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The earnings announcement premium around the globe ☆

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ABSTRACT

U.S. stocks have been shown to earn higher returns during earnings announcement months than during non-announcement months. We document that this earnings announcement premium exists across the globe. Moreover, it is not isolated to a few countries. Of the 20 countries with enough data to conduct a within-country analysis, nine exhibit a significantly positive premium. A cross-country analysis finds that the premium is strongest in countries with the greatest increase in idiosyncratic volatility around the time of their firms' earnings announcements, suggesting that uncertainty over the earnings information to be disclosed is a primary driver of the global announcement premium.

1. Introduction

It has been documented that U.S. stocks earn higher returns during months when earnings are announced than during non-announcement months. The magnitude of this earnings announcement premium has been estimated by Frazzini and Lamont (2007) to be over 7% per year.[1] While a number of potential explanations for the premium have been put forward, uncertainty remains over the reasons for its existence. There are two goals of this study. The first is to investigate the extent to which the earnings announcement premium extends globally, thereby providing out-of-sample evidence of its existence. The second is to exploit observed cross-country variations in the magnitude of the premium in order to gain insights into the factors driving this return.

Our sample consists of roughly 200,000 announcements of annual earnings from 46 foreign countries over a 20-year period. Using these announcements we estimate that the average monthly raw return to a strategy of investing in a portfolio of stocks expected to announce earnings during the month and shorting an equal dollar amount of a portfolio of expected non-announcers is 59.7 basis points, or 7.16% annualized. As shown in Fig. 1, a $1 investment in the long portfolio offset by a similar position in the short portfolio in 1991 would have grown to $4.14 by 2010. By comparison, investing $1 in a global portfolio, equally weighted by country, would have grown to $3.64 over that time period. Moreover, the long-short

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[1] Savor and Wilson (2011) and Bushman, McDermott, and Williams (2012) also document higher returns during earnings announcement months. Others have found a similar premium for shorter windows around earnings announcements. See, for example, Chari, Jagannathan, and Ofer (1988), Ball and Kothari (1991), Cohen, Dey, Lys, and Sunder (2007), and Berkman and Truong (2009). Aboody, Lehavy, and Trueman (2010) find a large earnings announcement premium for the stocks with the greatest prior 12-month return.

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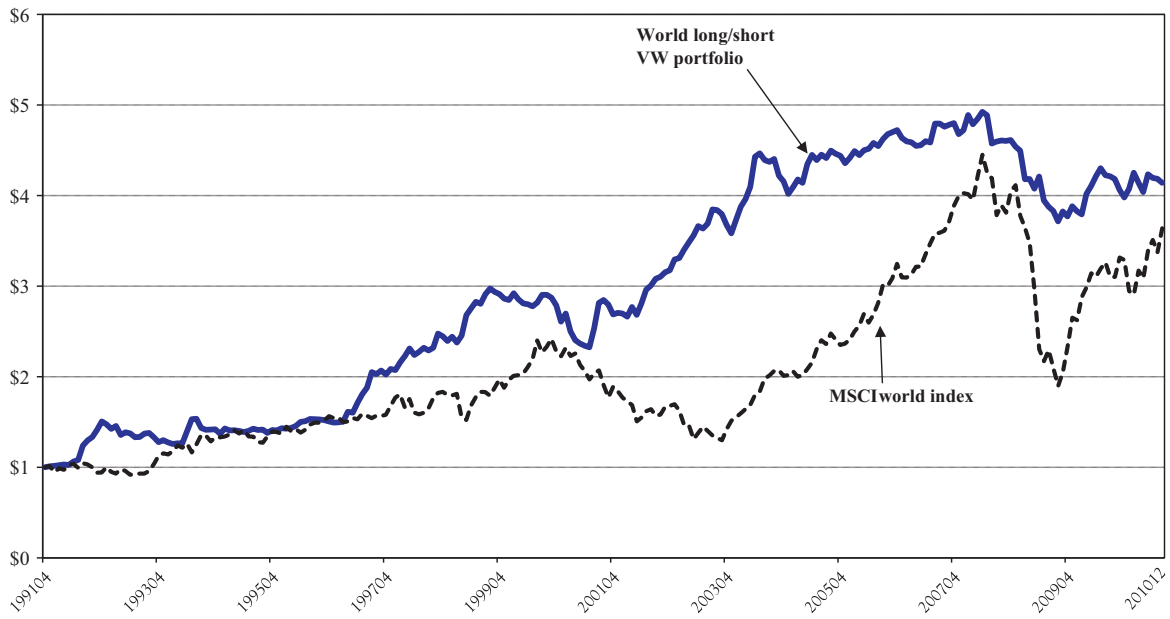

Fig. 1. Cumulative return on an investment of $1 in a long-short expected earnings announcement strategy. Our sample consists of approximately 200,000 annual earnings announcements issued by approximately 28,000 firms in 46 countries over the period from 1990 through 2009. This figure depicts the cumulative raw returns (denominated in U.S. dollars) to (1) a long-short strategy of buying all expected announcers in our global firm sample (excluding the U.S.) and shorting all expected non-announcers in a given month, where returns are value-weighted (VW) at the firm-level; and (2) a long position in the Morgan Stanley Capital International (MSCI) value-weighted global index (excluding the U.S.). Beginning with a $1 investment, returns are cumulated on a monthly basis from April 1991 through December 2010.

portfolio delivered a Sharpe ratio that is over 40% greater than that of the global portfolio.

At 59.7 basis points per month, the economic significance of the return to the long-short portfolio compares well to that of the premier asset-pricing anomalies studied in the literature: size, book-to-market, and momentum. In a global context, Fama and French (2012) find that factors associated with these three anomalies generate monthly returns of 10, 45, and 62 basis points, respectively, for the period from November 1990 to March 2011—largely coincident with the sample period we study. For the years 1981 through 2003, Hou, Karolyi, and Kho (2011) show that global size, book-to-market, and momentum factors are associated with monthly returns of 55, 51, and 63 basis points, respectively. It should be noted, however, that the amount of capital that could be deployed to exploit the announcement premium might be less than for these other anomalies since the announcer portfolio each month is limited to those firms expected to release earnings during the month.

To control for these three asset-pricing anomalies, we re-estimate the return premium using monthly cross-sectional regressions (employing a Fama-MacBeth approach), regressing individual firm returns on firm size, book-to-market ratio, and momentum, and adding an indicator variable for the firm's earnings announcement month. We also include country fixed effects. The coefficient on the indicator variable can be interpreted as the monthly earnings announcement premium. Over the 1991–2010 period, we find that it averages over 11% annually.

The positive earnings announcement premium is not isolated to just a few countries. Of the 20 countries with the greatest number of observations, it is significantly positive in nine. Across these nine countries there is substantial variation in the magnitude of the announcement premium. It ranges from a low of 63.4 basis points for France to a high of 235.5 basis points for the U.K.

We exploit these observed cross-country differences in order to gain insights into potential reasons for the premium's existence. As a prelude to our analysis, we document the patterns of daily abnormal returns, volume, and idiosyncratic volatility around the earnings announcements in our sample. Among our findings are that (a) the bulk of the premium is realized prior to (rather than after) the announcement day, (b) the higher pre-announcement returns are accompanied by reduced volume, and (c) the level of idiosyncratic volatility spikes in the three days centered on the earnings announcement date. These patterns suggest that uncertainty over the information to be released through earnings, and the accompanying abnormally high idiosyncratic volatility, cause investors to demand higher pre-announcement returns and lead to the observed earnings announcement premium. Consistent with this conjecture, we find that countries where firms have the greatest abnormal idiosyncratic volatility around earnings announcements generally have the largest announcement premium. To ensure that endogeneity issues between contemporaneous abnormal idiosyncratic volatility and returns are not responsible for this result, we repeat our analysis using each country's CIFAR financial disclosure score (Center for International Financial Analysis & Research, 1995) as an instrument for abnormal idiosyncratic volatility. The CIFAR score measures the extensiveness of a country's required financial statement disclosures and is a proxy for the amount of earnings information released, which, in turn, drives the spike in idiosyncratic volatility around earnings announcements. In line with our

prior results, the countries with the highest CIFAR scores have the greatest announcement premium.

Extending our analysis to interim earnings announcements provides additional support for the hypothesized link between abnormal idiosyncratic volatility and the announcement premium. For our international sample of interim announcements we do not find any reliable evidence of a premium, in contrast to annual announcements. At the same time, we show that the level of abnormal idiosyncratic volatility around these interim announcements is much smaller than around annual announcements, and is only marginally significant. In contrast, Frazzini and Lamont (2007) find a significant premium for interim announcements in the U.S. For these interim announcements we find that the level of abnormal idiosyncratic volatility is large and virtually identical in magnitude to that around annual announcements in the U.S.

Frazzini and Lamont (2007) conjecture that attention-based buying is a driver of the announcement premium in the U.S. We do not find evidence to support this conjecture for our international sample. An attention-based explanation predicts a positive relation between volume and returns around earnings announcements. However, when we partition countries based on the level of the pre-announcement abnormal volume of its firms (we focus on the pre-announcement window because most of the returns are concentrated there), we find a negative relation between volume and announcement premium.

Another potential explanation for the premium, put forward by Savor and Wilson (2011), is based on systematic risk. Building on the theoretical framework of Campbell and Vuolteenaho (2004), Savor and Wilson (2011) develop a model in which it is assumed that a firm's earnings news is comprised of market-wide and firm-specific components. A primary conclusion of their theoretical analysis is that the market-wide earnings component is a systematic risk and should be priced in equilibrium. The key empirical implication is that the magnitude of the earnings announcement premium should be a predictor of aggregate earnings growth. While they present evidence consistent with this hypothesis in the U.S., we do not find any reliable evidence that the earnings announcement premium predicts aggregate earnings growth internationally. In addition, we find little evidence of a systematic risk exposure in the earnings announcement premium. The estimated market beta on the long-short portfolios of the 20 countries with the greatest number of observations is, on average, zero. Moreover, the correlation in the long-short portfolio returns across countries is insignificantly different from zero.

The plan of this paper is as follows. In Section 2 we outline the data collection process and present descriptive statistics. In Section 3 we estimate the calendar-time returns to a strategy of purchasing (shorting) the shares of firms expected (not expected) to announce earnings during the month. An estimate of the magnitude of the earnings announcement premium, controlling for factors known to affect stock prices, appears in Section 4. In Section 5 we examine possible explanations for the existence of the announcement premium. Section 6 presents a summary and conclusions.

2. Data collection

Determining whether there exists a predictable return premium during earnings announcement months requires us to first decide whether the object of analysis is the *actual* or the *expected* announcement month. We choose the expected earnings announcement month for two reasons. First, using the actual month implicitly assumes that investors have perfect foresight of announcement timing (which they do not). Second, the use of the actual earnings announcement date causes an upward bias in returns due to the differential timing of good and bad news announcements (see Cohen, Dey, Lys, and Sunder, 2007). As our expectation of a firm's earnings announcement month, we use the month in which it released earnings during the prior fiscal year. We collect annual earnings announcement dates for the 20-year period from 1990 through 2009 and use them to form our expectations for the announcement months during 1991–2010.[2]

While the I/B/E/S database is a principal source of earnings announcement dates for U.S. companies, its use for foreign firms is problematic. The reason is that the date recorded for a foreign firm's earnings announcement is the date on which the earnings information was entered into the database, not necessarily the date on which the earnings were announced. For this reason we employ Bloomberg as our primary source for earnings announcement dates.[3] We use the I/B/E/S date only if Bloomberg does not record an earnings announcement for a given firm and fiscal year. If neither Bloomberg nor I/B/E/S provides a date, then the earnings announcement is recorded as missing. Most of our analysis excludes interim announcements.[4] We analyze interim announcements in Section 5, when we discuss possible explanations for the premium's existence.

Table 1 details our sample selection process. We begin with all annual earnings announcements in the Bloomberg database for which the announcement date is given and for which there is a valid SEDOL (Stock Exchange

[2] We assessed the accuracy of our expectation model for the earnings announcement month by calculating the distribution of the actual announcement month conditional on the expected month. In untabulated results, we find that the probability the actual and expected announcement months are the same ranges from 61.28% (for October) to 81.69% (for December). The month with the greatest number of expected announcements, February, has an accuracy rate of 76.94%. Later in the paper we estimate the impact that perfect foresight of the announcement month would have on the magnitude of the announcement premium.

[3] Griffin, Hirschey, and Kelly (2011) find that for a given firm and fiscal year, the date reported in the Bloomberg database is generally earlier than that reported in I/B/E/S. This is consistent with a delay in the reporting of the announcements in I/B/E/S. For an international sample of earnings announcements, DeFond, Hung, and Trezevant (2007) find that the I/B/E/S dates often differ from those reported in the financial press. Along the lines of Griffin, Hirschey, and Kelly (2011), we hand-check the accuracy of a small subset of announcement dates and find Bloomberg to be much more accurate than I/B/E/S.

