
Adj. R^2	0.259			0.258		
N(Month)	192			176		

United States

	All months			Excluding first tax (financial) month			January		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.000		-0.08	-0.002		-1.58	0.015	***	3.35
RM-RF	0.156	***	4.54	0.175	***	5.19	0.115		1.33
SMB	0.713	***	12.76	0.691	***	13.26	0.896	***	5.04
HML	-0.624	***	-9.00	-0.639	***	-9.59	-0.714	***	-3.29
RMW	-0.459	***	-6.31	-0.415	***	-5.75	-0.311		-1.55
CMA	-0.168		-1.38	-0.127		-1.27	-0.190		-0.71
Adj. R^2	0.712			0.728			0.660		
N(Month)	528			484			44		

Notes: *, **, and *** indicate statistical significance at the 10 percent, 5 percent, and 1 percent levels (two-tailed), respectively. *t*-stats are calculated using Newey-West standard errors with two lags. Firms with available

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AQ measures are ranked into one of ten decile portfolios at the end of each month according to their most recent AQ scores. The hedge portfolio is long in the poorest AQ decile (AQ10) and short in the best AQ decile (AQ1) for Japan, the UK and the United States. The AQ hedge portfolio returns for Australia is calculated as equal-weighted portfolio returns of [(AQ5+AQ4)/2-(AQ1+AQ2)/2], due to the relative smaller sample size. To calculate the hedge portfolio abnormal returns, the average monthly equal-weighted returns on the hedge portfolios are regressed on the monthly Fama and French (2015) five factors *(RM-RF, SMB, HML, RMW, CMA)*. The intercept from this time-series regression is the average monthly hedge portfolio abnormal return. The sample period covers January 2003 to June 2015 for Japan, October 1996 to March 2014 for the UK, October 1998 to September 2014 for Australia, and January 1971 to December 2014 for the United States. Panel A shows the full sample results; Panel B presents results after deleting firms with opening price below the threshold ($5 for the U.S., ¥180 for Japan, STG0.47 for the UK, and AUD0.19 for Australia)

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TABLE 3

Fama-MacBeth cross-sectional regression of returns against risk factors

$$R_{it} = \alpha_0 + \alpha_1 AQ_{it-1} + \alpha_2 BETA_{it-1} + \alpha_3 LN(SIZE)_{it-1} + \alpha_4 LN(BM)_{it-1} + \alpha_5 OP_{it-1} + \alpha_6 CV_{it-1} + \varepsilon_{it}$$

Panel A: Using the full sample

Japan

	All months			**Excluding first tax month**			**Excluding first financial month**		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.009		1.09	0.002		0.21	0.005		0.61
AQ	0.263	***	3.87	0.242	***	3.53	0.259	***	3.52
BETA	0.002		0.81	0.002		0.63	0.001		0.20
ln(SIZE)	-0.002	***	-2.66	-0.001		-1.45	-0.001	**	-2.01
ln(BM)	0.008	***	5.16	0.008	***	5.09	0.009	***	5.98
OP	0.091	***	5.18	0.090	***	5.08	0.092	***	4.70
CV	0.058		1.35	0.073		1.56	0.044		0.97
Adj. R^2	0.059			0.059			0.059		
N	2372			2372			2372		

United Kingdom

	All months			**Excluding first tax month**			**Excluding first financial month**		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.007	***	1.33	0.004	**	0.82	0.003		0.67
AQ	0.109	*	1.75	0.118	*	1.74	0.115	*	1.67
BETA	0.001		0.32	-0.001		-0.51	0.001		0.45
ln(SIZE)	0.001		1.04	0.001		1.17	0.001	*	1.87
ln(BM)	0.006	***	3.53	0.006	***	3.05	0.006	***	3.38
OP	0.003		0.18	0.002		0.11	0.004		0.20
CV	-0.010		-0.42	-0.008		-0.3	-0.026		-0.95
Adj. R^2	0.049			0.049			0.047		
N	916			916			916		

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TABLE 3, Panel A (continued)

