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ScienceDirect
The International Journal of Accounting
42 (2007) 341 – 379
The International Journal of Accounting

Cross-sectional variation in the economic consequences of international accounting harmonization: The case of mandatory IFRS adoption in the UK ☆

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Abstract

This study examines the economic consequences for UK firms of the European Union's decision to impose mandatory IFRS. We hypothesize that the impact varies across firms and is conditional on the perceived benefit. We estimate a counter-factual proxy for a UK firm's willingness to adopt IFRS from the prior GAAP choices of German firms. We show that this proxy predicts cross-sectional variations in both the short-run market reactions and the long-run changes in cost of equity that are associated with the decision. This implies that mandatory IFRS adoption does not benefit all firms in a uniform way but results in relative winners and losers.

Keywords: International Financial Reporting Standards; Mandatory adoption; Economic consequences

☆ The authors are from the Manchester Accounting and Finance Group, Manchester Business School. We would like to thank Rashad Abdel-khalik, Willem Buijink (our discussant), Kevin Chen, Ole-Kristian Hope, William Rees and other participants at the Illinois Accounting Symposium 2006, an anonymous referee, and Christian Leuz for their valuable comments.
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0020-7063/$30.00
doi:10.1016/j.intacc.2007.09.007

1. Introduction

The mandatory adoption of IFRS[1] in the European Union (EU) is one of the largest regulatory experiments in financial reporting ever undertaken, and may eventually prove to be a vital step towards global GAAP harmonization.[2] The EU and European Economic Area (EEA) include 30 countries with integrated financial markets and more than 7000 listed firms. Almost all EU/EEA listed firms are legally required to adopt IFRS in their consolidated statements no later than 2005.[3]

In this paper, we examine the economic consequences of mandatory IFRS adoption for United Kingdom (UK) listed firms. We study both the short-term price response to news about IFRS adoption, and the changes in the implied cost of equity for a large sample of firms between a date before the mandatory adoption was expected and a date by which mandatory adoption was effectively certain.

The short-run share-price response and long-run implied cost of equity methods complement each other when testing the effect of mandatory IFRS adoption. The potential advantage of focusing on short-run abnormal returns is that we are able to isolate specific days when news affects all firms in the sample. The disadvantage is that it is reliant on precise identification of the event days. In particular it assumes that there has been no leakage of the policy deliberations to the market. Unfortunately the dates on which the probability of mandatory adoption of IFRS changed are debatable. In contrast, an advantage of using the implied cost of equity method is that it is not sensitive to the identification of specific dates — we simply exclude the period of uncertainty and test the difference between the implied cost of equity before and after the announcement period. However, the estimation of the implied cost of equity is also potentially problematic, because it is often difficult to control for all factors affecting the implied cost of equity over a long period of time. Thus we view the two methodologies as being complementary and we believe that their joint use should increase the robustness of our conclusions.

We hypothesize that UK firms vary in their willingness to adopt IFRS, because the costs and benefits of IFRS adoption are likely to vary across firms. In terms of the literature on accounting choice, the decision to mandate IFRS for UK quoted firms was unusual in the sense that it cannot be simply portrayed as the imposition of a restriction on the accounting choices of UK firms. Prior to 2005 UK firms were not permitted to adopt IFRS for UK financial-reporting purposes. After 2005, UK firms are not allowed to use pre-2005 UK GAAP in their consolidated statements for financial reporting purposes. Thus the EU decision changed the choice set for UK firms by mandating a new set of rules for financial

[1] International Financial Reporting Standards (IFRS) is the name of accounting standards produced by the International Accounting Standards Board (IASB).

[2] The EU's motive for adopting the regulation is the creation of a more transparent and efficient capital market that will facilitate a lower cost of capital for EU firms(EC16/06/2002).

[3] EC 16/06/2002 requires all listed firms in a regulated market to comply with IFRS in their consolidated statements no later than 2005 unless they are listed in non-member state and have been using internationally accepted standards prior to September 2002. Member countries can allow adoption to be postponed until 2007 for firms that comply with US-GAAP. The UK has decided not to use this option and all listed firms in a regulated market are, therefore, required to comply with IFRS from 2005.

reporting that some UK firms might have adopted voluntarily, if they had been given the choice. If UK firms had been given a choice between UK GAAP and IFRS it is logically possible that some would have chosen not to adopt IFRS, and some would have chosen to abandon UK GAAP in favor of IFRS. Thus it is possible that some UK firms would have been constrained by the EU's decision, while others would have been liberated.

For the purposes of this paper we need a counter-factual proxy for what choices UK firms would have made, if they had been given an option to choose between UK GAAP and IFRS. One possibility, which we explore in this paper, is to exploit the information in the choices made by firms in an economy similar to the UK, but where firms had the choice to adopt IFRS before 2005. In particular Germany is a major EU economy that allowed early adoption of IFRS and that also experienced extensive early adoption. This combination of Germany and the UK as two major EU economies, but with very different IFRS adoption processes, produces a unique setting for testing the factors affecting the economic consequences of mandatory IFRS adoption.

We hypothesize that the characteristics of voluntary/early adopters of IFRS or US-GAAP[4] in an EU jurisdiction that allowed voluntary adoption of international accounting standards (IFRS or US-GAAP) might serve as a viable proxy for how UK firms might have behaved given the same choice. In particular we focus on the choices made by German firms.[5] In Germany, listed firms have had the option to choose between an international accounting regime (IFRS or US-GAAP) and domestic standards for their consolidated statements since 1998.[6] Economic theory predicts that firms committing to an international accounting regime are those that perceive the greatest net-benefit. We measure the degree of similarity to German voluntary adopters by estimating a logistic choice-model using German data and calculating the probability of voluntary adoption in the UK based on this model. We use the estimated probability of voluntary adoption from our model based on German firms as a counter-factual proxy for the probability of voluntary adoption by UK firms.

The advantage of this approach is that it focuses on actual observed choices, there is no potential for response bias, and it is based on a large population of firms. The disadvantage of this approach is that the German GAAP and financial disclosure regime is not the same as UK GAAP. The choice between UK GAAP and IFRS for UK firms is not the same as the choice between German GAAP and IFRS for German firms. For example, German IFRS adopters will typically experience a greater leap in disclosure quality. Due to these differences one might expect two sets of determinants for firms' willingness to switch to IFRS, i.e., one set that is common across both countries and another set that is country-

[4] For brevity we describe German firms that comply with either IFRS or US-GAAP in 2002 as voluntary adopters in this paper.

[5] An alternative approach might be to ask firms directly what they might have done if they had been given a choice. However this approach is also problematic for the following reasons: 1) firms may not know what they would have done given the choice, 2) many firms may be unwilling to respond to the survey (typical response rates are 20–30%), 3) some firms may not tell the truth.

[6] In April 1998 KapAEG was adopted in Germany allowing listed firms the option only to comply with either IFRS or US-GAAP in there consolidated statements.

specific. For instance, factors that correlate with corporate governance may be less transferable from Germany to the UK. Convincing outside investors that the firm is committed to improved corporate governance may be an underlying motive that is more important to firms in Germany due to the ownership and legal system they operate in. The implications of this fact for our research design is that we run the risk of identifying some determinants from the German adopters that are not necessarily relevant to the UK firms, which would add noise and reduce the power of the tests based on the UK sample. Thus our analysis reports two sets of results for the UK. One set assumes that the choice model for the UK is the same as for Germany. The other set attempts to isolate the Germany-specific choice drivers from the common drivers.

We find that the common-driver set produces consistent results both for implied cost of capital changes and for the short-run market responses. In both cases we find a significant positive cross-sectional association between the economic response to mandatory IFRS adoption and our counter-factual proxy for the probability of voluntary adoption by UK firms.

The study makes two main contributions. First, understanding that the costs and benefits of IFRS adoption varies systematically across firms is important not only to countries that have already decided to make IFRS mandatory, but also to countries that are currently considering taking this step.[7] Second, the study also makes a novel methodological contribution, by showing that under certain circumstances the information contained in voluntary GAAP choices in one economy can predict the economic responses to a mandatory GAAP change in a similar economy.

The remainder of the paper is organized as follows. Section 2 reviews the literature in the area and Section 3 develops the testable hypotheses of the paper. Section 4 describes the methodology and sample including the key dates that changed the likelihood of mandatory IFRS in the EU and the calculation of the counter-factual proxy for voluntary adoption in the UK. Section 5 presents the results and discusses the implications. Section 6 summarizes the paper.