[4] Among the reasons that we do not include interim announcements in most of our tests are that there are no uniform interim reporting standards across countries and that the coverage of interim announcements in both Bloomberg and I/B/E/S is spotty during the first several years of our sample period.

Table 1
Sample selection process.

This table details the compilation of our final sample. We begin with the set of all Datastream firms with valid SEDOLs as of February 2011. For these firms we collect all annual earnings announcements from Bloomberg between 1990 and 2009, eliminating those made more than 150 days after the end of the fiscal year and those without valid period-end dates. For fiscal periods without valid Bloomberg data, we use I/B/E/S as our source for earnings announcements. We exclude from our final sample (a) those securities that do not represent common equity, (b) those firms for which Datastream does not provide return data for the market of primary listing, as well as the exchange listing code and the listing currency, (c) firm-years where there is more than one annual earnings announcement, (d) delisted firms without a valid delisting date, (e) firms with beginning-of-year market capitalization of less than $1 million, and (f) securities that represent American Depositary Receipts (ADRs). For some of our analyses we need to calculate the book-to-market (B/M) ratio. For those analyses we eliminate firms without the necessary information to make this calculation.

	No. of annual announcements	No. of firms
Bloomberg earnings announcements with a valid SEDOL code on Datastream	260,659	34,652
Less: Bloomberg announcements without valid period-end dates	(15,396)	(478)
Less: Bloomberg announcements greater than 150 days old	(10,686)	(944)
Equals: total Bloomberg announcements	234,577	33,230
Plus: non-stale, nonoverlapping I/B/E/S earnings announcements	71,327	10,992
Less: years with more than one annual announcement per year	(2,287)	(30)
Equals: total earnings announcements	303,617	44,192
Less: firms with missing returns, missing exchange code and exchange currency; non-common equity; delisted firms without valid delisting date	(90,714)	(14,554)
Equals: announcements matched to valid Datastream data	212,903	29,638
Less: firms with less than $1 million market capitalization	(1,956)	(279)
Less: ADR issues	(10,236)	(587)
Equals: final earnings announcement sample	**200,711**	**28,772**
Less: firms without book-to-market data	(18,786)	(989)
Equals: final earnings announcement sample for analyses requiring B/M ratio	**181,925**	**27,783**

Daily Official List) code on Thomson Financial's Datastream database. This totals 260,659 annual announcements issued by 34,652 firms. To remain in the sample we also require either that the fiscal year-end date be given or that it be reliably estimable from the fiscal year-end dates of adjacent years. Imposing this requirement reduces our announcement sample by approximately 6%. We also dropped announcements that are made more than 150 days after the fiscal year-end date, under the assumption that such a long delay likely reflects either an erroneous announcement date or fiscal year-end date. This lowers our announcement sample by roughly 4%, leaving a final Bloomberg sample of 234,577 announcements by 33,230 firms.

To this set we add all those valid annual announcements in the I/B/E/S database that are not in Bloomberg. As before, a valid I/B/E/S announcement is one that includes the earnings amount and announcement date, and is issued no more than 150 days after the end of the fiscal year.[5] Using I/B/E/S we are able to supplement our sample by 71,327 announcements issued by 10,992 firms. As a last step, we eliminate instances in which a firm makes two annual earnings announcements in the same calendar year (about 3% of the total). Our final earnings announcement sample consists of 303,617 annual announcements made by 44,192 firms.

Our source for stock returns and other firm-specific financial information is Datastream, for which we convert all local currency-denominated returns to returns denominated in U.S. dollars.[6] Of the 44,192 firms with valid announcements, we drop those for which Datastream does not provide (a) return data for the market of primary listing, (b) the exchange listing code, and (c) the listing currency. We also exclude delisted firms without a valid delisting date and firms whose traded securities are not common equity [using the filters described in Griffin, Kelly, and Nardari (2010)]. Imposing these criteria reduces the number of firms in our sample by approximately one-third. We also drop firms with concurrently trading American Depositary Receipts (ADRs) in the U.S. (only for the time period during which the ADR is trading). We do so in order to ensure that the combination of a U.S. announcement premium and arbitrage are not driving the return pattern we document in foreign firms' shares. Finally, we exclude firms for those months in which their beginning-of-month market capitalizations (denominated in U.S. dollars) are less than $1 million. This leaves us with a final sample of 200,711 annual earnings announcements issued by 28,772 firms. For much of our analysis we require information on firms' book values. When we impose this requirement, our sample is reduced by 18,786 announcements and 989 firms.

The distribution of annual earnings announcements across years is shown in Table 2. The number of announcements in our sample increases almost monotonically over our sample period, more than doubling over the last ten years, from 9,439 in 1999 to 19,632 in 2009. This jump reflects the increasing comprehensiveness of the Bloomberg database over time. Bloomberg accounted for 44% of all sample announcements during the 1990–1999 period; that number increased to 93% during the years 2000–2009.

Our sample encompasses annual earnings announcements from 46 countries. For each country, Table 3 reports the average market capitalization of the firms

[5] Unlike some Bloomberg announcements, I/B/E/S always gives the fiscal year-end date.

[6] We use exchange rate data provided by Bloomberg to perform the conversion.

Table 2
Distribution of annual earnings announcements, by year.
This table reports the yearly number of annual earnings announcements coming from Bloomberg and from I/B/E/S for the period 1990–2009. See Table 1 for a detailed description of our sample.

Year	Number of announcements		
	Bloomberg	I/B/E/S	Total
All years	167,701	33,010	200,711
1990	8	550	558
1991	14	1,116	1,130
1992	39	1,287	1,326
1993	165	1,752	1,917
1994	309	2,069	2,378
1995	556	2,918	3,474
1996	1,299	3,050	4,349
1997	3,008	3,908	6,916
1998	4,954	3,240	8,194
1999	6,934	2,505	9,439
2000	8,754	2,159	10,913
2001	10,074	1,970	12,044
2002	11,695	1,222	12,917
2003	13,668	1,031	14,699
2004	14,861	832	15,693
2005	16,271	764	17,035
2006	17,742	843	18,585
2007	19,071	673	19,744
2008	19,173	595	19,768
2009	19,106	526	19,632

included in our sample, as a percentage of total market capitalization, as well as the number of earnings announcements provided.[7] Mean coverage across the countries in our sample is 41.8% of total market cap.[8] The mean (median) number of annual announcements per country is 4,363 (1,921). Japan has by far the largest number of announcements in our sample (over 40,000), covering 50.9% of the Japanese firms by market cap. Seven countries contribute at least 10,000 annual announcements. Of those, average coverage ranges from 30.0% (China) to 67.0% (Australia).

3. The global annual earnings announcement premium: calendar-time returns

We begin our analysis by estimating the calendar-time returns to a trading strategy based on the expected annual earnings announcement month. To do so, we construct two portfolios at the end of each month $t-1$ of our sample period. The first, referred to as the "announcer portfolio," is comprised of all firms that are expected to announce annual earnings during month t. The second, referred to as the "non-announcer portfolio," is comprised of all firms that are not expected to announce earnings during the month. Each portfolio's value-weighted raw return for month t is then calculated. (The weight for each component firm is based on the firm's dollar-denominated

Table 3
Distribution of earnings announcements by country.
This table presents the distribution of earnings announcements by country for our sample of 200,711 annual announcements over the period 1990–2009. The average coverage ratio for a country is equal to the total market capitalization of our firm sample in that country (denominated in U.S. dollars), measured at the end of each year, divided by the total market capitalization for that country (as reported by the World Bank), and then averaged across years. World Bank market capitalization includes only listed domestic companies at the end of the year, exclusive of investment companies, mutual funds, and other collective investment vehicles. See Table 1 for a detailed description of our sample.

Country	Number of announcements	Average coverage ratio (mkt cap) (%)
Japan	40,635	50.9
UK	15,450	36.9
Australia	13,574	67.0
South Korea	12,511	34.6
China	11,233	30.0
Canada	11,212	59.3
Taiwan	10,074	45.2
Malaysia	9,524	63.7
Hong Kong	8,259	53.1
France	5,902	43.6
India	5,492	30.2
Singapore	5,127	46.7
Germany	4,679	37.6
Thailand	4,573	57.8
Sweden	3,546	44.6
South Africa	3,375	27.8
Indonesia	2,700	46.1
Italy	2,369	44.5
Brazil	2,250	26.3
Greece	2,183	44.2
Norway	2,041	46.8
Switzerland	1,936	37.9
Turkey	1,922	48.3
Denmark	1,919	40.3
Netherlands	1,717	43.2
Israel	1,634	32.7
Chile	1,482	65.0
New Zealand	1,445	61.2
Poland	1,438	34.4
Finland	1,378	35.9
Belgium	1,329	52.3
Philippines	1,232	37.3
Spain	1,152	20.1
Mexico	963	33.4
Ireland	669	71.1
Argentina	644	21.7
Peru	549	46.3
Portugal	507	38.6
Pakistan	489	24.1
Austria	416	23.4
Jordan	334	46.8
Hungary	264	46.8
Czech Republic	258	46.6
Colombia	233	36.0
Luxembourg	76	30.9
Venezuela	16	0.0
Total	200,711	41.8

[7] Market capitalization data are provided by the World Bank and encompass the universe of stocks listed in each country. According to their definition, "[m]arket capitalization…is the share price times the number of shares outstanding. Listed domestic companies are the domestically incorporated companies listed on the country's stock exchanges at the end of the year. Listed companies do not include investment companies, mutual funds, or other collective investment vehicles." For each country-year, the year-end coverage percentage is calculated. The average over these years is the number reported in the table for that country.

[8] Including concurrently traded ADRs in the sample would increase mean coverage across the countries in our sample to 62.7% of total market cap (untabulated).

market value as of the end of month $t-1$.) The difference between the month t returns on these two portfolios is equal to the return from a trading strategy of purchasing the expected month t announcers and shorting the expected month t non-announcers. This return is sometimes referred to below as the "long-short return." If there are fewer than ten observations in either the announcer or non-announcer portfolio in a given month, that month is not included in our calendar-time return calculations.[9]

The first row of Table 4 presents the average monthly raw returns over our entire sample period. For the announcer portfolio, the average return is 123.7 basis points. The corresponding return for the non-announcer portfolio is 63.9 basis points. Both returns are significantly greater than zero. The long-short return is also significantly positive, averaging 59.7 basis points monthly, or 7.16% on an annual basis. The monthly Sharpe ratio for this strategy is 0.20. We alternatively compute the earnings announcement premium by value-weighting firm returns *within* country and then taking a value-weighted average of the country returns using country-level total market capitalization at the end of the prior calendar year.[10] Using this methodology, the average monthly return to the announcer portfolio drops to 116.4 basis points, while the non-announcer portfolio's average monthly return increases to 67.9 basis points (see the second row of Table 4). The long-short return is 48.5 basis points, or 5.82% annualized, with a monthly Sharpe ratio of 0.216. These numbers are not significantly different from those computed when firms are value-weighted across countries. As a point of reference, we replicate this analysis for U.S. firms. As reported in the last row of Table 4, the long-short return for the U.S. is 75.9 basis points per month, or 9.11% on an annual basis, with a corresponding Sharpe ratio of 0.217.

From Table 4 we also see that investing in a long portfolio comprised of expected announcing firms delivers a Sharpe ratio in the U.S. that is 37% higher than that of a long portfolio of expected non-announcers (0.271 compared to 0.198). Internationally, it is 79% higher (0.231 compared to 0.129) when value-weighting at the firm level and 84% higher (0.268 compared to 0.145) when value-weighting at the country level. The long-short portfolio has a slightly more modest Sharpe ratio: 0.217 for the U.S., 0.20 for a firm-level global portfolio, and 0.216 for a country-level global portfolio. However, since these long-short portfolios have virtually no correlation with their respective market indexes, adding them to an investor's opportunity set can lead to a substantial improvement in the investor's Sharpe ratio.[11]

4. The global earnings announcement premium: regression analysis

In this section we continue our analysis of the earnings announcement premium, within a Fama-MacBeth regression framework. For each month t of our sample period we estimate the following regression[12]:

$$\mathrm{Ret}_{ijt} = \alpha_t + \beta_{1t} ExpAnn_{ijt} + \beta_{2t} Mom_{ijt} + \beta_{3t} MktCap_{ijt} + \beta_{4t} BTM_{ijt} + \sum_{j=1}^{46} \gamma_j Country_j + \varepsilon_{ijt}, \quad (1)$$

where

Ret_{ijt}	natural log of one plus the raw return for firm i in country j during month t;
$ExpAnn_{ijt}$	indicator variable taking on a value of one if firm i in country j is expected to announce annual earnings during month t, and zero otherwise;
Mom_{ijt}	natural log of one plus the raw return for firm i in country j over months $t-11$ to $t-1$[13];
$MktCap_{ijt}$	natural log of market capitalization (in U.S. dollars) of firm i in country j at the end of month $t-1$;
BTM_{ijt}	natural log of the book-to-market ratio for firm i in country j as of the end of the fiscal year preceding month t[14];
$Country_j$	fixed-effect dummy variable, taking on a value of one for each firm in country j, and zero otherwise; and
ε_{ijt}	regression residual for firm i in country j for month t.