Australia

	All months			**Excluding first tax (financial) month**		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.058	**	2.53	0.040	*	1.71
AQ	0.056	***	2.72	0.064	***	2.92
BETA	-0.003		-1.12	-0.004	*	-1.64
ln(SIZE)	-0.003	***	-3.38	-0.002	**	-2.31
ln(BM)	0.003	*	1.61	0.002		1.04
OP	0.004		0.45	0.010		0.96
CV	0.002		0.11	-0.007		-0.38
Adj. R^2	0.042			0.039		
N	713			712		

United States

	All months			**Excluding first tax (financial) month**			**January**		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.034	***	4.97	0.018	***	2.68	0.214	***	7.41
AQ	0.031	**	2.29	0.006		0.47	0.306	***	3.75
BETA	0.001		0.53	-0.001		-0.64	0.018	***	3.16
ln(SIZE)	-0.001	***	-3.81	-0.001		-1.39	-0.010	***	-6.76
ln(BM)	0.004	***	5.63	0.003	***	4.52	0.015	***	3.77
OP	0.022	***	6.68	0.028	***	8.86	-0.038	***	-2.93
CV	-0.003		-0.50	-0.015	**	-2.24	0.125	***	5.09
Adj. R^2	0.042			0.037			0.093		
N	8052			8050			7840		

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TABLE 3 (continued)

Panel B: Using the priced-restricted sample

Japan

	All months			Excluding first tax month			Excluding first financial month		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	-0.008		-1.07	-0.015	*	-1.78	-0.009		-1.11
AQ	0.179	***	2.86	0.145	**	2.09	0.164	***	2.65
BETA	0.001		0.20	0.001		0.18	-0.001		-0.40
ln(SIZE)	0.000		-0.21	0.001		0.74	0.000		0.18
ln(BM)	0.011	***	6.32	0.011	***	6.35	0.012	***	6.66
OP	0.125	***	6.96	0.125	***	6.91	0.120	***	6.52
CV	0.052	**	2.13	0.062	**	2.39	0.042		1.63
Adj. R^2	0.057			0.057			0.056		
N	2285			2285			2285		

United Kingdom

	All months			Excluding first tax month			Excluding first financial month		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.005		0.98	0.003		0.56	0.002		0.37
AQ	0.138	**	2.09	0.149	**	2.10	0.133	*	1.83
BETA	-0.001		-0.29	-0.002		-0.96	0.000		0.02
ln(SIZE)	0.001		1.43	0.001		1.51	0.002	**	2.23
ln(BM)	0.006	***	2.81	0.005	**	2.53	0.006	***	2.81
OP	0.012		0.67	0.011		0.57	0.013		0.64
CV	-0.031		-1.20	-0.029		-1.08	-0.045		-1.60
Adj. R^2	0.052			0.051			0.044		
N	745			745			744		

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TABLE 3, Panel B (continued)

Australia

	All months			Excluding first tax (financial) month		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.022		1.28	0.007		0.35
AQ	0.041	***	2.37	0.050	***	2.79
BETA	0.001		0.28	-0.001		-0.27
ln(SIZE)	-0.001		-0.71	0.000		-0.04
ln(BM)	0.001		0.81	0.000		0.12
OP	0.012		1.06	0.010		0.83
CV	-0.009		-0.57	-0.015		-0.95
Adj. R^2	0.045			0.044		
N	582			581		

United States

	All months			Excluding first tax (financial) month			January		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.020	***	3.23	0.012	**	2.00	0.105	***	3.86
AQ	0.003		0.29	-0.010		-0.88	0.140	***	2.65
BETA	0.000		0.05	-0.001		-0.83	0.014	***	2.86
ln(SIZE)	-0.001	*	-1.90	0.000		-0.51	-0.005	***	-3.74
ln(BM)	0.002	***	2.88	0.001	*	1.74	0.012	***	2.93
OP	0.025	***	6.98	0.025	***	7.22	0.024		1.59
CV	-0.009		-1.36	-0.014	**	-2.06	0.045	***	3.38
Adj. R^2	0.048			0.045			0.074		
N	6973			6965			6492		