2. Literature review

2.1. Voluntary adoption

Until recently, empirical studies on the connection between GAAP changes and the cost of capital have focused on voluntary adoption of either IFRS or US-GAAP over domestic standards. The assumption is that the accounting regime affects the quality of information and that the quality of information in turn affects the cost of capital.

One stream of research examines proxies for the cost of capital either within an event study around the adoption of IFRS or US-GAAP, or cross-sectionally between firms that have adopted IFRS or US-GAAP and firms that use local-GAAP. Leuz and Verrecchia (2000) and Leuz (2003) take this approach by examining bid/ask-spreads, trading volume

[7] According to GAAP convergence (2002) over 90% of the 59 countries surveyed intend to convert national standards to IFRS.

and share price volatility as proxies for the information asymmetry component of the cost of capital. They find reduced information asymmetry when firms change from German GAAP (HGB) to either IFRS or US-GAAP, but no significant difference between IFRS and US-GAAP. Contrary to this conclusion, Daske (2006) finds no evidence of a reduced cost of capital when using both the residual-income valuation model and the Ohlson and Juettner-Nauroth (2005) abnormal earnings growth model to estimate the implied cost of equity. These three studies limit their sample to German firms, thus keeping the institutional settings constant. Cuijpers and Buijink (2005) use a European sample to test the affect of changing from local-GAAP to either IFRS or US-GAAP. They examine information asymmetry proxied by analyst following, forecast dispersion and stock return volatility and the implied cost of capital estimated using the method suggested by Easton, Taylor, Shroff, and Sougiannis (2002). They document a positive effect of adopting IFRS or US-GAAP on analyst following, but fail to find support for a lower implied cost of equity. Dargenidou, McLeay, and Raonic (2006) also use a European sample to test how the change from local GAAP to international accounting standards (IFRS or US-GAAP) affected the estimated cost of capital using the Ohlson and Juettner-Nauroth (2005) abnormal earnings growth model. Contrary to Daske (2006) and Cuijpers and Buijink (2005) they find that the cost of capital increased by more than 4% after voluntary adoption but that the effect is smaller for large firms. Daske, Hail, Leuz, and Verdi (2007) extend these studies by focussing on whether the impact varies with the degree of compliance. Survey evidence documents that compliance varies considerably among voluntary adopters (Cairns, 1999, 2000). Daske et al. (2007) show that the cost of capital is only reduced when adoption is serious (i.e., leads to improved accounting quality). The heterogeneity that they document in the voluntary setting arises from differences in compliance level. The heterogeneity that we explore in the mandatory setting arises because the willingness to adopt varies across firms. To summarize, prior research that uses proxies for either the cost of capital or components of the cost of capital have produced mixed results.

Another stream of research looks at the market reaction to the announcement of future compliance with IFRS or US-GAAP. The idea is that the market reaction around the announcement contains the change in the required risk premium, and thus impounds the change in the cost of capital. Pellens and Tomaszewski (1999) find insignificant market reactions to the announcement of future compliance with either IFRS or US-GAAP in Germany. The statistical power of their test is, however, low due to a sample size of only 16 firms. Karamanou and Nishiotis (2005) use an international sample of 54 firms adopting IFRS and show that firms experience abnormal positive returns around the announcement of future compliance with IFRS. They also find evidence that the positive market reaction is not identical among firms with different characteristics. Thus, firms with low valuations and high growth opportunities experience a stronger market reaction. Karamanou and Nishiotis suggest that firms use the adoption of IFRS to signal to the market that they are undervalued. This signalling motive does not apply to the mandatory-adoption setting as in the case of the UK. Indeed, the study of mandatory adoption differs from voluntary adoption in two ways. First, mandatory adoption eliminates the self-selection issues inherent in voluntary adoption. Second, the choice to voluntarily adopt is a signal that includes information in itself, which could be difficult to disentangle from the underlying issue a study seeks to examine.

2.2. Mandatory adoption

Three prior studies examine mandatory rather that voluntary adoption of IFRS. Comprix, Muller and Stanford-Harris (2003) examine abnormal returns around the dates of public announcements that increase the likelihood of mandatory IFRS in the EU. They apply the Sefcik and Thompson (1986) approach to evaluate the relationship between announcement returns and a number of firm and country characteristics. They find that firms that are a) audited by a big 5 auditor, b) located in countries that will experience the greatest increase in quality of financial information as a consequence of IFRS, and c) are subject to the highest level of legal enforcement experience significant positive returns. Apart from the nature of the auditor these characteristics are all country specific. Although the methodology of Comprix et al. (2003) is similar to the market-reaction test we conduct in our study, the underlying research question is different. While they predominantly examine differences among country characteristics we investigate the role of firm characteristics within one country, which keeps the institutional framework constant.

Armstrong, Barth, Jagolinzer, and Riedl (2006) also investigate market reactions to events that they argue would affect the likelihood of mandatory IFRS in the EU. Unlike Comprix et al. (2003), which studies the events from 2000 to 2002, they analyze later events between 2003 and 2004 that are related to the endorsement of the IFRS standards in general and IAS 32/39 in particular. In general, they find positive (negative) market reactions to events they classify as increasing (decreasing) the likelihood of mandatory IFRS and interpret this as evidence that investors perceive benefits of harmonized accounting standards under IFRS. The focus of Armstrong et al. (2006) is on whether mandatory IFRS is good or bad as perceived by investors. In our study, the focus is instead on the differences in the economic consequences of mandatory IFRS between firms that are likely to incur relative benefits and costs due to the decision. Differences in economic consequences are of particular interest when evaluating a mandatory policy change. Since all firms by definition are treated equally by such a policy, those that are disadvantaged by the policy are still forced to comply.

Instead of examining the short-term market reactions, Pae, Thornton, and Welker (2006) investigate the consequences of mandatory IFRS by looking at Tobin's Q. They show that firms with high agency cost measured by concentration of control and excess of the largest shareholder's voting rights over cash flow rights experience relative increases in valuation as a consequence of mandatory IFRS. The underlying idea in the study is that the shares of some firms trade at a discount due to weaker protection for minority shareholders and that mandatory IFRS force these firms to improve their disclosure, which in turn reduces the discount. Pae et al. (2006) use a European sample where early adoption is allowed in several countries and national disclosure quality generally is lower than IFRS. In this setting it makes sense to test whether the restrictions imposed by mandatory adoption of IFRS have benefited some investors. Contrary to this we restrict our test to a UK setting where firms have not had the option to comply with IFRS voluntarily. Furthermore, the disclosure quality in our UK setting is generally high and it is unsure whether IFRS is an improvement for all firms. Our sample is more relevant for an investigation into whether or not relative benefits differ across firms than whether or not it enhances protection of minority investors.

The key contribution of our study to the literature is the focus on firm-specific cross-sectional differences in the economic consequences of mandatory IFRS. We define the firm-specific differences through a counter-factual proxy for willingness to adopt. The idea is that heterogeneity in the economic consequences of mandatory IFRS arises because some firms are forced to comply against their will, while others have net benefits and would have complied voluntarily had they been given the opportunity. We are not aware of any prior studies that connect the voluntary accounting GAAP commitment of firms in one country to the economic consequences of a mandatory policy in another country.

3. Hypotheses development

The starting point of our analysis is the assumption that the costs and benefits of IFRS adoption, relative to firm value, will vary across firms. The mandatory adoption of IFRS imposes two kinds of changes on the financial-reporting practices of firms. First, firms are required to adopt a new set of accounting-measurement rules that in some cases will have a material effect on a firm's reported earnings and balance-sheet values, and in other cases will not. Second, IFRS introduces a new set of required disclosures that in some cases will be greater than the original disclosure requirements and in other cases less.

Empirical research suggests that the cost of capital is related to both disclosure and measurement policies. Examples of such studies are Botosan (1997), that examines the association between disclosure levels and the implied cost of equity, and Francis, LaFond, Olsson, and Schipper (2004), that examines the relationship between earnings attributes and the implied cost of equity. Both studies find that a lower quality of information is associated with a higher cost of capital.The main hypothesis of this paper is (stated in alternative form):

H 1. The cross-sectional variations in the economic consequences of mandatory IFRS adoption by UK firms are related to the probability that the firm would have adopted IFRS voluntarily if it had been given the choice.

In order to convert H 1 into an empirically testable proposition we need to identify specific, measurable, economic consequences, and we need to specify how to model the probability of (counter-factual) voluntary adoption by UK firms. For the purposes of this paper we focus on two, potentially related, types of economic consequences. First we consider the market response to news about the decision by the EU to mandate IFRS. Second we consider the relative change in implied cost of equity between the time when the EU started to consider IFRS adoption and the time when the decision to adopt IFRS was effectively final and binding on all member states.