The coefficient, β_{1t}, on the indicator variable, $ExpAnn_{ijt}$, can be interpreted as the average incremental monthly return generated during expected annual earnings announcement months relative to non-announcement months. In estimating its value, we control for the firm's prior return, the firm's size, and its book-to-market ratio, all factors that have been shown to be related to firms' stock returns. As mentioned previously, requiring that book value information be available reduces our sample of firms by 989, or approximately 3.4% of the total.

Table 5 reports the results of our main analysis. Over our entire sample period, the average coefficients on the momentum, size, and book-to-market control variables are all positive and significant (see Panel A). The average coefficient on the indicator variable, $ExpAnn_{ijt}$, is a significantly positive

[9] Only four of the 240 months of our sample period fall short of this requirement. All occur at the beginning of the sample period, January through April, 1991.

[10] We require there to be at least five observations in a country's announcer and non-announcer portfolios in a given month in order to include that country in the month's return calculations.

[11] For example, the MSCI World Index, excluding the U.S., has a Sharpe ratio of 0.136 over our sample period. Combining this asset with that of a country-level long-short portfolio yields a maximum Sharpe ratio of 0.303, with a 73% allocation to the long-short portfolio. We obtain qualitatively similar results when we consider a combination of (a) a U.S. market index and a U.S. long-short portfolio or (b) the MSCI World Index, including the U.S., and a country-level global portfolio

(*footnote continued*)
which includes the U.S. While these are ex post calculations and do not account for transactions costs, they nevertheless demonstrate the potential appeal to investors of incorporating the long-short portfolio into their investment opportunity set.

[12] For all full-sample regressions, we only include months where there are at least ten announcing and ten non-announcing firms.

[13] Results are quantitatively similar when we calculate momentum using months $t-12$ to $t-2$.

[14] If month t is less than four months after the previous fiscal year-end, then we calculate the book-to-market ratio using data as of the end of the fiscal year prior to that. The approximately 47,000 observations with a negative book value are dropped from this analysis.

Table 4
Returns on global portfolios of expected announcers and expected non-announcers.

This table shows average monthly raw returns, value-weighted at the firm level, on global portfolios of firms (46 foreign countries) for the period 1991–2010 ("firm-level global portfolios"). All returns are denominated in U.S. dollars. At the beginning of every calendar-month, stocks are assigned to one of two portfolios—expected announcers and expected non-announcers—using annual announcement dates predicted based on the previous year's actual annual announcement month. In order for a stock to be included in these portfolios in a given year, it must have an expected annual announcement month. Conditional on having an expected annual announcement month, each stock appears in the portfolio of expected announcers once each year and in the portfolio of non-announcers 11 times during the year. The long-short portfolio each month is comprised of a long position in expected announcers that month and a short position in expected non-announcers. To be included in our calculations, a monthly portfolio of expected announcers or a monthly portfolio of expected non-announcers must be comprised of at least ten firms; a monthly long-short portfolio must have at least ten firms on the long side and ten firms on the short side. We also report the long, short, and long-short average monthly returns, value-weighted within country and then value-weighted across countries ("country-level global portfolios"). Country value-weightings are computed each year, using country-level total market capitalizations as of the end of the prior calendar year. The last column provides the number of months that have sufficient data to calculate long-short portfolio returns. For comparison purposes, we report the corresponding portfolio returns for U.S. firms. *t*-Statistics appear below the reported monthly returns. ***=Significant at the 1% level; **=significant at the 5% level; *=significant at the 10% level. Sharpe ratios are also shown for each portfolio. See Table 1 for a detailed description of our sample.

	Portfolio of expected announcers			Portfolio of expected non-announcers			Long-short portfolio		Number of months included in long-short portfolio return calculations
	Monthly raw return	Sharpe ratio	Average monthly number of firms in portfolio	Monthly raw return	Sharpe ratio	Average monthly number of firms in portfolio	Monthly raw return	Sharpe ratio	
Firm-level global portfolios	**1.237*****	0.231	850	**0.639***	0.129	9294	**0.597*****	0.200	236
t-Statistic	*3.24*			*1.94*			*3.16*		
Country-level global portfolios	**1.164*****	0.268	46	**0.679****	0.145	46	**0.485*****	0.216	236
t-Statistic	*3.99*			*2.18*			*3.17*		
United States portfolio	**1.648*****	**0.271**	445	**0.889*****	**0.198**	4808	**0.759*****	**0.217**	240
t-Statistic	***4.20***			***3.07***			***3.35***		

Table 5
Fama-MacBeth regression analysis of global earnings announcement portfolio returns.

This table reports the average coefficients for monthly regressions of:

$$\mathrm{Ret}_{ijt} = \alpha + \beta_1 ExpAnn_{ijt} + \beta_2 Mom_{ijt} + \beta_3 MktCap_{ijt} + \beta_4 BTM_{ijt} + \sum_{j=1}^{46} \gamma_j Country_j + \varepsilon_{ijt},$$

where Ret_{ijt} is the natural log of one plus the raw return during month t for firm i in country j, denominated in U.S. dollars; $ExpAnn_{ijt}$ is an indicator variable equal to one if firm i of country j is expected to announce annual earnings in month t and is equal to zero, otherwise; Mom_{ijt} is the natural log of one plus the raw return for firm i of country j over months $t-1$ through $t-11$, denominated in U.S. dollars; $MktCap_{ijt}$ is the log of the market capitalization of firm i in country j at the end of month $t-1$, denominated in U.S. dollars; BTM_{ijt} is the log of the book-to-market ratio for firm i of country j as of the end of the prior fiscal year; $Country_j$ is an indicator variable equal to one for all firms in country j and equal to zero, otherwise; ε_{ijt} is the regression residual for firm i of country j in month t. The coefficients on the country indicator variables are not reported in the table. The average number of observations in the monthly regressions as well as the regression R^2 are also reported in the table. Panel A reports results for the full sample period, and by year. Panel B reports Fama-MacBeth regression estimates for size groups (small/medium/large) based on NYSE percentile break points. Firms are ranked based on market capitalization at the beginning of each month; those with market values below the 20th percentile are classified as small, those that fall between the 20th and 50th percentile are classified as medium-sized, and those with market values above the 50th percentile are classified as large. The median market capitalization value of firms within each size group is reported in parentheses. These regressions include only those months for which there are at least five announcer and five non-announcer firms within each size group during the month. Below each coefficient value is the corresponding *t*-statistic. Significance is determined based on the time-series distribution of equally weighted monthly coefficient estimates. ***=Significant at the 1% level; **=significant at the 5% level; *=significant at the 10% level. See Table 1 for a detailed description of our sample.

Panel A: Full sample period and by year

Year	Avg. # of monthly obs	Intercept	*ExpAnn*	*Mom*	*MktCap*	*BTM*	*R*-squared (%)
1991–2010	8,966	−1.302	**0.955*****	**0.008*****	**0.081****	**0.511*****	18.46
		−1.29	*11.08*	*4.43*	*2.29*	*9.21*	
1991	458	2.226	**0.836****	**0.018****	−0.159	−0.267	18.72
		0.39	*2.52*	*2.21*	*−0.41*	*−0.63*	
1992	846	−3.801	0.668	**0.018****	0.213	0.382	27.03
		−1.07	*1.19*	*2.53*	*1.01*	*1.65*	
1993	1,008	**5.051***	0.689	0.007	−0.181	**0.912*****	25.34
		1.69	*1.59*	*0.88*	*−0.83*	*3.37*	
1994	1,294	2.039	**0.556***	−0.002	**−0.122**	**0.411*****	29.99
		0.76	*1.96*	*−0.56*	*−0.63*	*2.96*	
1995	1,672	−0.981	**1.107*****	**0.013*****	0.151	0.211	20.67

Table 5 (*continued*)

Panel A: Full sample period and by year

		−0.92	*4.46*	*3.44*	*1.41*	*0.84*	
1996	2,596	2.361	**0.852****	**0.017*****	0.088	0.128	18.48
		0.48	*2.73*	*3.51*	*0.91*	*0.72*	
1997	3,384	−0.209	**1.518*****	**0.031*****	0.211	0.192	27.09
		−0.06	*3.29*	*3.51*	*1.06*	*1.46*	
1998	5,252	−7.148	**1.931*****	−0.012	0.118	0.100	30.91
		−1.31	*5.88*	*−0.91*	*0.71*	*0.45*	
1999	6,384	−3.109	**1.783*****	0.009	−0.016	−0.102	18.10
		−1.57	*4.03*	*1.31*	*−0.16*	*−0.30*	
2000	7,586	**−2.821***	**0.924****	−0.003	0.157	**1.025***	16.26
		−1.84	*2.19*	*−0.25*	*0.58*	*1.89*	
2001	9,172	**−4.171***	**1.561*****	0.016	0.165	**1.104*****	19.17
		−1.87	*4.26*	*1.28*	*1.46*	*3.95*	
2002	10,492	**−3.407****	**0.783****	**0.020****	0.146	**1.061*****	15.56
		−2.51	*2.72*	*2.37*	*1.32*	*9.99*	
2003	11,448	**3.098***	**1.359*****	−0.001	0.066	**0.735*****	14.95
		1.80	*3.87*	*−0.08*	*0.53*	*5.25*	
2004	13,106	−1.229	**0.648****	0.006	0.158	**0.819*****	14.97
		−0.38	*2.56*	*1.45*	*1.31*	*4.79*	
2005	14,068	**−5.759****	0.196	**0.011*****	0.103	**0.468*****	13.95
		−2.66	*0.67*	*5.04*	*0.98*	*5.31*	
2006	15,692	3.518	**0.531*****	**0.010****	**0.145*****	**0.723*****	13.07
		1.57	*3.98*	*2.46*	*2.87*	*4.68*	
2007	17,316	**−7.369****	**1.231*****	**0.009****	0.018	**0.516****	16.88
		−2.72	*3.85*	*2.57*	*0.14*	*2.72*	
2008	18,384	−6.321	0.056	**0.017***	0.018	**0.622*****	15.52
		−0.41	*0.13*	*1.78*	*0.46*	*4.69*	
2009	18,550	**3.356***	**0.991****	**−0.025****	0.131	**0.468*****	14.13
		1.84	*2.27*	*−2.37*	*0.88*	*2.84*	
2010	18,388	−0.184	**0.874***	**0.007****	0.093	**0.465*****	12.19
		−0.09	*1.84*	*2.29*	*0.72*	*3.34*	

Panel B: Earnings announcement premium across size groups

Firm size	Avg. *N*	Intercept	*ExpAnn*	*Mom*	*MktCap*	*BTM*	*R*-squared (%)
Small ($56m)	6,707	0.293	**1.139*****	**0.011*****	−0.110	**0.553*****	18.7
		0.25	*7.99*	*4.94*	*−1.51*	*8.46*	
Medium ($634m)	1,336	−0.571	**0.959*****	**0.009*****	0.064	**0.339*****	26.9
		−0.54	*6.59*	*3.71*	*0.85*	*4.23*	
Large ($3,003m)	874	−0.632	**0.412*****	0.002	0.072	**0.368*****	27.3
		−0.74	*2.93*	*0.56*	*1.43*	*4.35*	
Test of difference (Small − Large) = 1.139 − 0.412 = **0.727***** (Z = 3.63; $p < 0.01$)							

and economically large 0.955, which means that the average monthly incremental return during annual earnings announcement months is 95.5 basis points, or 11.46% on an annual basis.[15] Moreover, the coefficient on $ExpAnn_{ijt}$ is positive in every single year.[16] For 16 of the 20 years, it is significantly greater than zero. This suggests an earnings announcement premium that is robust over time.[17]

Regardless of the drivers of the premium, abnormal returns should reverse themselves over the 11 months following the expected earnings announcement month; over a full 12-month period, firms' abnormal returns should be insignificantly different from zero. To test this, we estimate the average premium τ months after the expected announcement month, $\tau = 1,2,\ldots,11$. We do so by re-estimating regression (1), replacing the indicator variable, $ExpAnn_{ijt}$, by an indicator variable that equals one if month t is τ months

[15] Re-estimating this regression including ADRs yields an almost identical average monthly incremental return during expected earnings announcement months (0.948, untabulated).