Notes: *, **, and *** indicate statistical significance at the 10 percent, 5 percent, and 1 percent levels (two-tailed), respectively. This table reports the mean values of the annual cross-sectional regression coefficient estimates obtained from regressing firms' monthly returns on *AQ*, firms' market beta (*BETA*), *ln(SIZE)*, *ln(BM)*, *OP*, and *CV* using the full sample (Panel A), and the price-restricted sample (Panel B), which excludes firms with opening price below the threshold ($5 for the United States, ¥180 for Japan, STG0.47 for the U.K, and AUD0.19 for Australia). Book value, *AQ*, *OP* and *CV* are matched into returns three months after the fiscal year-end. All explanatory variables are winsorized at the first and ninety-ninth percentiles. *t*-stats are calculated using Newey-West standard errors with two lags. The sample period covers January 2003 to June 2015 for Japan, July 1993 to March 2014 for the UK, October 1998 to September 2014 for Australia, and January 1971 to December 2014 for the United States.

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TABLE 4

Two stage cross-sectional regression

Panel A: First-stage regression: Time-series regressions of contemporaneous excess returns on factor returns using the full sample

$$R_{p,t} - R_{F,t} = \alpha_0 + \beta_{p,RM\text{-}RF}\,(R_{M,t} - R_{F,t}) + \beta_{p,SMB}SMB_t + \beta_{p,HML}HML_t + \beta_{p,AQ\ factor}\,AQfactor_t + \beta_{p,RMW}RMW_t + \beta_{p,CMA}CMA_t + \square_{i,t}$$

Japan

	All months			**Excluding first tax month**			**Excluding first financial month**		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.003	***	4.49	0.003	***	4.21	0.004	***	5.1
RM-RF	0.856	***	58.58	0.837	***	51.57	0.865	***	59.22
SMB	0.563	***	17.07	0.535	***	14.69	0.594	***	15.54
HML	0.239	***	6.12	0.268	***	6.64	0.235	***	5.64
AQfactor	0.237	***	3.20	0.279	***	3.53	0.197	**	2.46
RMW	-0.336	***	-12.91	-0.337	***	-12.03	-0.389	***	-12.67
CMA	-0.018		-0.65	-0.016		-0.51	0.010		0.35
R^2	0.595			0.594			0.604		

United Kingdom

	All months			**Excluding first tax month**			**Excluding first financial month**		
	Coef.		***t*-stat**	**Coef.**		***t*-stat**	**Coef.**		***t*-stat**
Intercept	0.004	***	7.12	0.004	***	6.41	0.004	***	5.99
RM-RF	0.862	***	46.52	0.856	***	48.01	0.883	***	46.85
SMB	0.698	***	24.97	0.717	***	23.05	0.684	***	23.44
HML	0.134	***	4.54	0.096	***	3.14	0.152	***	5.21
AQfactor	0.101	***	3.14	0.084	***	2.68	0.103	***	2.84
RMW	-0.078	***	-4.64	-0.105	***	-5.92	-0.057	***	-3.14
CMA	-0.008		-0.33	0.036		1.46	-0.008		-0.35
R^2	0.483			0.473			0.490		

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TABLE 4, Panel A (continued)

Australia

	All months		**Excluding first tax(financial) month**	
	Coef.	***t*-stat**	**Coef.**	***t*-stat**
Intercept	0.007 ***	4.39	0.006 ***	3.73
RM-RF	0.531 ***	26.01	0.555 ***	25.02
SMB	0.644 ***	15.44	0.577 ***	13.25
HML	0.132 ***	3.26	0.105 **	2.66
AQfactor	0.072 ***	2.37	0.088 ***	2.76
RMW	-0.101 ***	-3.99	-0.135 ***	-5.03
CMA	-0.116 ***	-3.00	-0.158 ***	-3.96
R^2	0.377		0.377	