The main hypothesis is divided into two testable hypotheses (stated in alternative form):

H 1A. The stock price reaction of UK firms to announcements that increased (decreased) the likelihood of mandatory IFRS adoption is positively (negatively) related to their degree of similarity to the characteristics of German voluntary IFRS adopters.

H 1B. The change in the implied cost of equity of UK firms before and after the mandatory IFRS adoption decision is negatively related to their degree of similarity to characteristics of German voluntary IFRS adopters.

Both hypotheses exploit the fact that an informationally efficient market should rapidly incorporate the expected costs and benefits of IFRS adoption into share prices. That is to say, that if the market expects UK firms with characteristics similar to German early adopters to derive a relative benefit from IFRS adoption over other firms then such firms should experience a reduction in their relative cost of capital after future mandatory IFRS adoption became known, and a relatively positive (negative) response to news indicating that mandatory IFRS adoption was more (less) likely.

Hypothesis H1A tests how the market initially received the news of mandatory IFRS adoption. Hypothesis H1B tests how the market perceives mandatory IFRS adoption in the longer run. Consistent results for H1A and H1B should increase the robustness of the conclusion with regard to the main hypothesis.

In thinking about these hypotheses it is important to recognize that our focus is on the possibility that some firms may benefit more than others from the implementation of IFRS. In particular we do not deny the possibility that the value of IFRS adoption could be relatively greater in Germany than in the UK. Indeed, while there does seem to be a common perception that IFRS could be beneficial for German firms (Leuz and Verrecchia, 2000), the general perception of IFRS seems less favourable for UK firms. Studies such as Ginger and Rees (2001) show that UK-GAAP and IFRS are generally assumed to be very close. Some practitioners hold the belief that UK GAAP is of higher quality than IFRS.[8] However, despite such perceptions recent empirical evidence has found that the investment decisions of fund managers in the UK have been affected by the transition to IFRS.[9]

Furthermore, we are not concerned in this paper with testing the overall effect of mandatory IFRS adoption on the cost of capital of UK firms. It could be that the median level of accounting information quality decreases in the UK due to IFRS being of a lower quality than UK accounting standards, but at the same time the effect of IFRS adoption could be smaller for UK firms similar in characteristics to German volunteer adopters. In this case our main alternative hypothesis would be accepted in the UK even though the overall affect of introducing IFRS was to decrease the quality of financial statements and increase the cost of capital.

Another issue we face in relation to the changes to measurement and disclosure policies due to IFRS is the differences between Germany and the UK. Economic intuition suggest that a firms' accounting policy choice is driven by it's perception of net benefits. If the perceived net-benefits are at least partly determined by a function of measurement and disclosure issues then it is unlikely to be identical for Germany and the UK due to their institutional differences. Ball (2006) and Nobes (2006) both analyse how differences in

[8] See Accountancy, January 1999, p. 6, and Accountancy, May 1999, p. 77, for examples.

[9] In a PwC/MORI (2005) survey of fund managers, 70% said that they found the first IFRS information fairly useful or very useful, 29% reported that the disclosure had influenced them to disinvest from a company, and 21% said that they had been influenced to not invest in a company and 13% had been influenced to invest.

financing, ownership, legal, and taxation systems across countries influence the development of their domestic accounting regulations and suggest that this is likely to have effects on the implementation of IFRS.

Thus, to the extent that these institutional differences affect the decision of IFRS adoption, there will be country-specific factors in Germany that are not transferable to the UK context and therefore induce noise into our counter-factual proxy for a UK firm's willingness to adopt IFRS. That is to say, our research design potentially underestimates the differences in economic consequences and therefore we run the risk of failing to reject that no differences exist when differences actually exist. Thus this issue, in effect, loads the dice in favor of our null hypothesis. The fact that we are able to reject the null, in spite of this issue, suggests that even stronger results in support of our alternative hypotheses could be found if a more powerful counter-factual proxy could be designed than the one we use here.

4. Methodology and sample

4.1. Development of the counter-factual proxy

In this section we explain the development of the counter-factual proxy for UK firms' willingness to adopt IFRS based on their degree of characteristics similar to German voluntary IFRS adopters. We use the observed voluntary GAAP choices of German firms to predict which UK firms would be more likely to adopt IFRS given the same choice. The following logistic regression models are used to explain the choice of German firms:

$$\text{Adopter}_i = \alpha_0 + \alpha_1 \text{FS}_i + \alpha_2 \text{DTM}_i + \alpha_3 \text{LMV}_i + \varepsilon_i \tag{1}$$

$$\text{Adopter}_i = \beta_0 + \beta_1 \text{FS}_i + \beta_2 \text{DTM}_i + \beta_3 \text{LMV}_i + \sum_{k=4}^{10} \beta_k \text{INDDUM}_{k,i} + \varepsilon_i \tag{2}$$

$$\text{Adopter}_i = \gamma_0 + \sum_{k=1}^{7} \gamma_k \text{INDDUM}_{k,i} + \varepsilon_i \tag{3}$$

The dependent variable (Adopter) is assigned the value of one if firm i complies with an international accounting regime in 2002 and the value of zero otherwise.[10] FS is the foreign sales divided by total sales, DTM is the long-term debt divided by the sum of its long-term debt and market value, LMV is the natural logarithm of the market value, and INDDUM are seven industry dummies set equal to one for the industry for which the firm belongs and zero otherwise. Model 1 (Eq. (1)) only includes the three firm-characteristic variables that measures foreign sales, leverage, and size. To capture the long-run norm, we

[10] We do not distinguish between IFRS and US-GAAP. The reason is that we are interested in firms that have net-benefit of committing to an international accounting regime regardless which one. We do, however, re-run the models excluding firms complying with US-GAAP. The results are consistent in all material aspects, which is also consistent with the result of Leuz (2003).

use the five-year mean value of these variables from 1998 to 2002. Model 2 (Eq. (2)) adds seven industry dummies to the existing independent variables in Model 1 in order to incorporate any industry effect. Model 3 (Eq. (3)) includes only these industry dummies. We group firms into industries using the Worldscope industry classification. The three variants enable us to observe whether firm characteristics or the industry is more relevant in capturing the economic consequences of mandatory IFRS. In the process of developing our choice model from the German sample, we experimented with additional variables such as operating margin as a proxy for performance, sales growth as a proxy for growth, and operating cash flow as a proxy for finance need. Unfortunately, none of these were statistically significant. This is not surprising since existing studies (e.g., Ashbaugh, 2001; Cuijpers and Buijink, 2005; Harris and Muller, 1999; Leuz, 2003; Leuz and Verrecchia, 2000; Tarca, 2004) find mixed results and the way in which these variables influence accounting-policy choice remains under debate. Therefore, to avoid weakening the power of our counter-factual proxy in distinguishing UK firms' willingness to adopt the IFRS, we excluded these variables from our final models in Eqs. (1) and (2).

In determining the explanatory variables in Eqs. (1), (2), and (3), we refer to the firm characteristics identified in the existing literature on voluntary accounting-regime choices such as Harris and Muller (1999), Leuz and Verrecchia (2000), Ashbaugh (2001), Leuz (2003), Tarca (2004) and Cuijpers and Buijink (2005). These studies generally argue that the decision to adopt an international accounting regime is a function of financial performance, leverage, firm size, finance need, and cross-listing. Tarca (2004) adds foreign exposure and industry as explanatory variables. An important issue in applying these to our research design is to find proxies that are the same under international accounting standards and German domestic standards. Significant differences could result in wrong conclusions. For instance, leverage measured as total liabilities divided by book value of equity, is larger under HGB than under IFRS (Hung and Subramanyam, 2004). This relationship is driven by the book value of equity being measured significantly lower under HGB than under IFRS, which is consistent with HGB being more ex-ante conservative than IFRS. Hung and Subramanyam (2004) perform a survey of reconciliation items disclosed by firms that adopt IFRS for the first time. Hung and Subramanyam (2004, Table 4) show that differences are significant for total assets and the book value of equity. We therefore avoid these accounting figures.[11] Furthermore, theoretical predictions suggest that information asymmetry, analyst following, and liquidity motivate disclosure choices and proxies for these are, therefore, possible explanatory variables. Despite this there is a likelihood that their incorporation in the choice-model estimated on German data will induce causality and/or endogeneity problems because prior studies (e.g., Cuijpers and Buijink, 2005; Leuz and Verrecchia, 2000) suggest that these variables are influenced by voluntary adoption. In line with the existing literature on accounting standard choices, we therefore exclude these variables (e.g., Ashbaugh, 2001; Cuijpers and Buijink, 2005; Harris and Muller, 1999; Leuz, 2003; Leuz and Verrecchia, 2000; Tarca, 2004).