[16] In untabulated results, we also find that the average incremental return is significantly positive regardless of the month in which earnings are expected to be announced.

[17] The average monthly incremental return of 0.955 is higher than the previously reported 0.597 average monthly return to the long-short portfolio. The difference between these two returns is due mainly to the fact that the observations in regression (1) are equally weighted, whereas those in the long-short portfolio are value-weighted. We

(*footnote continued*)
confirm this by estimating a value-weighted least-squares version of regression (1). In untabulated results, we find the average coefficient on $ExpAnn_{ijt}$ in the re-estimated regression to be 0.537, which is much closer to the long-short portfolio return. That the average monthly return decreases when observations are value-weighted is consistent with our finding, reported below, that the larger firms in our sample have a lower announcement premium.

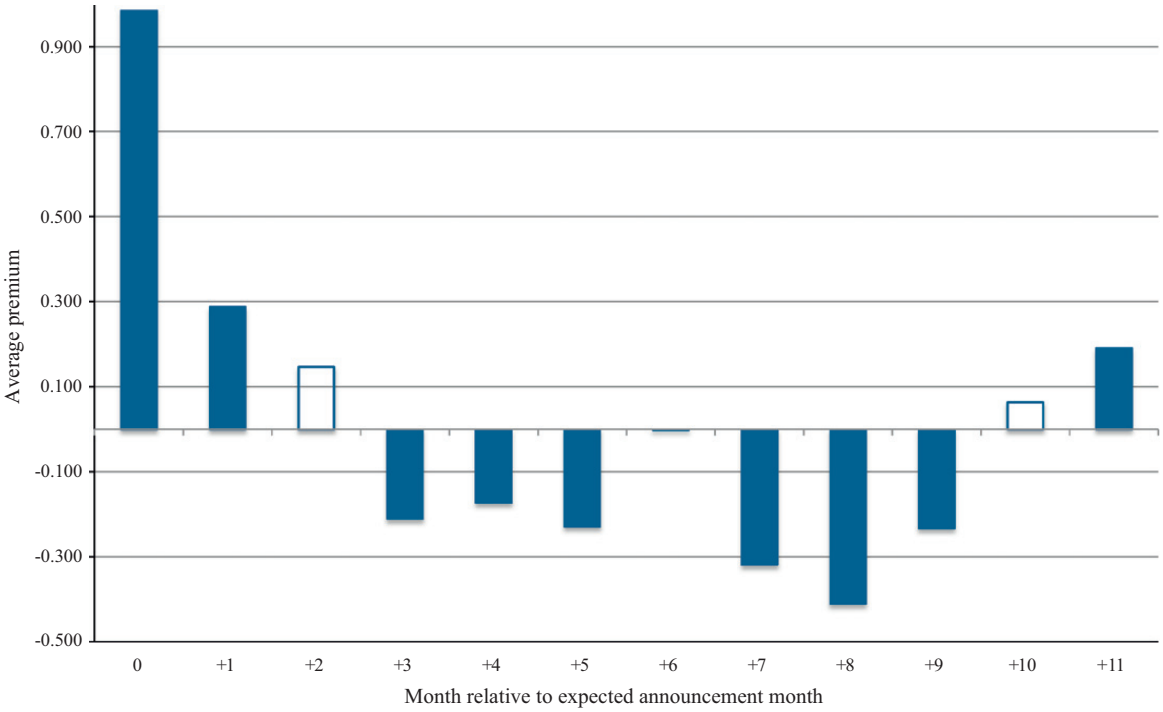

Fig. 2. Monthly premiums. Our sample consists of approximately 200,000 annual earnings announcements issued by approximately 28,000 firms in 46 countries over the period from 1990 through 2009. This figure charts the average premium τ months after the expected announcement month, $\tau = 0,1,2,...,11$, for expected announcement months from January 1991 through January 2010. We calculate the month τ premium by using the specification described in Table 5, replacing the indicator variable, $ExpAnn_{ijt}$, by an indicator variable that equals one if month t is τ months after the expected announcement month and zero, otherwise. A solid bar indicates that the average premium for that month is significantly different from zero.

after the expected announcement month and zero, otherwise. As shown in Fig. 2, only the expected announcement month and the two surrounding months have a significantly positive average premium. In six of the months the average premium is significantly negative. The sum of the coefficients on the indicator variables for the 12 months is an insignificant 9.6 basis points.

We next analyze whether firm size affects the magnitude of the annual earnings announcement premium. To do so we partition our sample into small, medium, and large firms, according to market capitalization at the end of each month, and then estimate regression (1) for each subsample.[18] As presented in Table 5, Panel B, the average coefficient on $ExpAnn_{ijt}$ is significantly greater for the smallest firms than it is for the largest ones (1.139 compared to 0.412).[19] The smallest firms generate an annual earnings announcement premium that is 72.7 basis points per month (8.72% annualized) greater than that of the largest firms. That small firms have the highest returns is characteristic of many anomalies in the finance and accounting literature. However, it contrasts with Frazzini and Lamont (2007), whose findings suggest either that size plays an insignificant role or that the larger firms generate a higher announcement premium.

We also find significant variation in the magnitude of the earnings announcement premium across countries. When calculating returns for a particular country, we only include those months for which at least five firms within that country are expected to announce earnings and at least five firms are expected not to make an announcement. Our country-level analysis encompasses only those countries for which at least 25% of the months in our sample period (60 out of 240 months) satisfy this criterion.

The results of our country analysis are presented in Table 6. For reference, we also include the U.S. For U.S. firms the average coefficient on the indicator variable, $ExpAnn_{ijt}$, is equal to 0.326, which is equivalent to an annualized earnings announcement premium of 3.91%.[20] Of the 20 foreign countries included in this analysis, 16 have a positive coefficient on $ExpAnn_{ijt}$. Nine are significantly greater than zero, ranging in magnitude from 0.634 for France (equivalent to an annualized announcement premium of 7.61%) to 2.355 for the U.K. (28.26%, annualized). All nine are greater in magnitude than that for the U.S. The average coefficient on $ExpAnn_{ijt}$ is negative in only four countries; none, though, is significantly different from zero.[21]

That the premium is significantly positive in almost half of the countries listed in Table 6 is not due to

[18] Small firms are defined as those with market capitalizations falling within the lowest two NYSE deciles, while large firms are those within the top five deciles.

[19] We obtain quantitatively similar results when we partition our firms evenly into quintiles according to size.

[20] This is lower than the long-short portfolio return of 9.11%. Contributing to the lower return is the equal-weighting of observations that occurs in the regression framework (as compared to the value-weighting of observations in the long-short portfolios) and the documented lower premium for small firms in the U.S. (Frazzini and Lamont, 2007).

[21] The intercept for China, 12.05, stands out for being unusually large. It is due primarily to non-announcer returns of over 90% during March and April, 2001 and average non-announcer returns of over 60% during February–April, 2007.

Table 6
Fama-MacBeth regression analysis of earnings announcement premium, by country.
This table reports the average coefficients for monthly regressions of:

$$Ret_{ijt} = \alpha + \beta_1 ExpAnn_{ijt} + \beta_2 Mom_{ijt} + \beta_3 MktCap_{ijt} + \beta_4 BTM_{ijt} + \varepsilon_{ijt}$$

for each country j, where Ret_{ijt} is the natural log of one plus the raw return during month t for firm i in country j, denominated in U.S. dollars; $ExpAnn_{ijt}$ is an indicator variable equal to one if firm i of country j is expected to announce annual earnings in month t and is equal to zero, otherwise; Mom_{ijt} is the natural log of one plus the raw return for firm i of country j over months $t-1$ through $t-11$, denominated in U.S. dollars; $MktCap_{ijt}$ is the log of the market capitalization of firm i in country j at the end of month $t-1$, denominated in U.S. dollars; BTM_{ijt} is the log of the book-to-market ratio for firm i of country j as of the end of the prior fiscal year; ε_{ijt} is the regression residual for firm i of country j in month t. Each country's regressions are estimated only for those months where there are at least five announcer and five non-announcer firms. We report results for those countries for which at least 60 months satisfy this criterion. In addition to the coefficient estimates, the table reports the number of months for which the country regressions are estimated, the average number of observations per month, and the R^2. Below each coefficient value is the corresponding t-statistic. Significance is based on the time-series distribution of equally weighted monthly coefficient estimates. ***=Significant at the 1% level; **=significant at the 5% level; *=significant at the 10% level. See Table 1 for a detailed description of our sample.

Country	Number of months	Avg. # of monthly obs.	Average coefficient estimates					R-squared (%)
			Intercept	*ExpAnn*	*Mom*	*MktCap*	*BTM*	
United States	240	5,151	**−1.570***	**0.326*****	**0.009*****	**0.135****	**0.389*****	3.54
			−1.93	*2.77*	*3.32*	*2.35*	*4.84*	
Japan	203	2,258	**−1.832***	**0.799*****	−0.003	**0.128****	**0.701*****	5.45
			−1.85	*4.26*	*−0.78*	*2.06*	*6.63*	
China	62	745	**12.050****	−0.306	−0.004	**−0.726***	0.435	14.36
			2.30	*−1.04*	*−0.33*	*−1.86*	*0.88*	
South Korea	134	716	0.244	−0.279	0.002	−0.078	**1.224*****	6.17
			0.14	*−0.53*	*0.42*	*−0.53*	*5.90*	
UK	235	686	−1.123	**2.355*****	**0.020*****	0.078	**0.346*****	4.59
			−1.50	*13.81*	*8.33*	*1.39*	*4.19*	
Australia	164	679	−0.063	**1.687*****	**0.014*****	0.014	**0.477*****	5.55
			−0.06	*4.85*	*3.91*	*0.19*	*3.96*	
Taiwan	60	627	1.287	−0.269	0.009	−0.101	0.551	10.96
			0.47	*−0.54*	*1.22*	*−0.49*	*0.84*	
Canada	204	554	−0.342	0.026	**0.016*****	0.043	**0.329****	5.66
			−0.33	*0.09*	*5.14*	*0.65*	*2.43*	
Malaysia	181	529	**−3.686****	0.23	−0.004	**0.321****	**0.703*****	7.19
			−2.29	*0.86*	*−0.81*	*2.67*	*5.69*	
Hong Kong	172	462	1.505	0.366	0.004	−0.112	**0.586*****	6.91
			0.94	*1.00*	*0.99*	*−0.92*	*3.86*	
India	97	459	0.315	0.287	0.006	0.151	**1.052****	9.69
			0.11	*0.63*	*0.81*	*0.69*	*2.38*	
Germany	105	348	−1.580	**0.937****	**0.013****	**0.140****	**0.281***	6.62
			−1.63	*2.01*	*2.22*	*2.05*	*1.73*	
France	144	339	0.129	**0.634***	**0.014****	0.026	**0.342****	7.68
			0.15	*1.77*	*2.73*	*0.41*	*2.27*	
Singapore	154	303	−2.056	**1.473*****	0.000	0.172	**0.698******	8.36
			−1.30	*4.46*	*0.01*	*1.52*	*4.28*	
Thailand	67	256	−1.722	−0.061	0.002	0.139	**0.964*****	8.43
			−0.78	*−0.07*	*0.36*	*0.66*	*3.02*	
Sweden	72	233	0.477	0.478	0.007	−0.020	0.115	7.94
			0.30	*0.94*	*1.05*	*−0.15*	*0.50*	
South Africa	151	196	**−5.300*****	**1.965*****	**0.008***	**0.429*****	**0.576****	7.85
			−4.10	*4.46*	*1.91*	*3.99*	*2.54*	
Switzerland	61	125	0.516	**0.919*****	**0.016***	0.042	**0.328****	11.25
			0.41	*3.17*	*1.69*	*0.45*	*1.96*	
Brazil	63	118	−2.056	0.408	**−0.010**	0.147	**0.549****	9.31
			−0.89	*0.44*	*−1.13*	*0.75*	*2.13*	
Denmark	76	109	−0.505	0.812	**0.031*****	0.004	**0.620****	10.81
			−0.26	*1.63*	*4.00*	*0.02*	*2.21*	
New Zealand	92	72	0.041	**1.172****	**0.042*****	−0.021	0.317	15.50
			0.02	*2.22*	*2.90*	*−0.13*	*1.30*	

significant positive cross-country correlations in long-short portfolio returns. In untabulated results we find that the average pairwise return correlation for these nine countries is only 0.04 (insignificantly different from zero) and there are almost as many negative correlations as positive ones. Further, of those that are positive, only five are significantly different from zero.