United States

	All months		**Excluding first tax(financial) month**		**January**	
	Coef.	***t*-stat**	**Coef.**	***t*-stat**	**Coef.**	***t*-stat**
Intercept	0.001 **	2.06	0.000	0.39	0.002	0.92
RM-RF	0.950 ***	65.04	0.966 ***	64.51	0.911 ***	59.44
SMB	0.519 ***	18.37	0.555 ***	19.88	0.571 ***	14.09
HML	0.303 ***	9.28	0.287 ***	8.91	0.148 ***	3.81
AQfactor	0.641 ***	8.92	0.537 ***	7.98	0.735 ***	9.54
RMW	0.080 ***	4.51	0.066 ***	3.63	0.039	1.11
CMA	-0.045 **	-2.1	-0.033 *	-1.66	0.017	0.32
R^2	0.814		0.801		0.848	

Notes: *** and ** indicate statistical significance at the 1 percent and 5 percent levels, respectively (two-tailed). This table reports the mean coefficients and mean R^2 of 100 time-series regressions of size and book-to-market portfolios' current monthly excess return (stock return less the risk-free rate) on the Fama-French five factors and the AQ factor for all months (non-January months). As in Fama and French (1992), firms are first ranked into 10 size portfolios according to their market value at the end of the sixth month after the reporting year. Since for most firms, the end month of financial reporting year is December in the United States and the UK, June in Australia, and March in Japan, firms' market value at the end of June of year *t* is used to rank the size portfolios for the United States and the UK. For Australia, firms' market value at the end of December of year *t* is used for the ranking; and for Japan, firms' market value at the end of September of year *t* is used to construct the size portfolio. Within each size decile portfolio, firms are further ranked into 10 portfolios based on their market-to-book ratios at the end of year *t–1* (December) for the United States and the UK, and at the end of financial year *t* for Australia (June) and Japan (March). Portfolios are re-balanced every year, and portfolio betas are computed using the whole-period returns. AQ*factor* is the monthly return to the accruals quality factor–mimicking portfolio, calculated as the difference between the monthly equal-weighted portfolio returns of firms in the poorest two AQ quintiles and firms in the best two AQ quintiles. $R_{p,t}$ is the monthly

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equal-weighted portfolio returns; *RM-RF* is the monthly value-weighted market return les risk-free rate; *SMB* is the monthly return on the size factor–mimicking portfolio; *HML* is the monthly return on the book-to-market factor-mimicking portfolio; *RMW* is the monthly return on the operating profitability factor-mimicking portfolio; and *CMA* is the monthly return on the investment factor-mimicking portfolio. The sample period covers January 2003 to June 2015 for Japan, July 1993 to March 2014 for the UK, October 1998 to September 2014 for Australia, and January 1971 to December 2014 for the United States

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TABLE 4 (continued)

Panel B: Second stage regression: Cross-sectional regressions of excess returns on factor betas

$$R_{pi,t} - R_{F,t} = \lambda_0 + \lambda_1\hat{\beta}_{pi,RM-RF} + \lambda_2\hat{\beta}_{pi,SMB} + \lambda_3\hat{\beta}_{pi,HML} + \lambda_4\hat{\beta}_{pi,AQfactor} + \lambda_5\hat{\beta}_{pi,RMW} + \lambda_6\hat{\beta}_{pi,CMA} + \lambda_7 Low\_priced + \mu_{i,t}$$