[11] In Eqs. (1) and (2) our measure of leverage is affected by the accounting regime. According to Hung and Subramanyam (2004) total liabilities tend to be lower under HGB than under IFRS. This means that our measure of leverage will be biased towards zero, thus underestimating leverage's effect on the choice of an international accounting regime.

Due to the difficulty in identifying variables that are available under the same definition in both Germany and the UK, as well as limitations in data availability, we exclude direct proxies of corporate-governance variables in our study. However, we do not rule out the possibility that the leverage variable and/or industry dummies we include in Eqs. (1), (2), (3) may indirectly capture cross-sectional variations in corporate governance structure. Existing studies argue that leverage proxies the level of "insider" finance available to the firm, which in turn affects a firm's disclosure incentives, because insiders do not rely on public disclosures (e.g., Cuijpers and Buijink, 2005; Meek, Roberts, and Gray, 1995; Zarzeski, 1996). Variations across industries have also been documented for governance mechanisms such as ownership structure (Demsetz and Lehn, 1985), executive compensation (Aggarwal and Samwick, 1999), and board structure (Agrawal and Knoeber, 2001). Gillian, Hartzell, and Starks (2003) also find that the strength of monitoring through board and charter provisions is industry specific. Another variable that could possibly serve this purpose is free-float but the trade-off here is a significantly reduced sample size. Since Leuz and Verrecchia (2000) find that free-float is insignificant (p-value 0.96) in Germany, leaving this variable out of the equation is unlikely to bias our results significantly. In terms of finance need, the most commonly applied proxy in the voluntary-adoption literature is a dummy variable taking the value of one if the firm issued equity after adoption, which generally turns out significant in prior studies (e.g., Ashbaugh, 2001; Harris and Muller, 1999). The intuition is that firms planning to issue capital in the future are more likely to adopt voluntarily to decrease the cost of capital. However, in the context of mandatory adoption this line of argument does not hold and it is unclear from which time period we should collect the data in order to make the test comparable to prior studies. An alternative approach would be to define it in terms of capital intensity (long-term assets/total assets), following Leuz and Verrecchia (2000). The problem of applying this proxy in our study is that both long-term assets and total assets are highly affected by the change from German GAAP to an international accounting regime. This makes it difficult to distinguish between the effect of changing the accounting regime and the effect of the finance need the proxy was meant to capture. Our existing models avoid the empirical measurement problems of corporate governance and finance need variables. Assuming that corporate governance, finance need, and other possible variables have accounting-policy-choice implications, their exclusion would reduce the explanatory power of the counter-factual proxy applied in the UK and reduce the chances of finding results in support of our hypothesis. While we do not deny the theoretical relevance of these variables to the context of our study, we leave them for future research as the objective for this paper is to investigate whether there are cross-sectional differences in economic consequences not to identify all possible drivers of these differences.

We compute the counter-factual proxy for a UK firm's willingness to adopt IFRS as follows:

$$\theta_{1,j} = \hat{\alpha}_0 + \hat{\alpha}_1 FS_j + \hat{\alpha}_2 \mathrm{DTM}_j + \hat{\alpha}_3 \mathrm{LMV}_j \tag{4}$$

$$\theta_{2,j} = \hat{\beta}_0 + \hat{\beta}_1 FS_j + \hat{\beta}_2 \mathrm{DTM}_j + \hat{\beta}_3 \mathrm{LMV}_j + \sum_{k=4}^{10} \hat{\beta}_k \mathrm{INDDUM}_{k,j} \tag{5}$$

$$\theta_{3,j} = \hat{\gamma}_0 + \sum_{k=1}^{7} \hat{\gamma}_k \mathrm{INDDUM}_{k,j} \tag{6}$$

$$\mathrm{Pr}_{l,j} = \frac{1}{1 + e^{-\theta_{l,j}}} \qquad (l = 1, 2, \text{ and } 3) \tag{7}$$

where $\mathrm{Pr}_{l,j}$ are the probabilities of voluntarily IFRS adoption by firm j based on model l (i.e. Models 1, 2, and 3), coefficients $\hat{\alpha}_0$ to $\hat{\alpha}_3$, $\hat{\beta}_0$ to $\hat{\beta}_9$, and $\hat{\gamma}_0$ to $\hat{\gamma}_6$ are estimated from Eqs. (1), (2), and (3) respectively, and FS_j, DTM_j, LMV_j, and $\mathrm{INDDUM}_{k,j}$ are proxies of foreign exposure, leverage, size, and industry dummies for firm j. The definition of these independent variables follows their German sample counterparts described earlier and are measured as the five-year mean over the same period of 1998–2002. Model 4 assumes that FS, DTM, and LMV are the common drivers between the UK and Germany. Model 5 assumes that the industry dummies that are relevant for German voluntary adoption are also relevant drivers for the UK. This assumption will be incorrect if the industry dummies proxy the governance motive for IFRS adoption by German firms.

4.2. Test of H1A

Hypothesis H1A assumes there is a positive relationship between a UK firm's degree of similarity to the characteristics of German voluntary IFRS adopters and their stock-price reaction to announcements relating to mandatory IFRS adoption. To test H1A we apply the Sefcik and Thompson (1986) portfolio-weighting approach commonly used to test the effect of firm characteristics on stock market reaction to time clustered events (e.g., Comprix et al., 2003; Li, Pincus, and Rego, 2004). This approach involves the following steps. First, construct a matrix F defined as follows:

$$F = [\mathrm{Int} \quad \mathrm{Pr}] \tag{8}$$

where Int is an $N \times 1$ vector of one and Pr is an $N \times 1$ vector of the degree of UK firms' similarity to German volunteer adopters (based separately on Models 1, 2, and 3) and N is the number of sampled UK firms. Second, create portfolio weights W as follows:

$$W = \begin{bmatrix} W'_{\mathrm{Int}} \\ W'_{\mathrm{Pr}} \end{bmatrix} = (F'F)^{-1}F' \tag{9}$$

where W'_{Int} is the row of portfolio weights based on Int, W'_{Pr} is the row of portfolio weights based on Pr, and F is the $N \times 2$ matrix defined in Eq. (8). Third, compute the returns (R_{Pr}) of the portfolio weighted on the information pertaining to Pr as follows:

$$R_{\mathrm{Pr},t} = W'_{\mathrm{Pr}} R_{i,t} \tag{10}$$

where $R_{i,t}$ is the $N \times 1$ vector of individual firm stock returns on day t, and t covers 521 trading days from 01/01/1999 to 31/12/2000. Finally, we run the following time-series regression:

$$R_{\mathrm{Pr},t} = \alpha + \beta R_{m,t} + \sum_{k=1}^{K} \delta_k D_{k,t} + \varepsilon_t \tag{11}$$

where α is the intercept, β is the risk coefficient, $R_{m,t}$ is the return on Financial Times All Shares Index on day t, δ_k is the risk-adjusted abnormal returns pertaining to event k, $D_{k,\mathrm{t}}$ is a dummy variable for the kth event during the three-day period (days −1, 0, and +1 relative to the announcement date) and is set to one (−1) if the event is assumed to be favorable (unfavorable) to mandatory IFRS adoption and zero otherwise. Eqs. (9) and (10) can also be implemented on the *Int* portfolio, but for brevity we do not report the results. The risk-adjusted abnormal return δ_k reflects the effect of Pr on the stock price reaction to the events examined. Sefcik and Thompson (1986) argue that these estimates are equivalent to those in a cross-sectional regression of abnormal returns on firm characteristics but fully control for the cross-correlation and heteroskedasticity in firm disturbances, which is essential in time-clustered event studies.[12]

Eq. (11) estimates the relationship between UK firms' degree of similarity to German voluntary IFRS adopters and their stock market reaction to the relevant announcements. In H1A, we expect that UK firms with higher Pr values should enjoy a relatively more positive market reaction to announcements that are favourable to mandatory IFRS adoption. This is because they share greater similarity in characteristics to the German voluntary adopters whose accounting regime commitment is due to their perceived net benefits.