Finally, we estimate the effect that imperfect knowledge of the announcement month has on the magnitude of the announcement premium. We do so by re-estimating

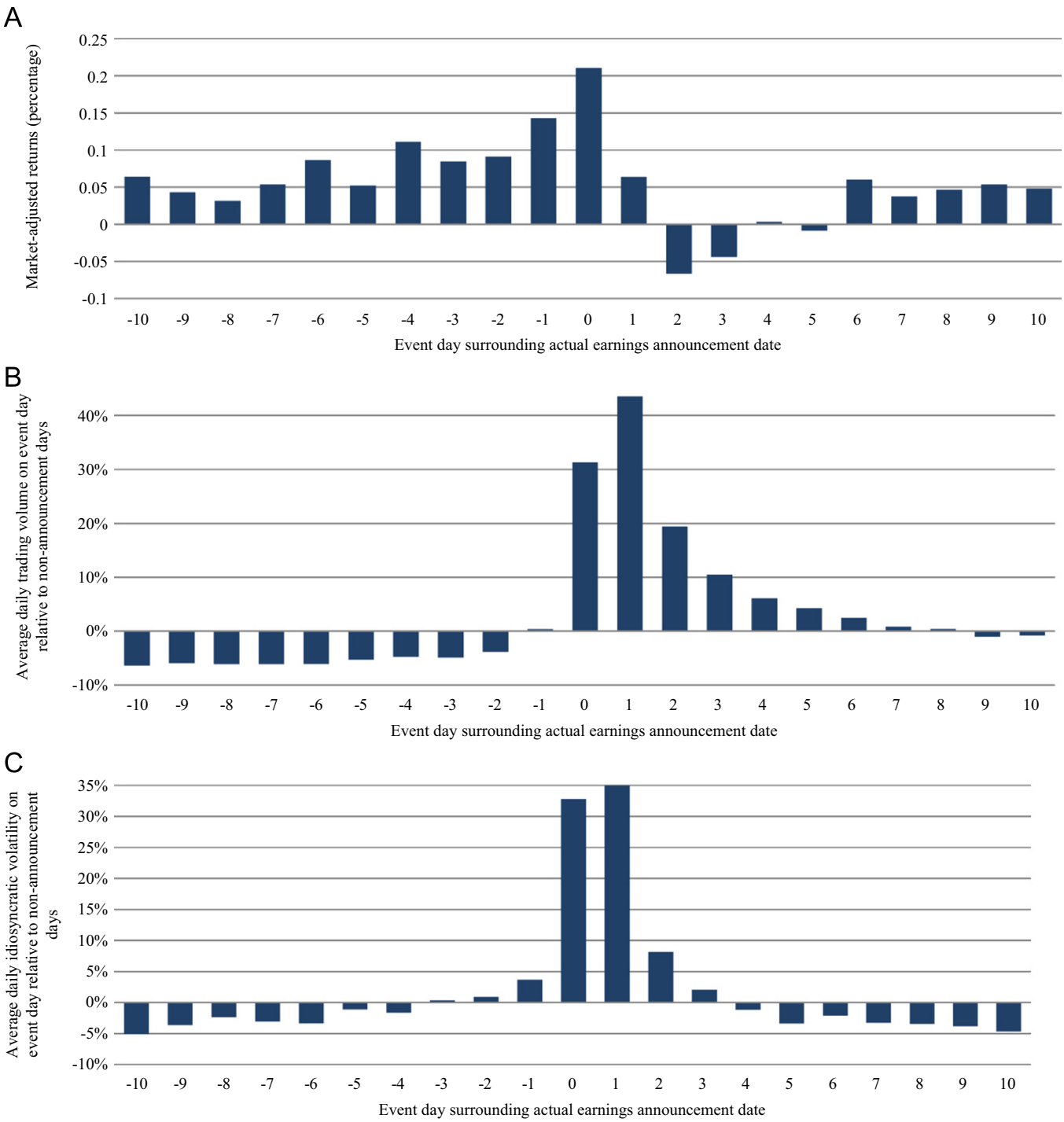

Fig. 3. Returns, volume, and idiosyncratic volatility surrounding earnings announcements. The following panels show the average daily market-adjusted returns, abnormal trading volume, and abnormal idiosyncratic volatility for the 21 trading days surrounding a firm's actual earnings announcement date, over the period January 1991 through December 2010. Event day 0 is the day of the earnings announcement. Panel A plots the average market-adjusted return for each event day *t*. For a given firm-announcement, the market-adjusted return on event day *t* is equal to the firm's raw return that day less the country-specific market return. All returns are denominated in U.S. dollars. The average over all observations is the average market-adjusted return on event day *t*. Panel B plots the average abnormal volume on event day *t*. For a given firm-announcement, abnormal volume on event day *t* is defined as the firm's volume that day divided by the average daily volume over the firm's non-announcement period, minus one. The non-announcement period is defined as 125 trading days before to 125 trading days after the earnings announcement, excluding the 21 days surrounding the announcement date. The average over all firm-announcements is the average abnormal volume on event day *t*. Panel C plots the average abnormal idiosyncratic volatility on event day *t*. For a given firm-announcement, abnormal idiosyncratic volatility on event day *t* is equal to the square root of the squared residual (from a firm-year market model estimated with three lags) on day *t*, divided by the square root of the average squared residual during the non-announcement period, minus one. The average over all firm-announcements is the average abnormal idiosyncratic volatility on event day *t*. (A) Daily market-adjusted returns surrounding earnings announcements, (B) Average daily trading volume surrounding earnings announcements (relative to non-announcement period) and (C) Average daily idiosyncratic volatility surrounding earnings announcements (relative to non-announcement period).

regression (1), replacing the expected announcement month by the actual month of the earnings release. In untabulated results we find that the coefficient on the new indicator variable, $ActAnn_{ijt}$, equals 113.4 basis points, which is insignificantly different from that on $ExpAnn_{ijt}$.

5. Potential explanations for the earnings announcement premium

5.1. Preliminaries

In this section we use the documented cross-country differences in the magnitude of the earnings announcement premium in order to provide insights into the premium's origins. To put our analyses into context, it is helpful to first examine how the premium varies on a daily basis in the days surrounding an earnings announcement.[22]

In Fig. 3, Panel A, we plot the average daily market-adjusted return over the period beginning ten trading days before the actual earnings announcement date and ending ten trading days after.[23] The most striking aspect of this graph is how much larger the average daily market-adjusted returns are over the pre-announcement period (event days −10 through −2, where event day 0 is the announcement date) than over the post-announcement period (event days 2 through 10). Pre-announcement returns average 6.88 basis points per day, while post-announcement returns average only 1.43 basis points. During the three-day announcement window (event days −1 through 1), the daily average is 13.92 basis points.

We also plot average daily abnormal volume over the same period (Fig. 3, Panel B).[24] For a given firm and earnings announcement, abnormal volume on any event day t is defined as the percentage by which that day's volume exceeds (or falls short of) average daily volume during the non-announcement period (defined as 125 trading days before to 125 trading days after the earnings announcement, excluding the 21 days surrounding the announcement date). Over our entire sample, average abnormal volume for event day t is the average of these firm- and announcement-specific percentages. In sharp contrast to the return pattern in Panel A, average daily abnormal volume is much lower during the pre-announcement period than it is after. Before the announcement it averages 5.5% below average daily non-announcement period volume, but is 4.7% above that level afterward. During the three-day announcement window, volume averages 25.1% above average daily non-announcement period volume.

In Panel C of Fig. 3 we plot average daily abnormal idiosyncratic volatility for each event day t. To calculate this measure, we first estimate a yearly market model for each firm i, where the dependent variable is the firm's daily return and the independent variable is its country's daily market return (with three lags of market return included so as to mitigate problems associated with stale prices of thinly traded stocks).[25] Denote by ε_{it}^2 the squared residual from this regression (that is, the square of the idiosyncratic return) for event day t. Denote by $\bar{\varepsilon}_{inon}^2$ the average of the daily squared residuals over the non-announcement period. The abnormal idiosyncratic volatility for firm i on event day t, $AIVOL_{it}$ is then given by:

$$AIVOL_{it} = \sqrt{\frac{\varepsilon_{it}^2}{\bar{\varepsilon}_{inon}^2}} - 1. \tag{2}$$

The abnormal idiosyncratic volatility for event day t, $AIVOL_t$, is the average of $AIVOL_{it}$ over all firms and all years. By subtracting one in expression (2), $AIVOL_t$ can be thought of as the percentage by which event day t idiosyncratic volatility exceeds (or falls short of) average volatility during the non-announcement period.

As seen in Panel C, idiosyncratic volatility during each day of the pre-announcement and post-announcement periods is less than during the non-announcement period (by an average of 2.1% and 1.3%, respectively). There is a big spike in idiosyncratic volatility, however, over the three-day announcement window, where it averages 23.8% more than during the non-announcement period.[26]

5.2. Information uncertainty as an explanation for the earnings announcement premium

The spike in idiosyncratic volatility during the announcement window is not surprising, given the amount of firm-specific information that is disclosed through earnings. It suggests that uncertainty over the nature of the information to be revealed may be driving the higher returns during the earnings announcement month. Merton (1987) and Malkiel and Xu (2002) show analytically that higher idiosyncratic share price volatility leads to higher expected returns as long as investors are constrained in their security holdings, whether as a result of incomplete information, transactions costs, or institutional restrictions on holdings. Consistent with their model, Malkiel and Xu (2002) find empirically that idiosyncratic risk is an important determinant of the

[22] In order to ameliorate the effect of daily return data errors, we set to missing (1) any daily return greater than 200%, (2) two consecutive daily returns, r_d and r_{d+1}, if $(1+r_d)(1+r_{d+1})-1<0.20$and either r_d or r_{d+1} is greater than 100% (this eliminates daily returns characterized by a sharp increase, followed by a sharp reversal), or (3) any daily return of zero for which there is no valid volume data. [These are also the filters employed by Griffin, Kelly, and Nardari (2010).] If this procedure results in the removal of any daily return during the 21 days surrounding an earnings announcement date, we drop that firm-announcement from the analysis in this section. After imposing these requirements, our daily sample consists of approximately 102,000 annual announcements over 43 countries. (We lose Luxembourg, Jordan, and Venezuela because of the lack of daily country index returns.)

[23] For a given firm-announcement, the market-adjusted return on event day t is equal to the firm's raw return that day less the country-specific market return. The average over all observations is the average market-adjusted return on event day t.

[24] Volume data come from Bloomberg, when available; otherwise, Datastream is our source for volume.

[25] For analyses involving abnormal idiosyncratic volatility, we include only those firm-years where no more than 25% of the days have returns of zero (to ensure that our firm-specific market model estimations are reliable).

[26] Ball and Kothari (1991) and Cohen, Dey, Lys, and Sunder (2007) also find an increase in idiosyncratic volatility around earnings announcements.

cross-sectional differences in expected returns. While Ang, Hodrick, Xing, and Zhang (2006) challenge this finding and show a negative relation between idiosyncratic volatility and expected returns, Fu (2009) argues that their result stems from the use of an inappropriate proxy for investors' expectation of idiosyncratic volatility. Employing a more precise expectation model, Fu (2009) again finds a positive relation between expected returns and idiosyncratic volatility. In our setting, these results imply that investors will demand a higher expected return prior to an earnings announcement in anticipation of a sharp increase in idiosyncratic volatility when earnings are disclosed. This is consistent with the return pattern in Panel A of Fig. 3.

To formally test this hypothesis, we exploit the documented differences in the earnings announcement premium across countries; the countries where firms have a greater increase in idiosyncratic volatility at the time of the earnings announcement should exhibit a higher premium. As a first step in our analysis we calculate, for each firm and calendar year, the level of firm-specific abnormal idiosyncratic volatility during the firm's three-day announcement window. As before, we estimate a simple market model (again, with three lags of market return included) and extract each trading day's regression residual. The firm-specific abnormal idiosyncratic volatility around the firm's earnings announcement is calculated by dividing the standard deviation of the residuals over the three-day announcement window by the standard deviation of the residuals over the non-announcement period. We then subtract one from this ratio, allowing us to interpret the resulting number as the percent by which announcement-window idiosyncratic volatility exceeds (or falls short of) that during the non-announcement period. A country's abnormal idiosyncratic volatility is the average of these firm-specific percentages across all the firms in the country and across all the years for which we have data.