Japan

	All months				Excluding first tax month				Excluding first financial month			
	Model 1		Model 2		Model 1		Model 2		Model 1		Model 2	
	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat
Intercept	-0.001	-0.08	0.000	-0.03	-0.003	-0.32	-0.003	-0.3	-0.005	-0.39	-0.005	-0.45
$\hat{\beta}_{i,RM-RF}$	0.012	0.99	0.012	0.97	0.013	0.97	0.013	0.97	0.018	1.22	0.019	1.32
$\hat{\beta}_{i,SMB}$	-0.002	-0.47	-0.004	-0.98	-0.005	-1.06	-0.007	-1.37	-0.003	-0.72	-0.005	-0.99
$\hat{\beta}_{i,HML}$	0.008 ***	2.73	0.008 ***	2.89	0.006 **	2.02	0.006 **	2.11	0.008 ***	2.64	0.008 ***	2.69
$\hat{\beta}_{i,AQfactor}$	0.008 ***	2.71	0.006 **	2.35	0.007 **	2.29	0.006 **	2.05	0.008 ***	2.75	0.007 ***	2.64
$\hat{\beta}_{i,RMW}$	-0.003	-0.58	-0.003	-0.61	-0.003	-0.63	-0.003	-0.63	-0.004	-0.68	-0.003	-0.57
$\hat{\beta}_{i,CMA}$	-0.008	-1.64	-0.008	-1.7	-0.007	-1.44	-0.008	-1.51	-0.007	-1.32	-0.007	-1.35
Low_priced			0.006 *	1.85			0.004	1.13			0.005	1.52
Adj. R^2	0.050		0.058		0.050		0.058		0.052		0.060	

United Kingdom

	All months				Excluding first tax month				Excluding first financial month			
	Model 1		Model 2		Model 1		Model 2		Model1		Model 2	
	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat
Intercept	0.008 **	2.27	0.005	1.6	0.006 *	1.78	0.005	1.41	0.003	0.99	0.002	0.53
$\hat{\beta}_{i,RM-RF}$	-0.003	-0.68	0.000	-0.05	-0.003	-0.70	-0.001	-0.31	0.001	0.25	0.002	0.56
$\hat{\beta}_{i,SMB}$	-0.003	-0.87	-0.004	-1.53	-0.003	-1.05	-0.004	-1.41	-0.004	-1.5	-0.005 *	-1.85
$\hat{\beta}_{i,HML}$	0.009 ***	2.85	0.009 ***	2.91	0.007 ***	2.65	0.008 **	2.64	0.011 **	3.52	0.011 ***	3.5
$\hat{\beta}_{i,AQfactor}$	0.009 **	2.44	0.008 **	2.34	0.010 ***	2.80	0.010 ***	2.73	0.006 *	1.91	0.006 *	1.88
$\hat{\beta}_{i,RMW}$	0.002	0.63	0.002	0.59	0.000	0.08	0.000	0.08	0.000	0.02	0.000	0.07
$\hat{\beta}_{i,CMA}$	0.007 **	2.27	0.007 **	2.22	0.008 **	2.30	0.008 **	2.27	0.007 **	2.17	0.007 **	2.18
Low_priced			0.004 ***	2.74			0.003 ***	2.00			0.003 *	1.65
Adj. R^2	0.016		0.018		0.016		0.018		0.016		0.017	

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TABLE 4, Panel B (continued)

	All months				**Excluding First Tax (Financial) Month**			
	Model 1		Model 2		Model 1		Model 2	
	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat
Intercept	0.003	0.8	0.003	0.99	-0.001	-0.46	-0.001	-0.3
$\hat{\beta}_{i,RM-RF}$	-0.009 *	-1.76	-0.009 *	-1.78	-0.003	-0.49	-0.004	-0.69
$\hat{\beta}_{i,SMB}$	0.014 ***	3.35	0.004	1.02	0.009 **	2.03	0.001	0.38
$\hat{\beta}_{i,HML}$	0.015 ***	5.86	0.014 ***	5.54	0.015 ***	5.63	0.015 ***	5.44
$\hat{\beta}_{i,AQfactor}$	0.011 ***	2.99	0.007 **	1.95	0.013 ***	3.77	0.010 ***	2.99
$\hat{\beta}_{i,RMW}$	-0.026 ***	-5.76	-0.023 ***	-5.10	-0.020 ***	-4.54	-0.018 ***	-4.11
$\hat{\beta}_{i,CMA}$	0.012 ***	4.49	0.010 ***	3.60	0.012 ***	4.19	0.010 ***	3.68
Low_priced			0.022 ***	8.24			0.016 ***	6.72
Adj. R^2	0.017		0.020		0.016		0.018	