4.3. Test of H1B

Hypothesis H1B assumes that long-run changes in the cost of capital of UK firms after the mandatory IFRS adoption decision is negatively related to their degree of similarity to the characteristics of German voluntary adopters. To test this hypothesis, we calculate the implied cost of equity capital based on the Ohlson and Juettner-Nauroth (2005) abnormal earnings valuation model and the Easton (2004) PEG model. Under the Ohlson and Juettner-Nauroth (2005) approach, the implied cost of equity capital (ICE_{OJ}) is defined as follows:

$$\mathrm{ICE}_{OJ} = A + \sqrt{A^2 + \left(\frac{\mathrm{eps}_{t+1}}{P_t}\right)\left[\left(\frac{\mathrm{eps}_{t+2} - \mathrm{eps}_{t+1}}{\mathrm{eps}_{t+1}}\right) - (\gamma - 1)\right]} \tag{12}$$

$$A = \frac{1}{2}\left[(\gamma - 1) + \left(\frac{\mathrm{dps}_{t+1}}{P_t}\right)\right] \tag{13}$$

where eps_{t+1} and eps_{t+2} are one- and two-years ahead analyst-consensus forecasts of earnings per share, dps_{t+1} is the one-year-ahead analyst-consensus forecast of dividend per share, P_t is the current price, and $(\gamma-1)$ is the perpetual growth rate at which the short-term growth decays asymptotically. We follow Gode and Mohanram (2003) in setting the $(\gamma-1)$ equal to the risk free rate minus 3%, which is the long-term inflation rate. If $(\gamma-1)$ is negative, we set its value

[12] Our use of Sefcik and Thompson's (1986) approach to test market reaction is similar to Comprix et al. (2003), which studies a similar context. However, another study by Armstrong et al. (2006) applies the traditional cross-sectional regression of announcement time of abnormal returns on company characteristics. To address the cross-correlation issue, they implement a separate test to compare announcements with randomly sampled non-announcement returns. Nevertheless, this methodology could not simultaneously address the time-clustering issue along with the research design seeking to test the relationship between market reaction and company characteristics. The Sefcik and Thompson's (1986) approved is directly designed for this purpose.

to zero following Claus and Thomas (2001). Following Chen, Jorgensen, and Yoo (2004), when $\text{eps}_{t+2} < \text{eps}_{t+1}$ we assign short-term earnings growth ($\text{eps}_{t+2} - \text{eps}_{t+1}$) to zero. When the value inside the root is negative, we assume the $\text{ICE}_{OJ} = A$. Under the Easton (2004) approach, the implied cost of equity capital (ICE_{PEG}) is calculated as follows:

$$\text{ICE}_{\text{PEG}} = \sqrt{\frac{(\text{eps}_{t+2} - \text{eps}_{t+1})}{P_t}} \tag{14}$$

These two models are preferred over the residual-income valuation model applied in Gebhardt, Lee, and Swaminathan (2001) as they do not require clean-surplus assumptions. The value of ICE_{PEG} is equivalent to ICE_{OJ} if we assume $(\gamma - 1) = 0$ and $\text{dps}_{t+1} = 0$. Because ICE_{PEG} requires that $\text{eps}_{t+1} < \text{eps}_{t+2}$, it tends to skew the sample toward growth stocks.

Due to inherent measurement problems associated with the estimation of the cost of equity from historical returns (Fama and French, 1997; Elton, 1999), inferring the cost of equity from analyst forecasts and market prices through accounting-based valuation models has emerged and promulgated in contemporary literature. Although some studies suggest possible weaknesses of such an approach (Easton, 2006; Guay, Kothari, and Shu, 2005) other studies show that their estimates capture common proxies of risk (e.g., Gebhardt et al., 2001; Gode and Mohanram, 2003; Botosan and Plumlee, 2005). In our study, we apply this established methodology to evaluate H1B, which associates our counter-factual proxy of UK firms' willingness to adopt IFRS with long-run changes in the level of market-perceived risk around the decision-making period. It provides a more direct proxy for the impact on firms' cost of equity capital than short-term market reaction tests. As mentioned earlier, tests of both H1A and H1B mutually complement each other and the triangulation of the results suggests a more powerful inference to our research question.

Following Daske (2006), we compute the implied cost of equity capital on a monthly basis from January 1996 to December 1998 (pre-announcement period) and September 2001 to October 2004 (post-announcement period). We calculate the change (ΔICE_{OJ} and ΔICE_{PEG}) by subtracting the median implied cost of equity of the pre-announcement period from the median implied cost of equity of the post-announcement period.[13] To determine the relationship between UK firms' degree of similarity to German voluntary adopters and changes in cost of capital, we run the following cross-sectional regressions:

$$\begin{aligned} \Delta\text{ICE}_{\text{OJ},j} = {} & \lambda_0 + \lambda_1 \text{Pr}_j + \lambda_2 \Delta\text{MV}_j + \lambda_3 \Delta\text{BM}_j + \lambda_4 \Delta\text{DM}_j + \lambda_5 \Delta\text{SG}_j \\ & + \lambda_6 \Delta\text{OPM}_j + \varepsilon_j \end{aligned} \tag{15}$$

where $\Delta\text{ICE}_{\text{OJ},j}$ is the change in implied cost of equity capital of firm j from pre- to post-announcement period, Pr_i is the degree of similarity with German voluntary adopters of firm j (based separately on Models 1, 2, and 3), ΔMV_j, ΔBM_j, ΔDM_j, ΔSG_j, and ΔOPM_j are the changes in three-year median-market value, book-to-market value, debt-to-market value, sales growth, and operating-profit margin, respectively, for firm j from pre- to post-

[13] Our research design seeks to detect cross-sectional variations at a company-specific level. To achieve this, we need to estimate the changes of implied cost of equity on an individual-company basis instead of using the approach that Easton et al. (2002) suggest.

announcement period.[14] Market value, book-to-market values, and debt-to-market values have been confirmed by previous research to be correlated with implied cost of equity (Botosan & Plumlee, 2005; Chen et al., 2004; Gebhardt et al., 2001; Gode & Mohanram, 2003) and are, therefore, incorporated as control variables. In addition, we control for growth proxied by sales growth and profitability proxied by operating-profit margin. The same regression of Eq. (15) is applied using ΔICE_{PEG} as a dependent variable. The coefficient λ_1 gives the relationship between Pr and long-run changes in cost of capital, after controlling for changes in various firm-specific attributes over the same period. In H1B we assume that there should be a negative relationship between the Pr and long-run changes in the cost of capital. This is again based on the assumption that higher Pr firms bear greater resemblance to German firms that adopted an international accounting regime and therefore have a net economic benefit upon the adoption of IFRS.

4.4. Key dates

The timeline of events leading to the mandatory adoption of IFRS is crucial for this study. For the event study we need to know the days when the market revised its expectations about IFRS adoption. For the cost of capital study we need to identify a pre-decision period when the issue of IFRS adoption was far from being resolved and a post-decision period when the issue of IFRS adoption was clearly settled.

To narrow down the period when expectations changed we first searched The Financial Times and Accountancy for all articles related to IFRS between 1st January 1999 and 31st December 2002. 1999 was the first year where commentators began concrete discussions on mandatory IFRS adoption and 2002 was the year the final directive was formally adopted by the council of ministers. This search revealed that after 31 December 2000 most commentators expected mandatory IFRS adoption in the EU by 2005.[15] This is prior to the formal adoption of the directive but consistent with evidence from Binder (1985) that suggests that formal regulatory announcements are generally anticipated in event studies using public announcements. We, therefore, narrowed down the search from 1st January 1999 to 31st December 2000. Seven events appeared relevant. These events are tabulated in Table 1.

Event 1 is the commission's presentation of its preferred option to the Financial Services Policy Group. The preferred option included all the important elements of the final directive. This event was not widely discussed in the Financial Times but due to the early stage and the importance of the Financial Services Policy Group this event is included. Thus, we classify Event 1 as favorable.

The year after the first event is the period with most uncertainty. It involved prolonged discussions about the future structure of the International Accounting Standard Committee (IASC, the IASB's predecessor). US stakeholders, among them the Securities and Exchange Commission (SEC), wanted less political influence in the accounting standard-setting process, while the EU commission took the opposite view. The future structure of the IASC was the

[14] We calculate the percentage change for market value.

[15] E.g., KPMG's (2002) comment on the final adoption of the directive: "In June 2002, the Council of Ministers of the European Union adopted the much anticipated regulation requiring all listed groups in the European Union to apply International Financial Reporting Standards (IFRS) in their financial statements by 2005."