We rank countries according to their level of abnormal idiosyncratic volatility and partition them into three groups—high, medium, and low. We then estimate the average earnings announcement premium for each group using regression (1).[27] We replace the indicator variable for the expected announcement month, $ExpAnn_{ijt}$, in expression (1) with that for the actual announcement month, $ActAnn_{ijt}$, since for this part of the analysis we are not focused on implementable strategies. The results of our analysis are reported in Table 7, Panel A.[28] Consistent with heightened idiosyncratic volatility being a driver of the earnings announcement premium, the countries with high abnormal idiosyncratic volatility have the largest average earnings announcement premium, 156.9 basis points (18.83% on an annual basis). The average announcement premium for the medium abnormal idiosyncratic volatility group is smaller, 91.8 basis points, but still significant. The average announcement premium for the low abnormal idiosyncratic volatility countries is insignificant. The difference between the premium for the high and low groups is a significant 171.9 basis points (an economically large 20.63% annually).

In order to ensure that these results are not driven by an endogeneity problem between contemporaneous idiosyncratic volatility and return, we re-estimate our regressions using a country's CIFAR financial disclosure score (Center for International Financial Analysis & Research, 1995) as an instrument for abnormal idiosyncratic volatility. This score measures the extent to which 85 different financial items are disclosed in the 1995 annual reports of the firms in that country; the higher the score, the more items are disclosed. It has been used in many studies to measure disclosure quality and/or transparency, which underlie the increase in idiosyncratic volatility around earnings announcements.[29] In untabulated results we show that the score has a positive correlation of 67% with the level of abnormal idiosyncratic volatility during the earnings announcement window, providing support for its use as an instrumental variable. An additional justification for using the CIFAR score as an instrumental variable is that its compilation precedes in time the majority of our earnings announcements.

We replicate our prior analysis, partitioning countries into high, medium, and low groups, according to their CIFAR score.[30] As reported in Table 7, Panel B, the high disclosure countries have the greatest average earnings announcement premium, a significant and economically large 182.4 basis points, or 21.89% annually. The average premium for the medium disclosure countries is a smaller, but still significant, 44.1 basis points. The average premium for the low disclosure countries is insignificantly different from zero. These results are qualitatively very similar to those obtained when partitioning on abnormal idiosyncratic volatility and provide additional support for the hypothesis that heightened idiosyncratic volatility around earnings announcements is a driver of the observed earnings announcement premium.[31]

[27] The average abnormal idiosyncratic volatility of countries in the high group is 33%. Among the countries in this group are France, Hong Kong, Japan, and the U.K. The average abnormal idiosyncratic volatility of countries in the low group is −4%. Among the countries in this group are Greece, South Korea, and Taiwan. While not in our sample, the U.S. is a high abnormal idiosyncratic volatility country.

[28] Of the 46 countries in our sample, we exclude three—Jordan, Luxembourg, and Venezuela—because the firms in those countries have too few observations to be able to reliably estimate the market model for them.

[29] See, for example, La Porta, Lopez-de-Silanes, and Shleifer (1998), Bushman, Piotroski, and Smith (2004), DeFond, Hung, and Trezevant (2007), and Leuz (2010).

[30] Among the countries with high CIFAR scores are Australia, the U.K., and France. Among the countries with low CIFAR scores are Germany, Greece, and Italy. While not in our sample, the U.S. has a medium CIFAR score. The Czech Republic, China, Hungary, Indonesia, Jordan, Luxembourg, Peru, Poland, Portugal, and Venezuela are not included in our analysis because those countries do not have a CIFAR score.

[31] We also test for a relation between the announcement premium and idiosyncratic volatility at the firm level. In this case we use as a proxy for contemporaneous idiosyncratic volatility an average of the firm's idiosyncratic volatility over the prior four years. (If fewer than four years are available, all prior years are used.) We rank our firm-year observations according to the value of this proxy and then partition the observations into three groups—low, medium, and high. For each group we estimate regression (1). In untabulated results, we find that the announcement premium for the low group, 0.41, is insignificantly different from zero, while for the high group it is a significant 0.97. The difference, 0.56, is significantly greater than zero.

Table 7
Earnings announcement premium and abnormal idiosyncratic volatility.

This table reports the average coefficients for monthly regressions of:

$$\mathrm{Ret}_{ijt} = \alpha + \beta_1 ActAnn_{ijt} + \beta_2 Mom_{ijt} + \beta_3 MktCap_{ijt} + \beta_4 BTM_{ijt} + \sum_j \gamma_j Country_j + \varepsilon_{ijt},$$

where Ret_{ijt} is the natural log of one plus the raw return during month t for firm i in country j, denominated in U.S. dollars; $ActAnn_{ijt}$ is an indicator variable equal to one if firm i of country j announces annual earnings in month t and is equal to zero, otherwise; Mom_{ijt} is the natural log of one plus the raw return for firm i of country j over months $t-1$ through $t-11$, denominated in U.S. dollars; $MktCap_{ijt}$ is the log of the market capitalization of firm i in country j at the end of month $t-1$, denominated in U.S. dollars; BTM_{ijt} is the log of the book-to-market ratio for firm i of country j as of the end of the prior fiscal year; $Country_j$ is an indicator variable equal to one for all firms in country j and equal to zero, otherwise; ε_{ijt} is the regression residual for firm i of country j in month t. The coefficients on the country indicator variables are not reported in the table. The average number of observations in the monthly regressions as well as the regression R^2 are also reported in the table. In Panel A, we report Fama-MacBeth regression estimates for high, medium, and low average abnormal idiosyncratic volatility countries (for each country, the average is taken over all firm-announcements within that country). For each firm and earnings announcement, abnormal idiosyncratic volatility is computed as the standard deviation of the residuals (generated from yearly market models with three lags) over the three-day window surrounding the announcement, divided by the standard deviation of the residuals over the 250 days surrounding the announcement (excluding the announcement window), minus one. In Panel B, we report Fama-MacBeth regression estimates for high, medium, and low CIFAR (Center for International Financial Analysis & Reporting) accounting disclosure score countries (as reported in Bushman, Piotroski, and Smith, 2004). This score measures the extent to which 85 different financial items are disclosed in the 1995 annual reports of the firms in that country; the higher the score, the more items are disclosed. Of the 46 foreign countries originally included in our data, we omit Jordan, Luxembourg, and Venezuela from the current analysis due to lack of sufficient data to reliably estimate market models within these countries. In addition, we also omit China, Czech Republic, Hungary, Poland, Peru, and Indonesia due to missing CIFAR data. These regressions include only those months for which there are at least five announcer and five non-announcer firms within each group during the month. Below each coefficient value is the corresponding t-statistic. Significance is determined based on the time-series distribution of equally weighted monthly coefficient estimates. ***=Significant at the 1% level; **=significant at the 5% level; *=significant at the 10% level. See Table 1 for a detailed description of our sample.

Panel A: High, medium, and low idiosyncratic volatility countries

Country	# Of months	Avg. # of monthly obs.	Average coefficient estimates					R-squared (%)
			Intercept	*ActAnn*	*Mom*	*MktCap*	*BTM*	
High abnormal idiosyncratic volatility countries (14)	236	4,124	−0.872	**1.569*****	**0.007*****	**0.072***	**0.468*****	14.7
			−1.49	*13.79*	*3.08*	*1.76*	*7.01*	
Medium abnormal idiosyncratic volatility countries (15)	236	2,682	−0.363	**0.918*****	**0.010*****	0.005	**0.472*****	17.8
			−0.52	*3.92*	*4.28*	*0.19*	*4.60*	
Low abnormal idiosyncratic volatility countries (14)	236	1,764	−0.657	−0.150	0.004	**0.138***	**0.590*****	20.1
			−0.49	*−0.45*	*0.65*	*1.69*	*6.32*	
Test of difference (High−Low)=1.569 −(−0.150)=**1.719***** (Z=4.09; $p<0.05$)								

Panel B: High, medium, and low CIFAR countries

Country	# Of months	Avg # of monthly obs.	Average coefficient estimates					R-squared (%)
			Intercept	*ActAnn*	*Mom*	*MktCap*	*BTM*	
High CIFAR score countries (12)	228	2,744	−**1.208****	**1.824*****	**0.010*****	**0.098****	**0.491*****	12.7
			−2.02	*15.61*	*4.61*	*2.31*	*7.61*	
Medium CIFAR score countries (13)	228	3,848	−0.529	**0.441*****	0.003	**0.080***	**0.535*****	15.9
			−0.77	*3.06*	*1.27*	*1.78*	*7.30*	
Low CIFAR score countries (11)	228	1,490	−4.185	−0.266	−0.014	0.339	0.279	18.6
			−1.26	*−0.54*	*−0.60*	*1.55*	*1.52*	
Test of difference (High−Low)=1.824−(−0.266)=**2.090**** (Z=2.41; $p<0.05$)								

Table 8
Fama-MacBeth regression analysis of interim earnings announcement premium.
Panel A reports the average coefficients for monthly regressions of:

$$\mathrm{Ret}_{ijt} = \alpha + \beta_1 ExpAnn_{ijt} + \beta_2 Mom_{ijt} + \beta_3 MktCap_{ijt} + \beta_4 BTM_{ijt} + \beta_5 ExpInt_{ijt} + \sum_{j=1}^{46} \gamma_j Country_j + \varepsilon_{ijt},$$

where Ret_{ijt} is the natural log of the raw return during month t for firm i in country j, denominated in U.S. dollars; $ExpAnn_{ijt}$ is an indicator variable equal to one if firm i of country j is expected to announce annual earnings in month t and is equal to zero, otherwise; Mom_{ijt} is the natural log of the average raw return for firm i of country j over months $t-1$ through $t-11$, denominated in U.S. dollars; $MktCap_{ijt}$ is the log of the market capitalization of firm i in country j at the end of month $t-1$, denominated in U.S. dollars; BTM_{ijt} is the log of the book-to-market ratio for firm i of country j as of the end of the prior fiscal year; $ExpInt_{ijt}$ is an indicator variable equal to one if firm i of country j is expected to announce interim earnings in month t and is equal to zero, otherwise; $Country_j$ is an indicator variable equal to one for all firms in country j and equal to zero, otherwise; ε_{ijt} is the regression residual for firm i of country j in month t. The coefficients on the country indicator variables are not reported. The average number of observations in the monthly regressions as well as the regression R^2 are also reported in the table. In this regression we only include months for which there are at least ten announcer and ten non-announcer firms. Below each coefficient value is the corresponding t-statistic. We determine significance based on the time-series distribution of equally weighted monthly coefficient estimates. In Panel B we report the average abnormal idiosyncratic volatility (across all firms) around annual and, separately, around interim announcements, both for the world (excluding the U.S.) and for the U.S. For each firm and earnings announcement, abnormal idiosyncratic volatility is computed as the standard deviation of the residuals (generated from yearly market models with three lags) over the three-day window surrounding the announcement, divided by the standard deviation of the residuals over the 250 days surrounding the announcement (excluding the announcement window), minus one. The difference between the average abnormal idiosyncratic volatility for annual and interim announcements is also reported, along with the corresponding t-statistic. ***=Significant at the 1% level; **=significant at the 5% level; *=significant at the 10% level; ###=significantly different from 1 at the 1% level; ##=significantly different from 1 at the 5% level; #=significantly different from 1 at the 10% level. See Table 1 for a detailed description of our sample.

Panel A: Controlling for interim announcement months

Avg. # of monthly obs.	Intercept	*ExpAnn*	*Mom*	*MktCap*	*BTM*	*ExpInt*	*R*-squared (%)
8996	−1.312	**0.943*****	**0.008*****	**0.080****	**0.511*****	0.164	18.48
	−1.31	*11.15*	*4.23*	*2.25*	*9.22*	*1.17*	

Panel B: Average abnormal idiosyncratic volatility around annual and interim earnings announcements, for the world (excluding the U.S.) and for the U.S.

	World (excluding U.S.)			U.S.		
	Annual	*Interim*	*Difference*	*Annual*	*Interim*	*Difference*
Average abnormal idiosyncratic volatility	**1.14**###	**1.03**#	**0.110*****	**1.26**###	**1.25**###	**0.010***

Interim earnings announcements provide another means by which to test this hypothesis. Frazzini and Lamont (2007) find a positive and significant premium for interim disclosures for U.S. firms. To estimate the interim announcement premium internationally, we add to regression (1) an indicator variable, $ExpInt_{ijt}$, which is equal to one if firm i in country j is expected to announce interim earnings (semiannual or quarterly) during month t, and is equal to zero, otherwise.[32] The coefficient on this variable can be interpreted as the average incremental monthly return during expected interim announcement months relative to non-announcement months. As reported in Table 8, Panel A, the coefficient on $ExpInt_{ijt}$ is an insignificant 16.4 basis points. In contrast, the coefficient on $ExpAnn_{ijt}$ in this regression is a significantly positive 94.3 basis points. In contrast to U.S. firms, there is no evidence that an interim announcement premium exists in foreign countries.