United States

	All months				**Excluding first tax (financial) month**				**January**			
	Model 1		Model 2		Model 1		Model 2		Model 1		Model 2	
	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat	Coef.	*t*-stat
Intercept	0.031 ***	9.58	0.028 ***	8.58	0.029 ***	9.52	0.030 ***	9.56	0.050 ***	6.48	0.030 ***	3.52
$\hat{\beta}_{i,RM-RF}$	-0.022 ***	-6.02	-0.020 ***	-5.17	-0.021 ***	-6.05	-0.022 ***	-5.95	-0.041 ***	-3.63	-0.021 *	-1.69
$\hat{\beta}_{i,SMB}$	-0.001	-0.88	-0.002	-1.13	-0.003 *	-1.96	-0.003 **	-2.23	0.030 ***	4.98	0.028 ***	4.74
$\hat{\beta}_{i,HML}$	0.005 ***	3.45	0.005 ***	3.16	0.004 ***	2.76	0.004 ***	2.69	0.012 *	1.83	0.014 **	2.29
$\hat{\beta}_{i,AQfactor}$	-0.001	-0.72	-0.002 *	-1.81	-0.005 ***	-3.55	-0.004 ***	-3.48	0.037 ***	5.49	0.030 ***	5.49
$\hat{\beta}_{i,RMW}$	-0.003	-1.49	-0.003	-1.58	-0.002	-1.17	-0.002	-1.11	-0.022 ***	-5.12	-0.013 ***	-2.91
$\hat{\beta}_{i,CMA}$	0.003 **	2.52	0.003 **	2.5	0.002 *	1.72	0.002 *	1.67	0.03 ***	6.07	0.021 ***	3.93
Low_priced			0.003 *	1.76			-0.003 *	-1.79			0.124 ***	6.65
Adj. R^2	0.022		0.027		0.019		0.023		0.05		0.07	

Notes: *, **, and *** indicate statistical significance at the 10 percent, 5 percent, and 1 percent levels (two-tailed), respectively.

This table presents the average month-by-month cross-sectional regression coefficients of firm-specific excess returns on full-period factor betas using Fama-Macbeth's (1973) two-pass methodology. The full-period betas are estimated from the first-stage multivariate time-series regression model of 100 portfolio excess return on Fama-

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French five factors and the AQ factor from January 2003 to June 2015 for Japan, July 1993 to March 2014 for the UK, October 1998 to September 2014 for Australia, and January 1971 to December 2014 for the United States. *Low_priced* is a dichotomous variable indicating the presence of a firm with opening price below the threshold ($5 for the United States, ¥180 for Japan, STG0.47 for the UK, and AUD0.19 for Australia). Adj. R^2 is the average adjusted R^2 of the month-by-month cross-sectional regressions.

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TABLE 5

Summary results of the documented AQ premium from the three asset pricing tests

Panel A: Summary of full sample results

		AQ hedge portfolio abnormal returns			**Fama-Macbeth regression**			**Two stage regression (2nd stage)**		
		Intercept		*t*-stat	Coef. of AQ		*t*-stat	Coef. of AQ factor loading		*t*-stat
Japan	All months	0.010	***	2.80	0.263	***	3.87	0.008	**	2.71
	Non-first tax months	0.010	***	3.01	0.242	***	3.53	0.007	**	2.29
	Non-first financial months	0.009	**	2.70	0.259	***	3.52	0.008	**	2.75
UK	All months	0.015	***	3.74	0.109	*	1.75	0.009	**	2.44
	Non-first tax months	0.015	***	3.76	0.118	*	1.74	0.010	***	2.80
	Non- first financial months	0.018	***	3.54	0.115	*	1.67	0.006	*	1.91
Australia	All months	0.009	***	3.23	0.056	***	2.72	0.011	***	2.99
	Non-first tax (financial) months	0.010	***	3.20	0.064	***	2.92	0.013	***	3.77
U.S.	All months	0.005	**	2.52	0.031	**	2.29	-0.001		-0.72
	Non-first tax (financial) months	-0.001		-0.45	0.006		0.47	-0.005		-3.55
	January	0.053	***	6.78	0.306	***	3.75	0.037	***	5.49