Table 1
Events changing the likelihood of mandatory IFRS adoption in the EU

Event	Effect on likelihood	Date	Source	Description
1	Favorable	28/01/1999	FT04/04/1999	The Financial Services Policy Group, representatives of EU finance ministers, were presented with the European Commission's preferred option for accounting harmonization
2	Unfavorable	22/03/2000	IASC website 1)	Preliminary announcement of IOSCO endorsement of IFRS
3	Unfavorable	18/05/2000	FT25/05/2000	Formal IOSCO endorsement of IFRS
				IOSCO conditional endorsement decreased the likelihood of mandatory IFRS in the EU because it was the culmination of US influence in the IASC (later IASB)as it was connected to a changed structure much as proposed by the FASB and a changed team of trustees with more influence from the US. Furthermore, the OSCO's conditional endorsement was disappointing to the commission because it allows countries a series of opt-outs. Opt-outs that were expected to be used by the US
4	Favorable	09/06/2000	IASC website 1)	Fritz Bolkerstein, EU commissioner, makes a preparatory announcement that IFRS will be proposed as compulsory for all EU-listed companies
5	Favorable	13/06/2000	FT14/06/2000	The European Commission's propose that all listed EU companies should prepare their consolidated financial reports in accordance with IFRS
6	Favorable	17/07/2000	FT17/07/2000	ECOFIN meeting supporting the commission's proposal on this day and Britain calls for the completion of the European single market in financial services to be brought forward
7	Favorable	27/11 2000	FT28/11/2000	PricewaterhouseCoopers survey published. The survey shows support for IFRS among CFOs of European companies

This table presents the events that changed the likelihood of mandatory IFRS adoption in the EU. The events were identified by searching The Financial Times (FT) and Accountancy from January 1 1999 to December 31 2000. The dates shown are the event dates. The two dates marked by 1) are obtained from Comprix et al. (2003). 2) refer to Financial Times 25/05/2000 "Brussels' lost voice" for further discussion of the implication of IOSCO endorsement and EU mandating IFRS. Favorable means that the event increased the likelihood of mandatory IFRS in the EU and unfavorable means that the event decreased the likelihood of mandatory IFRS in the EU.

main obstacle for the International Organization of Securities Commissions' (IOSCO) endorsement of IFRS. IOSCO is an organization of securities commissions in the world working to promote high standards of regulation (www.iosco.org). IOSCO endorsement was to be the culmination of the core standard project. A process was begun in 1997, jointly by IOSCO and IASC, to reach a core set of accounting standards to be used for cross-border listings (www.iasb.org). The discussion on structure ended in May 2000 when the IASC decided to follow almost entirely the US proposal for a new structure. In addition to this restructuring, a number of high-ranking positions in the IASC's board of trustees went to US officials. The new structure and the US influence on the board of trustees paved the way for the IOSCO endorsement.[16] The endorsement itself turned out to be only conditional. It allowed

[16] Accountancy June 2000, p. 7: "The (core standard) project took a long time, partly because of the SEC's apparent hostility to anything but US GAAP. However, since the IASC finalized its restructuring proposals – in which the SEC played a major role – it seems to have moderated its approach."

countries a number of significant opt-outs.[17] The new structure, the appointment of US trustees, and the weak conditional endorsement lead to a decreased likelihood of mandatory IFRS adoption in the EU. We use the dates of endorsement and expected endorsement as proxies for moves by the IASC toward the US and away from the EU commission. Thus, we classify Event 2 and Event 3 as unfavorable.

The period of uncertainty ended in the second half of 2000 when four announcements made it clear that the EU would proceed with the regulation regardless of the new structure of the IASC and its reduced influence on the board of trustees. In June and July the commission pledged its support first through the responsible commissioner and second by a formal communication. Third, ECOFIN, consisting of the EU fiscal ministers, supported the commission's communication. And fourth, towards the end of 2000 a survey showed strong support among firms in the EU for mandatory adoption of IFRS. We classify Events 4 to 7 as favorable.

Although, all dates of changed likelihood are in 1999 and 2000 we acknowledge that some uncertainty might still have existed at the beginning of 2001. We therefore exclude the first nine months of 2001 in the analysis of long-run changes in the implied cost of capital. We define the pre-announcement period as the 36 months from 01/01/1996 to 31/12/1998 and the post announcement period as the 36 months from 01/10/2001 to 30/09/2004.

When building the counter-factual proxy (in Section 4.1) used to extract the characteristics of voluntary adopters, we need to define a voluntary *commitment* and to connect this definition to the choice of an international accounting regime. Following Leuz and Verrecchia (2000) we define a voluntary commitment as a decision by the firm about what it will disclose before it knows the content of the information. A decision to disclose after the information is known to the firm is a voluntary disclosure. The voluntary adoption of an international accounting regime is a commitment because it is not possible to change back to domestic standards in years where the firm for some reason decides compliance is undesirable.

The strength of a commitment is determined by a combination of how rigid it is and how long into the future the commitment stretches. Thus, a commitment made before the EU decision to require mandatory IFRS is stronger than a commitment made after the decision. This is because the latter only lasted till 2005, whereas the length of the former was unsure at the time it was made. From the above discussion of key dates, we know that most of the uncertainty as to whether IFRS would become mandatory had diminished by the end of 2000.

[17] Accountancy, June 2000, p.7 "... the US regulator (SEC) still has concerns about the quality of IFRSs and is thought to be behind IOSCO's less-than-wholehearted endorsement." FT The Financial Times, May 25, 2000 p2: "The IOSCO agreement will appear anaemic to the Commission. While committing to international standards, the deal permits national regulators several degrees of freedom. They are allowed to require extra disclosure, to apply their own interpretations and to demand items be reconciled to domestic standards.Brussels will portray this, with some justification, as an exercise in American ego massaging. The US Securities and Exchange Commission thinks its standards are the best, implying that some IASC rules are deficient by comparison. The pick and choose approach of the IOSCO deal could allow the US to keep its beloved rules more or less intact.The European Commission will perceive similar US dominance in the team of trustees. Mr Volcker is backed up by another regulatory heavyweight in the form of David Ruder, a one-time SEC chairman. None of Europe's seven representatives have held such elevated positions."Paul Volcker was appointed within days of the IOSCO endorsement. The above view is further supported by a commentary in the Financial Times: The Financial Times, June 6, 2000 p25: ....."The Commission is equally critical of the IOSCO deal. It views as disastrous a decision to allow countries a series of opt-outs from global rules. Regulators will have the power to require reinterpretation, greater disclosure and reconciliation to national standards."

The final decision was, however, not formally made before 2002 even if no dates in 2001 and 2002 appear to significantly change the likelihood of mandatory IFRS. Furthermore, the decision to change accounting regime requires a certain period of preparation. KPMG (2002) states that firms need to start the process of transition in 2002 if they are to be ready to comply with IFRS from 2005. Based on these factors we use the accounting-regime choices of German firms in 2002 as the dependent variable in the choice model. That is to say, firms that complied with an international accounting regime no later than 2002 made a commitment before knowing that IFRS would later become mandatory and, therefore, it is assumed that they perceived net economic benefits from committing to comply with IFRS.

Like Comprix et al. (2003), our study differs substantially from Armstrong et al. (2006) in the choice of time frame to extract announcement events. Their market-reaction study is conducted on a set of later events that took place between 2002 and 2004. The events are all related to EFRAG's endorsement of IFRS in general and IAS 32/39 in particular. They argue that these events capture a change in the likelihood of mandatory IFRS. There is no doubt that these events are significant in the process towards mandatory IFRS adoption worldwide, but it is unlikely that they capture any changes in the likelihood of mandatory IFRS adoption in the EU. Firms adopting IFRS on the 1st of January 2005 have 1st of January 2004 as the day of transition (firms with a calendar fiscal year). In order to be able to meet this deadline firms need a certain period of preparation. KPMG estimates that firms adopting in 2005 need to begin preparation in 2002 (KPMG 2002). In addition to this, the final decision to impose mandatory IFRS was made in June 2002 (EC16/06/2002). Thus it is unlikely that any significant uncertainty as to whether IFRS (except for IAS 32/39) would become mandatory remained in 2003 and 2004.