If heightened idiosyncratic volatility around earnings announcements is a driver of the premium, then these results would imply that volatility should be significantly lower around interim announcements than around annual announcements for our international sample, but not necessarily for the U.S. We estimate abnormal idiosyncratic volatility around interim announcements in a manner similar to that for annual announcements; the relevant results are presented in Table 8, Panel B. For our international sample, idiosyncratic volatility during the annual announcement window averages a significant 14% above its level during the non-announcement period. For interim announcements, it is less than one-quarter as large—a marginally significant 3% above that during the non-announcement period. The difference of 11 percentage points is reliably positive. For the U.S. sample, in contrast, average idiosyncratic volatility is significantly greater than zero for both the annual and interim announcements, and is of similar magnitude (26% and 25%, respectively). This difference, although statistically significant, is economically negligible. Taken together, these results provide additional support for the hypothesis that idiosyncratic volatility is a driver of the earnings announcement premium.

5.3. Investor attention

Frazzini and Lamont (2007) conjecture that investor attention is partly responsible for the existence of the announcement premium. Under the attention hypothesis, individual investors purchase stocks that grab their attention, thereby temporarily increasing share prices. If heightened

[32] Similar to the determination of the expected annual announcement month, the expected interim announcement month is set equal to the actual interim announcement month during the prior year.

Table 9
Earnings announcement premium and abnormal trading volume.

The table reports the average coefficients for monthly regressions of:

$$Ret_{ijt} = \alpha + \beta_1 ActAnn_{ijt} + \beta_2 Mom_{ijt} + \beta_3 MktCap_{ijt} + \beta_4 BTM_{ijt} + \sum_j \gamma_j Country_j + \varepsilon_{ijt},$$

where Ret_{ijt} is the natural log of one plus the raw return during month t for firm i in country j, denominated in U.S. dollars; $ActAnn_{ijt}$ is an indicator variable equal to one if firm i of country j announces annual earnings in month t and is equal to zero, otherwise; Mom_{ijt} is the natural log of one plus the raw return for firm i of country j over months $t-1$ through $t-11$, denominated in U.S. dollars; $MktCap_{ijt}$ is the log of the market capitalization of firm i in country j at the end of month $t-1$, denominated in U.S. dollars; BTM_{ijt} is the log of the book-to-market ratio for firm i of country j as of the end of the prior fiscal year; $Country_j$ is an indicator variable equal to one for all firms in country j and equal to zero, otherwise; ε_{ijt} is the regression residual for firm i of country j in month t. The coefficients on the country indicator variables are not reported in the table. For a given firm-announcement, abnormal trading volume for the pre-announcement period (event days −10 through −2) or the post-announcement period (event days 2 through 10) is defined as the firm's average daily volume over the period divided by the average daily volume over the firm's non-announcement period, minus one. The non-announcement period is defined as 125 trading days before to 125 trading days after the earnings announcement, excluding the 21 days surrounding the announcement date. We average across firms and announcements in each country to produce a country-level measure of abnormal daily volume for the period, winsorizing firm-announcement abnormal volume observations within each country to reduce the effect of outliers. Panel A (Panel B) reports Fama-MacBeth regression results across subsamples of countries ranked on pre-announcement (post-announcement) period abnormal volume. Also reported are the average number of observations in the monthly regressions as well as the regression R^2. Below each coefficient value is the corresponding t-statistic. Significance is determined based on the time-series distribution of equally weighted monthly coefficient estimates. ***=Significant at the 1% level; **=significant at the 5% level; *=significant at the 10% level. See Table 1 for a detailed description of our sample.

Panel A: Fama-MacBeth regressions across country-level partitions based on pre-announcement period abnormal volume

Abnormal volume partition	Avg. # of monthly obs.	Intercept	*ActAnn*	*Mom*	*MktCap*	*BTM*	*R*-squared (%)
High (14 countries)	2,004	**−1.812***	0.332	**0.007****	**0.146****	**0.543*****	20.3
		−1.71	*1.41*	*2.25*	*1.99*	*5.82*	
Medium (14 countries)	3,842	−0.586	**0.500*****	0.003	**0.080***	**0.534*****	16.5
		−0.58	*3.61*	*0.87*	*1.76*	*7.46*	
Low (14 countries)	2,296	−0.736	**1.734*****	**0.011*****	0.054	**0.399*****	14.2
		−1.22	*15.39*	*5.65*	*1.26*	*5.56*	
Test of difference (High−Low)=0.332−1.734=**−1.402***** ($Z=-5.25$; $p<0.01$)							

Panel B: Fama-MacBeth regressions across country-level partitions based on post-announcement period abnormal volume

Abnormal volume partition	Avg. # of monthly obs.	Intercept	*ActAnn*	*Mom*	*MktCap*	*BTM*	*R*-squared (%)
High (14 countries)	2,794	−0.850	**1.667*****	**0.014*****	0.062	**0.455*****	15.8
		−1.38	*14.69*	*6.78*	*1.29*	*5.58*	
Medium (14 countries)	3,546	−0.989	**0.839*****	0.003	**0.097****	**0.542*****	17.8
		−1.40	*4.78*	*0.49*	*2.55*	*6.94*	
Low (14 countries)	1,796	0.372	**0.994*****	**0.011*****	0.025	**0.555*****	18.5
		0.35	*3.65*	*4.51*	*0.28*	*6.04*	
Test of difference (High−Low)=1.667−0.994=**0.673**** ($Z=2.24$; $p<0.05$)							

investor attention to earnings announcements is a driver of the premium, then we should find a positive relation between volume and returns around earnings announcements.

To test this hypothesis, we compute the pre-announcement period abnormal volume for each firm and earnings announcement, measured by the average daily volume during the pre-announcement period as a percent of the average daily non-announcement period volume. We focus on the pre-announcement period because the bulk of the market-adjusted return is realized during this time (recall Fig. 3, Panel A). We rank each country according to the average abnormal volume over all firm-announcements for that country, and then partition the countries into high, medium, and low abnormal volume groups.[33] The attention-grabbing hypothesis would predict that the premium will be largest (smallest) in the high (low) group.

We rerun regression (1) for each group, again replacing $ExpAnn_{ijt}$ with $ActAnn_{ijt}$.[34] Results of this regression analysis are reported in Table 9, Panel A. The volume-return relation is opposite to what the investor attention hypothesis predicts: the average earnings announcement premium is largest for the low pre-announcement volume group and smallest for the high pre-announcement volume group. The difference between the two, 140.2 basis points, is economically large and significantly greater than zero.

To gain insight into what might be causing these inconsistent results, we repeat our analysis, this time ranking countries according to average post-announcement period

[33] Among the countries in the high abnormal volume group are China, Germany, Italy, and Spain. Among those in the low abnormal volume group are Australia, France, Greece, and the U.K.

[34] For each regression we only include months for which there are at least five announcing and five non-announcing firms. We also delete observations for which volume is missing on one or more days during the period.

abnormal volume.[35] Since the market-adjusted return is negligible during this period (Fig. 3, Panel A), we would not expect to find any association between a country's ranking and its announcement premium. Results are presented in Table 9, Panel B. In contrast to our previous findings, in this period we find that the high volume group is associated with the largest average premium. The difference between the premiums for the high and low volume groups, 67.3 basis points, is significantly greater than zero.

While these results are not supportive of the investor attention hypothesis, they are consistent with the documented positive association between idiosyncratic volatility and the announcement premium. To see this, we refer to the theoretical model of George, Kaul, and Nimalendran (1994), who analyze trading volume in a setting where informed traders take positions ahead of a forthcoming disclosure and discretionary liquidity traders decide whether to postpone their trading until after that disclosure. George, Kaul, and Nimalendran (1994) show that if liquidity traders are sufficiently sensitive to the adverse selection costs of trading against informed traders, total pre-announcement volume will be suppressed relative to total post-announcement volume. Building upon this, Chae (2005) predicts that the drop (increase) in volume before (after) an earnings announcement will be sharper the higher the idiosyncratic volatility at the time of the announcement.[36] His empirical results support this prediction. Given the positive association between idiosyncratic volatility and the announcement premium, this implies that the level of abnormal volume during the pre-announcement (post-announcement) period should be negatively (positively) related to the premium, which is what we find.

5.4. Seasonality pattern in stock returns

In recent work, Heston and Sadka (2008, 2010) document a seasonality pattern in stock returns. They find that outperforming (underperforming) stocks in a given month tend to outperform (underperform) during the same month in subsequent years. Given that firms tend to announce annual earnings in the same calendar month each year, the earnings announcement premium could potentially be a manifestation of return seasonality. To examine this possibility, we control for seasonality and re-estimate the premium. We do so by modifying regression (1) as follows:

$$\begin{aligned} Ret_{ijt} = {} & \alpha + \beta_{1t} ExpAnn_{ijt} + \beta_{2t} Mom_{ijt} + \beta_{3t} MktCap_{ijt} \\ & + \beta_{4t} BTM_{ijt} + \beta_{5t} PriorRet_{ijt} + \beta_{6t} ExpAnn_{ijt} \cdot PriorRet_{ijt} \\ & + \sum_{j=1}^{45} \gamma_j Country_j + \varepsilon_{ijt}, \end{aligned} \tag{3}$$

where $PriorRet_{ijt}$ is the raw stock return of firm i in country j during month $t-12$. In addition to $PriorRet_{ijt}$, we include in regression (3) the interaction between the prior year's return and the expected announcement month indicator variable. This interaction term tests whether the impact of the seasonality effect is greater during earnings announcement months.

Regression results are reported in Table 10, Panel A. Consistent with Heston and Sadka (2008, 2010), the average coefficient on the prior return variable is positive and significant; the interaction term, though, is not. Importantly for our analysis, the average coefficient on $ExpAnn_{ijt}$ remains significantly positive and economically large. In fact, its value, 0.89, is insignificantly different from the average coefficient of 0.955 originally estimated in regression (1). We conclude from this that the return premium during earnings announcement months is not simply a manifestation of the seasonality pattern documented by Heston and Sadka (2008, 2010).

5.5. Unexpectedly strong earnings during sample period

Another potential explanation for our results is that, globally, firms had unexpectedly strong earnings during our sample period and that this is reflected in positive market returns concentrated in earnings announcement months. To test this, we estimate the following regression:

$$LS_j = \alpha + \gamma AggGrowth_j + \varepsilon_j. \tag{4}$$

The dependent variable, LS_j, is the average monthly return over our sample period on a long-short portfolio (long in the shares of firms in country j during their actual earnings announcement months and short in the shares of those firms during their non-announcement months). The independent variable, $AggGrowth_j$, is the average annual aggregate earnings growth for country j over our sample period. Annual aggregate earnings growth for each calendar year t, $AggGrowth_{jt}$, is equal to the aggregate difference between the annual earnings announced during years t and year $t-1$ for all firms in the country, divided by the sum of the market values of these firms as of the end of calendar $t-1$.[37]

If unexpectedly good earnings news is driving the observed global earnings announcement premium, then we should observe a significantly positive relation between a country's average annual aggregate earnings growth and the size of its announcement premium (that is, γ should be significantly greater than zero). This is not what we find. As reported in Table 10, Panel B, γ, while positive, is insignificantly different from zero. As robustness checks, we re-estimate this regression, first using the average annual growth of a country's gross domestic product (GDP) as the independent variable and then using the country's average annual market return. In untabulated results, we find that in

[35] For the post-announcement period, the high abnormal volume group includes China, France, Hong Kong, and the U.K. Included in the low abnormal volume group are Australia, Germany, Greece, and South Korea.

[36] This follows under the reasonable assumption that idiosyncratic volatility at the time of the earnings announcement and the adverse selection costs of trading are both positively related to the amount of information held by privately informed traders.

[37] We exclude from our long-short return computations those months without a valid long-short portfolio (that is, in which there are fewer than five announcer firms and five non-announcer firms in the country). There is one country without a valid long-short portfolio in any month of our sample period; consequently, it is not included in this analysis.

Table 10
Earnings announcement premium, stock return seasonality, and strong unexpected earnings.