Panel B: Summary of Restricted Sample Results (US observations excluded from the sample have mean monthly returns 1.25 percent greater than those retained)

		AQ hedge portfolio abnormal returns			**Fama-Macbeth regression**			**Two-stage regression(2nd stage)**		
		Intercept		*t*-stat	Coef. of AQ		*t*-stat	Coef. of AQ factor loading		*t*-stat
Japan	All months	0.004	**	2.01	0.179	***	2.86	0.006	**	2.35
	Non-first tax months	0.005	**	2.43	0.145	**	2.09	0.006	**	2.05
	Non-first financial months	0.004	*	1.89	0.164	***	2.65	0.007	**	2.64
UK	All months	0.007	*	1.83	0.138	**	2.09	0.008	**	2.34
	Non-first tax months	0.008	*	1.90	0.149	**	2.10	0.010	***	2.73
	Non- first financial months	0.009	*	1.81	0.133	*	1.83	0.006	*	1.88
Australia	All months	0.005	*	1.68	0.041	**	2.37	0.007	**	1.95
	Non-first tax (financial) months	0.006	**	1.99	0.050	***	2.79	0.010	***	2.99
U.S.	All months	0.000		-0.08	0.003		0.29	-0.002		-1.81
	Non-first tax (financial) months	-0.002		-1.58	-0.010		-0.88	-0.004		-3.48
	January	0.015	***	3.35	0.140	***	2.65	0.030	***	5.49

Notes: *, **, and *** indicate statistical significance in predicted direction at the 10 percent, 5 percent, and 1 percent levels (two-tailed), respectively. This table presents the summary results of the documented AQ premium from the three asset pricing tests (Tables 2-4).

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TABLE 6

Summary results of the documented AQ premium from the three asset pricing tests for alternate filters

Panel A: Kim and Qi filter (excluding firms with opening and closing prices below the threshold – U.S. observations excluded from sample have mean monthly returns 0.761 percent lower than those retained)

		AQ hedge portfolio abnormal returns			**Fama-Macbeth Regression**			**Two stage regression (2nd stage)**		
		Coef. of intercept		*t*-stat	Coef. of AQ		*t*-stat	Coef. Of AQ factor loading		*t*-stat
Japan	All months	0.009	***	4.25	0.269	***	3.46	0.009	***	3.25
	Non-first tax months	0.009	***	4.47	0.236	***	2.76	0.009	***	3.01
	Non-first financial months	0.008	***	3.95	0.245	***	2.98	0.010	***	3.36
UK	All months	0.017	***	3.42	0.147	**	2.21	0.010	***	2.78
	Non-first tax months	0.017	***	3.39	0.158	**	2.21	0.011	***	3.11
	Non- first financial months	0.019	***	2.98	0.140	*	1.93	0.007	**	2.15
Australia	All months	0.011	***	3.58	0.048	**	2.42	0.014	***	3.97
	Non-first tax (financial) months	0.011	***	3.37	0.054	***	2.71	0.018	***	4.97
U.S.	All months	0.010	***	7.03	0.044	***	3.63	0.007	***	5.01
	Non-first tax (financial) months	0.008	***	5.84	0.027	**	2.20	0.006	***	4.21
	January	0.028	***	5.08	0.233	***	3.58	0.030	***	5.78

Panel B: Ex post filter (excluding firms with closing price below the threshold – U.S. observations excluded from sample have mean monthly returns 2.14 percent lower than those retained)