4.5. The German sample

The German sample is based on all existing and dead firms available from Datastream. We exclude financial institutions, firms with negative common equity, preferred shares, foreign firms, and those cross-listed on a non-German stock exchange. Since cross-listed firms are often required to provide higher-quality disclosure, their motive, cost, and benefit of adopting IFRS voluntarily are likely to be different from the rest of the firms in our German sample. Apart from that, since the UK-GAAP demands relatively higher disclosure standards than the HGB, the impact of involuntary disclosure due to cross-listing is higher for German than UK firms. This in turn leads to a disparity in the perceived benefit of IFRS adoption between firms from these two countries that are cross-listed abroad. Thus, the inclusion of cross-listing in the Pr value estimation is likely to pick up effects that are neither comparable nor transferable between the two samples. Following Daske (2006) and Karamanou and Nishiotis (2005) we should be concerned with the quality of information on accounting standards in commercial databases. The possibility of errors is largest for domestic standards, because a classification identical to the year before rarely results in an error when the firm already complies with an international accounting regime. This is because firms rarely switch from international to domestic standards. The opposite is not true. We, therefore, test all German firms classified as domestic-standard firms in their annual reports for 2002. For firms classified as complying with IFRS or US-GAAP we match the Compustat classification to the Datastream classification and check all data that do not agree with the annual reports for 2002. If we are unable to find the annual report for 2002 in Thompson One Banker we classify the accounting-standards variable for that firm as missing. The data for all other variables in Eqs. (1), (2) and (3) are obtained from Datastream. The final German sample size for Models 1, 2,

Table 2
Descriptive statistics for the UK and Germany

Panel A: Descriptive statistics on firm-specific characteristics for Models 1 and 2

		Germany				UK							
						IBES sample				Full sample			
		OBS	FS	DTM	LMV	OBS	FS	DTM	LMV	OBS	FS	DTM	LMV
Model 1 Pr	Mean		35.0%	14.5%	11.8		24.3%	14.4%	12.0		23.2%	11.7%	11.2
	Median		34.1%	7.7%	11.7		15.2%	8.7%	12.1		9.2%	6.0%	11.1
	Standard deviation		27.5%	17.1%	1.6		26.4%	14.0%	1.5		31.0%	14.1%	1.7
	Sample size	389				469				1310			
Model 2 Pr	Mean		34.9%	14.6%	11.8		24.3%	14.4%	12.0		23.2%	11.7%	11.2
	Median		33.6%	7.6%	11.7		15.2%	8.7%	12.1		9.2%	6.0%	11.1
	Standard deviation		27.6%	17.1%	1.6		26.4%	14.0%	1.5		31.0%	14.1%	1.7
	Sample size	382				469				1309			

Panel B: Descriptive statistics on industries (Model 3 only)

Industry (Worldscope classification)	Germany		UK			
			IBES sample		Full sample	
	OBS	%	OBS	%	OBS	%
Resource/utility (5200, 5800, 8200)	3	4	16	3	230	9
Electronics (3700, 4000)	104	17	74	16	377	15
Construction/manufacture (1300, 1600, 1900, 2500, 2800, 4900, 5500, 6100, 7300)	168	27	139	30	547	22
Retail/Transport/Media (6400, 7000, 7900)	40	6	62	13	265	10
Non-cyclical consumer goods (2200, 4600, 7600)	35	6	24	5	101	4
Drugs/healthcare (3400)	16	3	18	4	85	3
Other (3100, 6700, 8500, 8800)	231	37	136	29	933	37
Total	623	100	469	100	2538	100

Panel C: German firms' accounting-standard choice

	Model 1		Model 2		Model 3	
	OBS	%	OBS	%	OBS	%
Adopter						
IFRS	129	33	126	33	216	35
US-GAAP	70	18	66	17	114	18
	199	51	192	50	330	53
Non-adopters (HGB)	190	49	190	50	293	47

(continued on next page)

Table 2 (*continued*)

Panel D: Descriptive statistics for the counter-factual proxy of UK firms' willingness to adopt IFRS (Pr)

		Model 1	Model 2	Model 3
IBES sample	Mean	0.48	0.49	0.51
	Median	0.47	0.48	0.45
	Standard dev.	0.17	0.26	0.21
Full sample	Mean	0.43	0.45	0.53
	Median	0.41	0.43	0.63
	Standard dev.	0.26	0.26	0.19

This table shows the sample size and descriptive statistics on the data used to estimate the logistic regression models of voluntary adoption of international accounting regimes (IFRS/US-GAAP) in Germany and to calculate the counter-factual proxy (i.e., the likelihood of voluntary adoption) in the UK. Firm characteristics such as foreign sales (FS), leverage (DTM), and size (LMV) are measured as the five-year mean from 1998 to 2002. Industry dummies are assigned based on industry groups from Worldscope. Panel A presents the sample for Models 1 and 2 and Panel B provides the sample for Model 3. OBS is the sample size. FS is calculated as the foreign sales to total sales. DTM is calculated as long-term debt/(long-term debt + market capitalization). LMV is the natural logarithm of market value. Panel C describes the accounting-standard choice of German firms in the sample used in building the logistic regression models. Panel D presents the distribution of the counter-factual proxy of willingness to adopt IFRS among UK firms (Pr). IBES sample is applied to test of H1B. Full sample is only applied to test of H1A.

and 3 are 403, 386, and 641, respectively. Table 2 shows the size and relevant descriptive statistics for the German sample. Table 2 panel C describes the accounting standards that the German firms in our sample use in 2002 across the three models. Firms are equally divided between adopters and non-adopters. For instance, adopters account for 51%, 50%, and 53% in the dependent variable of the logistic regressions in Eqs. (1), (2), and (3) (Models 1, 2, and 3), respectively. This even distribution ensures that we do not skew heavily toward either adopters or non-adopters in our estimation.

4.6. The UK sample

Our UK sample is based on all existing and dead UK firms in Datastream. To construct the full UK sample, we exclude financial institutions, firms with negative book value of equity, preferred stocks, foreign firms, and those cross-listed on a non-UK stock exchange. The purpose of the latter criteria is to address the issue that cross-listed firms are already under more scrutiny from foreign investors and analysts. Thus, the costs and benefits that cross-listed firms incur from mandatory IFRS adoption is not comparable with the average firms in our UK sample. However, in Section 5.5.3 we re-introduce ADR firms into the UK sample for robustness check. A relatively small number of firms that already complied with either US-GAAP or IFRS are also excluded.[18] The data for foreign sales, leverage, and size used to compute the counter-factual proxy in Eqs. (4), (5) and (6) are from Datastream. For Models 1, 2, and 3 we have 1310, 1309, and 2538 observations, respectively. As discussed in

[18] Compliance with IFRS or US-GAAP prior to 2005 is only possible by supplementary reporting (i.e. producing two sets of financial statements), thus imposing additional costs on the issuer.

Sections 4.2 and 4.3, we apply two sets of tests on the UK firms. The market-reaction test of the H1A can be implemented using the full sample. We obtain the daily returns of the individual firms and the Financial Times All Share Index from Datastream. The test of H1B imposes additional data constraints. To measure changes in implied cost of equity over the pre-announcement (January 1996 to December 1998) and post-announcement (September 2001 to October 2004) periods, we obtain analyst-earnings forecasts and actual current price from the IBES. The data to compute changes in other control variables over the pre- and post-announcement periods, i.e. size, book-to-market value, leverage, sales growth, and operating profit margin in Eq. (15), are from Datastream. This results in a reduced common sample of 469 observations. Panels A and B of Table 2 present the sample size and relevant descriptive statistics for the UK sample applied in our study. In comparison to the German sample, the level of dependence on foreign revenue appears smaller in the UK, although the difference is not significant. In terms of leverage and size, our samples for these two countries are fairly similar. Table 2 panel D describes the counter-factual proxy of willingness to adopt IFRS. Mean and median are both close to 50%, suggesting a fairly equal distribution. This ensures that our subsequent tests are not conducted on UK samples that are heavily skewed toward firms that are either very similar or different in characteristics to German voluntary IFRS adopters.

5. Empirical findings

5.1. The counter-factual proxy

Table 3 presents the results of the logistic regression models (Eqs. (1)–(3)) we use to extract the characteristics of German voluntary adopters.