Panel A reports the average coefficients for monthly regressions of:

$$Ret_{ijt} = \alpha + \beta_1 ExpAnn_{ijt} + \beta_2 Mom_{ijt} + \beta_3 MktCap_{ijt} + \beta_4 BTM_{ijt} + + \beta_5 PriorRet_{ijt} + \beta_6 ExpAnn_{ijt} \cdot PriorRet_{ijt} + \sum_{j=1}^{45} \gamma_j Country_j + \varepsilon_{ijt},$$

where Ret_{ijt} is one plus the natural log of the raw return during month t for firm i in country j, denominated in U.S. dollars; $ExpAnn_{ijt}$ is an indicator variable equal to one if firm i of country j is expected to announce annual earnings in month t and is equal to zero, otherwise; Mom_{ijt} is the natural log of one plus the raw return for firm i of country j over months $t-1$ through $t-11$, denominated in U.S. dollars; $MktCap_{ijt}$ is the log of the market capitalization of firm i in country j at the end of month $t-1$, denominated in U.S. dollars; BTM_{ijt} is the log of the book-to-market ratio for firm i of country j as of the end of the prior fiscal year; $PriorRet_{ijt}$ is the raw stock return of firm i in country j during month $t-12$, denominated in U.S. dollars; $Country_j$ is an indicator variable equal to one for all firms in country j and equal to zero, otherwise; ε_{ijt} is the regression residual for firm i of country j in month t. The coefficients on the country indicator variables are not reported in the table. Panel B reports results from the country-level regression: $LS_j = \alpha + \gamma AggGrowth_j + \varepsilon_j$. The dependent variable, LS_j, is the average monthly return during our sample period on a portfolio long in the shares of firms in country j during their actual earnings announcement month and short in the shares of those firms during their non-announcement months, denominated in U.S. dollars. The independent variable, $AggGrowth_j$, is the average annual aggregate earnings growth for country j over our sample period, where aggregate earnings growth for each calendar year t is computed as the aggregate difference between the annual earnings announced during years t and $t-1$ for all firms in country j, divided by the sum of the market values of these firms as of the end of calendar year $t-1$, all denominated in U.S. dollars. Panel C reports the results of replicating the Fama-MacBeth regression results (Table 5) for two subsamples: one comprised of country-years with positive annual aggregate earnings growth and the other comprised of country-years with negative annual aggregate earnings growth. Included in the long-short return calculations are only those months with at least five announcer and five non-announcer firms in country j. The average market return for country j is computed only for those months as well. The sample period covers 1991–2010. The average number of observations in the monthly regressions is also reported in Panels A and C, with Panel B reporting the number of countries in the regression. The regression R^2 is also reported. Below each coefficient value is the corresponding t-statistic. Significance in Panels A and C is based on the time-series distribution of equally weighted monthly coefficient estimates. ***=Significant at the 1% level; **=significant at the 5% level; *=significant at the 10% level. See Table 1 for a detailed description of our sample.

Panel A: Testing for seasonality in returns

Avg. # of monthly obs.	Intercept	*ExpAnn*	*Mom*	*MktCap*	*BTM*	*PriorRet*	*ExpAnn* × *PriorRet*	*R*-squared (%)
8,919	**−1.655****	**0.890*****	**0.008*****	**0.072****	**0.513*****	**0.028*****	0.011	18.64
	−2.46	*9.84*	*4.38*	*2.05*	*9.45*	*5.53*	*1.49*	

Panel B: Testing for unexpectedly strong earnings driving returns using average annual aggregate earnings growth

Number of countries	Intercept	*AggGrowth*	*R*-squared (%)
44	0.008	0.193	3.1
	0.45	*1.09*	

Panel C: Testing for unexpectedly strong earnings driving returns using the subsamples of positive and negative aggregate earnings growth months

Growth period	Avg. # of monthly obs.	Intercept	*ExpAnn*	*Mom*	*MktCap*	*BTM*	*R*-squared (%)
Positive (228 months)	5,990	**−0.496**	**1.017*****	**0.009*****	**0.118*****	**0.541*****	16.76
		−0.49	*8.61*	*5.31*	*2.92*	*8.96*	
Negative (228 months)	3,090	**−1.821*****	**0.923*****	**0.005***	0.066	**0.529*****	16.86
		−3.64	*4.04*	*1.73*	*0.93*	*6.39*	
Test of difference (Positive − Negative) = 1.017 − 0.923 = 0.094 (Z = 0.33; not significant)							

both cases, the regression coefficient is again insignificantly different from zero.[38]

As a second test of this hypothesis, we partition our country-years into those with positive aggregate earnings growth and those with negative aggregate earnings growth and re-estimate regression (1) for each subsample. If unexpectedly strong earnings is a driver of the announcement premium, then we would expect to see a greater premium for the positive aggregate earnings growth subsample than for the negative growth subsample. As seen in Table 10, Panel C, there is no reliable evidence that this is the case. The difference between the premium for the positive aggregate earnings growth subsample (101.7 basis points) and the negative growth subsample (92.3 basis points) is an insignificant 9 basis points. We also find (in untabulated results) that the average return on the long-short portfolio in country-years with positive aggregate earnings growth is not significantly different from the average return in country-years with negative growth. None of these test results provides support for the conjecture that surprisingly favorable earnings news over our sample period is a driver of the earnings announcement premium.

[38] We separately calculate the correlation between the average monthly long-short portfolio return for each country-year and the annual aggregate earnings growth for that country-year. In untabulated results, we find that the correlation is an insignificant 0.065. We replicate this analysis using annual GDP growth instead of annual aggregate earnings growth and again find a correlation (−0.056) that is insignificantly different from zero.

5.6. Earnings-related systematic risk

Building on the work of Campbell (1993) and Campbell and Vuolteenaho (2004), Savor and Wilson (2011) develop a model in which it is assumed that a firm's earnings news is comprised of systematic (market-wide) and idiosyncratic components. In their model, investors use the information in each earnings announcement to revise their expectations for the systematic component of the earnings of all non-announcing firms. (They can do so only imperfectly because they do not observe the two components separately.) The key insight of their model is that, under reasonable parameterizations, investors should be compensated for bearing earnings-related systematic risk and that firms with higher risk levels should yield larger earnings announcement premia. Using U.S. data, they find evidence consistent with their conjecture.

To test whether earnings announcement risk provides a partial explanation for the earnings announcement premium in an international setting, we focus on the primary implication of the Savor and Wilson (2011) analysis: that the magnitude of the earnings announcement premium should be a predictor of aggregate earnings growth. As in Savor and Wilson (2011), we measure the announcement premium by the return on a long-short portfolio (long in announcing firms' shares and short in the shares of non-announcers), and then estimate the following regression:

$$AggGrowth_{jt+1} = \alpha + \beta_1 LS_{jt} + \beta_2 Mkt_{jt} + \beta_3 AggEarn_{jt} + \sum_{j=1}^{46} \gamma_j Country_j + \varepsilon_{jt}, \quad (5)$$

where:

$AggEarn_{jt}$ the aggregate annual earnings announced in calendar year t for all firms in country j, divided by the sum of the market values of these firms as of the end of calendar year t;

LS_{jt} the long-short portfolio return for year t in country j, calculated by summing the monthly long-short premia during the year, with each month's return weighted by the number of announcements that month;

Mkt_{jt} value-weighted market return on Datastream's total market index for country j in year t.

All other variables are as previously defined.

According to Savor and Wilson (2011), we should expect a positive relation between aggregate earnings growth and the long-short portfolio return; that is, the coefficient, β_1, should be positive. However, as reported in the first row of Table 11, it is insignificantly different from zero. As a robustness check, we alternatively compute the yearly long-short return by compounding the monthly returns and then rerun regression (5). As seen in the second row of the table, the coefficient on the compounded return is again insignificantly different from zero.

Savor and Wilson (2011) also posit that the magnitude of the premium should be persistent over time. This follows from their model as long as firms differ in their exposure to earnings announcement risk and the differential exposure is reasonably stable over time. They find evidence of this for U.S. firms. To test for persistence globally, we rank firms in our sample each year according to the size of the firm's annual earnings announcement premium over the prior two years, calculated as the difference between its average return during earnings announcement months and during non-announcement months. We then partition the firms into five portfolios, according to the size of their two-year prior average announcement premium. In untabulated results, we document a significant difference between the *current* earnings announcement premium for the quintile of firms with the largest *prior* premium and that for the quintile with the smallest *prior* premium. While this finding is consistent with earnings announcement risk being a driver of the announcement premium, it is also consistent with the abnormal idiosyncratic volatility hypothesis (as long as differences across firms in the levels of abnormal idiosyncratic volatility are relatively stable over time). Since it cannot distinguish between these alternative explanations, it is a weak test of Savor and Wilson (2011).

The model of Savor and Wilson (2011) also implies that the market beta of a portfolio of announcers should exceed that of non-announcers. Alternatively stated, the market beta of a long-short portfolio should be positive. To test this, we regress the yearly long-short portfolio returns of each country on the country's value-weighted market portfolio return. In untabulated results, we find that of the 20 countries with the greatest number of observations, the long-short portfolio betas are insignificantly different from zero for 14 of them; only three of them are significantly positive and three are significantly negative.[39] As a whole, these results are not consistent with the conjecture that earnings announcement risk is a driver of the announcement premium.

6. Summary and conclusions

We show that the earnings announcement premium, documented previously for U.S. stocks, extends globally. For the firms within our 46-country sample, a calendar-time strategy of holding shares of firms during their earnings announcement months and shorting them during all other months generates a monthly return of 59.7 basis points over our 1991–2010 sample period. Within a regression framework, controlling for size, momentum, and the book-to-market ratio, and introducing country fixed-effects, we find that the average monthly incremental return during earnings announcement months is a significant and economically large 95.5 basis points. The premium is pervasive over time and across countries. It is significantly positive in 16 of the 20 years of our sample period and for nine of the 20 countries with sufficient data to perform a within-country analysis. We also find that the phenomenon is more pronounced for the smallest stocks. This stands in contrast to Frazzini and Lamont (2007) who find either that there is no significant relation

[39] Using intraday price data, Patton and Verardo (2012) find evidence that betas increase on earnings announcement days for firms in the Standard and Poor's (S&P) 500 index. Due to our monthly horizon, we are less likely to pick up such a short-term effect.

Table 11
Earnings announcement premium and annual aggregate earnings growth.
This table reports the average coefficients for yearly regressions of:

$$AggGrowth_{jt+1} = \alpha + \beta_1 LS_{jt} + \beta_2 Mkt_{jt} + \beta_3 AggEarn_{jt} + \sum_{j=1}^{46} \gamma_j Country_j + \varepsilon_{jt},$$

where $AggEarn_{jt}$ is the aggregate annual earnings announced during calendar year t for all firms in country j, divided by the sum of the market values of these firms as of the end of calendar year t; $AggGrowth_{jt+1}$ is the aggregate difference between the annual earnings announced during calendar year $t+1$ and calendar year t for all firms in country j, divided by the sum of the market values of these firms as of the end of calendar year t; LS_{jt} is the long-short portfolio return for year t in country j, calculated alternatively by (1) summing the monthly long-short premia during year t, with each month's return weighted by the number of announcements that month (referred to as frequency weighting), and (2) compounding the monthly long-short premia; Mkt_{jt} is the value-weighted market return on Datastream's total market index for country j in year t; $Country_j$ is an indicator variable equal to one for country j and equal to zero, otherwise; and ε_{jt} is the regression residual for country j in year t. All returns and local earnings are denominated in U.S. dollars. All explanatory variables are truncated at the top and bottom 1% of their distributions. Also presented is the adjusted R^2 and the number of country-year observations in each regression specification. Below each coefficient value is the corresponding t-statistic. ***=Significant at the 1% level; **=significant at the 5% level; *=significant at the 10% level. See Table 1 for a detailed description of our sample.

	Intercept	*LS*	*Mkt*	*AggEarn*	*R*-squared (%)	# Of country-year obs.
(1) Frequency-weighted long-short return	**0.031****	0.042	**0.051*****	**−0.472*****	37.55	652
	2.12	*1.57*	*4.34*	*−4.12*		
(2) Compounded long-short return	**0.031****	0.018	**0.051*****	**−0.471*****	37.65	652
	2.06	*1.45*	*4.39*	*−4.21*		

between size and the announcement premium for U.S. stocks or that the larger firms generate a higher premium.

An advantage of working with a global sample is that we can exploit cross-country differences in the magnitude of the premium in order to shed new light on the reasons for its existence. Consistent with the conjecture that uncertainty over the earnings information to be released causes investors to demand higher pre-announcement returns, leading to the observed earnings announcement premium, we find that the size of the premium varies directly with the level of abnormal idiosyncratic volatility during the earnings announcement window. In contrast, there is no evidence that the premium is compensation for increased levels of systematic risk around the time of firms' earnings releases or that it is driven by heightened investor attention to earnings announcements.

If, as our results suggest, the premium is a function of abnormal idiosyncratic volatility around earnings announcements, a natural question to ask is why diversified arbitrageurs do not eliminate the premium. A major reason is likely to be transactions costs. Exploiting the announcement premium requires portfolio turnover of 100% per month, as each stock is bought at the start of its announcement month and sold at the end of the month. Griffin, Kelly, and Nardari (2010) estimate round-trip transactions costs internationally for the years 1994 through 2005 and find that, except for the largest companies over the last few years of their sample period, they exceed 1%. Given that the documented monthly international announcement premium is approximately 1%, arbitrage profits may be eliminated by these transactions costs.

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