		AQ hedge portfolio abnormal returns			**Fama-Macbeth regression**			**Two stage regression (2nd stage)**		
		Coef. of intercept		*t*-stat	Coef. of AQ		*t*-stat	Coef. Of AQ factor loading		*t*-stat
Japan	All months	0.011	***	5.37	0.289	***	3.92	0.011	***	3.75
	Non-first tax months	0.012	***	5.56	0.269	***	3.33	0.010	***	3.54
	Non-first financial months	0.011	***	4.95	0.266	***	3.43	0.011	***	3.79
UK	All months	0.022	***	4.37	0.147	**	2.21	0.011	***	3.08
	Non-first tax months	0.023	***	4.32	0.155	**	2.17	0.012	***	3.38
	Non-first financial months	0.024	***	3.69	0.142	*	1.95	0.008	**	2.36
Australia	All months	0.014	***	4.89	0.042	**	2.11	0.019	***	5.22
	Non-first tax (financial) months	0.014	***	4.59	0.047	**	2.34	0.022	***	6.27
U.S.	All months	0.017	***	11.25	0.064	***	5.19	0.011	***	8.34
	Non-first tax (financial) months	0.015	***	10.21	0.048	***	3.79	0.011	***	7.71
	January	0.035	***	6.03	0.247	***	3.83	0.037	***	6.72

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TABLE 6 (continued)

Panel C: Either price (excluding firms with opening or closing price below the threshold – U.S. observations excluded from sample have mean monthly returns 0.151 percent lower than those retained)

		AQ hedge portfolio abnormal returns			Fama-Macbeth Regression			Two stage regression (2nd stage)		
		Coef. of intercept		*t*-stat	Coef. of AQ		*t*-stat	Coef. Of AQ factor loading		*t*-stat
Japan	All months	0.006	***	2.87	0.198	***	3.46	0.008	***	2.88
	Non-first tax months	0.006	***	3.09	0.178	***	2.81	0.007	***	2.61
	Non-first financial month	0.006	***	3.04	0.185	***	3.32	0.009	***	3.11
UK	All months	0.012	***	3.13	0.138	**	2.09	0.009	***	2.65
	Non-first tax months	0.013	***	3.17	0.146	**	2.06	0.011	***	3.01
	Non- first financial months	0.014	***	2.82	0.134	*	1.85	0.007	**	2.09
Australia	All months	0.008	***	3.09	0.042	**	2.11	0.011	***	3.24
	Non-first tax (financial) months	0.009	***	3.35	0.047	**	2.34	0.022	***	6.27
U.S.	All months	0.006	***	5.29	0.022	**	2.20	0.004	***	3.75
	Non-first tax (financial) months	0.004	***	3.82	0.010		0.96	0.003	***	2.98
	January	0.022	***	4.63	0.155	***	2.90	0.023	***	5.32

Panel D: Lag dollar volume (excluding firms with Lag dollar volume below the threshold-U.S. observations excluded from sample have mean monthly returns 0.724 percent greater than those retained)

		AQ hedge portfolio abnormal returns			Fama-Macbeth Regression			Two stage regression (2nd stage)		
		Coef. of intercept		*t*-stat	Coef. of AQ		*t*-stat	Coef. Of AQ factor loading		*t*-stat
Japan	All months	0.013	***	3.11	0.293	***	3.13	0.007	***	2.64
	Non-first tax months	0.014	***	3.16	0.268	***	2.98	0.006	**	2.38
	Non-first financial months	0.012	***	2.73	0.270	***	2.67	0.007	**	2.57
UK	All months	0.028	***	4.03	0.105	*	1.72	0.015	***	3.39
	Non-first tax months	0.030	***	4.02	0.107	*	1.62	0.016	***	3.55
	Non- first financial months	0.033	***	3.69	0.111	*	1.69	0.011	***	2.74
Australia	All months	0.012	***	3.7	0.052	**	2.34	0.012	***	3.26
	Non-first tax (financial) months	0.012	***	3.65	0.064	***	2.67	0.015	***	3.91
U.S.	All months	0.010	***	5.09	0.049	***	2.97	0.007	***	4.39
	Non-first tax (financial) months	0.006	***	3.52	0.032	*	1.88	0.004	**	2.46
	January	0.045	***	5.53	0.237	***	3.34	0.046	***	6.37

Notes: *, **, and *** indicate statistical significance in predicted direction at the 10 percent, 5 percent, and 1 percent levels (two-tailed), respectively.

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