The results of Model 1 suggest that large firms with a low level of debt financing and a large foreign exposure are most likely to adopt an international accounting regime voluntarily. This is consistent with existing studies on voluntary commitments. For instance, Cuijpers and Buijink (2005) and Dargenidou et al. (2006) argue that larger firms tend to have lower disclosure costs. Meek et al. (1995) suggest that there is a greater demand for information and need for disclosure among large firms due to their higher political and agency costs. Tarca (2004) argues that higher foreign-sales firms have greater need to disseminate information abroad due to their exposure to foreign operations and capital markets. Tarca (2004) also suggests that firms with lower leverage depend more on equity capital and, therefore, have a greater need to reduce information asymmetry. Following the hypotheses developed in Section 3, we expect that firms with these characteristics have the largest net-benefit of mandatory IFRS adoption in the UK. In Model 2, the addition of industry dummies subsumes the relationship between voluntary adoption and leverage. This is not surprising since capital structures are known to be industry-specific. Another possible reason is that leverage may be jointly capturing the influence of a firm's dependence on equity financing as well as its corporate-governance structure. The latter effect in the leverage variable may be subsumed by the industry dummies since corporate-governance mechanisms are documented to be industry-specific (e.g., Agrawal and Knoeber, 2001; Gillian et al. 2003). However, as described in Footnote 12, the leverage coefficient is biased towards zero; we, therefore, keep leverage in Eq. (2). Model 3 applies only the industry dummies. Notice that the Electronics and

Table 3
Characteristics of German early adopters of international accounting regimes (IFRS or US-GAAP)

Variables	Predicted sign	Model 1	Model 2	Model 3
FS	+	2.0685	2.002	
		[5.01]	[3.91]	
DTM	−	−2.6003	−1.1696	
		[−3.68]	[−1.59]	
LMV	+	0.2630	0.3908	
		[3.40]	[3.89]	
Resource/utility	?		−1.6459	−0.7300
			[−2.64]	[−1.84]
Electronics	?		0.4025	0.6816
			[1.03]	[2.53]
Construction/manufacture	?		−1.6030	−1.0587
			[−5.01]	[−5.04]
Retail/transport/media	?		−1.4058	−1.2533
			[−2.29]	[−3.44]
Non-cyclical consumer goods	?		−3.8042	−3.3257
			[−4.01]	[−4.49]
Drugs/health care	?		NA	2.1857
				[2.10]
Intercept	?	−3.4120	−4.3002	0.5224
		[−3.77]	[−3.80]	[3.84]
Pseudo R^2		0.1143	0.2425	0.1274
Observations		389	382	623

This table presents the results of the logistic regression models of German firms' accounting-regime choice with *t*-statistics (in brackets). The dependent variable is set to one for German voluntary IFRS/US-GAAP adopters and zero otherwise and the independent variables are firm characteristics (Model 1), both firm-specific characteristics and industry dummies (Model 2), and industry dummies only (Model 3). FS is foreign sales to total sales. DTM is long-term debt/(long-term debt + market value). *LMV* is the natural logarithm of market value. See Table 2 for the Worldscope industry-classification code for the industry dummies, In Model 2, there is no observation for the Drug/healthcare firms because they all complied either with IFRS or US-GAAP.

Drug/Healthcare industries are significantly associated with voluntary adoption. This may be capturing the adoption of non-domestic standards by the high-tech growth firms listed in the *Neuer Markt* of Germany.

5.2. Market reaction

Table 4 presents the abnormal stock returns over three alternative window periods for portfolios based on UK firms during public announcements of events that changed the probability of mandatory IFRS adoption in the EU, i.e., the test of H1A.

We apply the Sefcik and Thompson (1986) approach and weight the UK firms' portfolios (Eqs. (8)–(10)) by their Pr values, which are the counter-factual proxies (Eqs. (4)–(7)) measuring UK firms' willingness to adopt IFRS (Eqs. (1)–(3)). We run a time-series regression of the UK Pr weighted portfolios' return on the market portfolio return proxied by FT All Shares Index and a set of event dummies representing the window periods covering each of the seven sampled events (Eq. (11)). The market-portfolio return controls for systematic risk and the event dummies capture the risk-adjusted abnormal returns pertaining to the corresponding event windows. These risk-adjusted abnormal returns are conditional on the Pr values through the portfolio weights. The event dummies are set to one (−1) for announcements that are favorable (unfavorable) to mandatory IFRS adoption in the EU and zero otherwise. For brevity, we only show the coefficients and *t*-statistics for the event dummies. The last column is based on a dummy variable that combines the values of all seven events. This enables us to make a joint inference that aggregates the net effect of Pr on market reaction across all seven events.

Panels A, B, and C of Table 4 show the results of the tests for days −1, 0, 1, days −1, 0, and day 0, respectively, across the full sample. Across all panels of Table 4 the portfolio weighted on Pr values calculated from Model 1 (Eq. (4)) has a significantly positive relationship with the net risk-adjusted abnormal returns of all sampled events as shown in the last column. The net risk-adjusted abnormal returns of this particular portfolio are 0.672%, 0.718%, and 0.485% in panels A, B, and C, respectively. These findings provide evidence in support of H1A by indicating that UK firms with higher Pr values, i.e., the degree of similarity with the German voluntary adopters, are associated with significantly higher market reactions relative to their lower Pr counterparts. To evaluate the role of the industry effect in the share-price response, we construct Models 2 and 3 (Eqs. (5) and (6)). (The former adds industry dummies into Model 1 and the latter contains only the industry dummies.) The net risk-adjusted abnormal returns for the portfolio weighted on Model 2 Pr are insignificant in panel C and only marginally significant in Panels A and B. The net risk-adjusted abnormal returns for the portfolio weighted on Model 3 Pr are statistically insignificant throughout all three panels. The general observation in Table 4 is that industry dummies only seem to add noise to our analyses of the UK sample despite the finding in Table 3, which shows that in the German sample industry membership has explanatory power for voluntary IFRS adoption.

We argue that a firm's willingness to adopt IFRS may be determined by factors that are common across countries as well as factors that are country-specific due to institutional differences. Our finding in Table 4, that the Pr values based on Model 1 are associated with significant net-market reactions, suggest that foreign sales, size, and leverage are common factors in both Germany and the UK. The observation that Pr values incorporating industry dummies in Models 2 and 3 exhibit a weaker association with abnormal returns could be due to the possibility that the industry effects are correlated with country-specific factors that are only applicable in Germany and not in the UK. We argue that this may be the case because the industry effects may be capturing the influence of corporate-governance structure on the adoption of IFRS by German firms. As shown in Table 3, the association between IFRS adoption in Germany and firm leverage, which is often applied to capture-corporate governance structure (e.g., Cuijpers & Buijink, 2005), is weakened in the presence of industry dummies. This provides weak evidence that the industry effect captures cross-sectional variations in governance structure of German firms and corroborates existing studies that document such relationship (e.g., Agrawal & Knoeber, 2001; Gillian et al., 2003).

In terms of individual events, notice that the portfolios weighted in Model 1 Pr have a consistently and significantly positive relationship with risk-adjusted abnormal returns associated with Event 1. According to Table 1, this is the day when the Financial Services Policy Group, comprised of representatives of the EU finance ministers, were presented with the European Commission's preferred option for accounting harmonization. This is our earliest sampled major announcement with favorable implications for the EU decision to impose mandatory IFRS adoption. Events 2 and 3 are considered to have unfavorable implications for mandatory adoption of IFRS. We expect the Pr value to be negatively

Table 4
Abnormal stock returns of UK firms conditional on their degree of similarity to German voluntary IFRS adopters during announcements of decisions to impose mandatory IFRS adoption in the EU

Events	1	2	3	4	5	6	7	Net
Dates	28-01-1999	22-03-2000	18-05-2000	09-06-2000	13-06-2000	17-07-2000	27-11-2000	
Effect	Favorable	Unfavorable	Unfavorable	Favorable	Favorable	Favorable	Favorable	
Panel A: Window -1,0,1								
Model	0.01117	0.01386	0.01204	0.00453	−0.00272	0.00725	−0.00311	0.00672
1 Pr	[4.59]	[1.97]	[4.28]	[1.16]	[−1.07]	[6.26]	[-0.96]	[3.14]
Model	0.00344	0.01471	0.01644	0.00362	−0.00665	0.01074	−0.01282	0.00488
2 Pr	[3.78]	[2.13]	[4.48]	[1.31]	[−3.87]	[10.31]	[−2.51]	[1.77]
Model	−0.00845	0.02054	0.02126	0.00328	−0.00898	0.01396	−0.02309	0.00351
3 Pr	[−6.20]	[2.04]	[6.64]	[2.58]	[−4.54]	[4.21]	[−2.38]	[0.84]
Panel B: Window -1,0								
Model	0.01381	0.01590	0.00934	0.00459	−0.0057	0.00867	0.00134	0.00718
1 Pr	[8.74]	[1.55]	[3.62]	[1.18]	[−10.48]	[12.46]	[1.98]	[2.55]
Model	0.00434	0.01947	0.01298	0.00374	−0.00833	0.00948	−0.00570	0.00573
2 Pr	[6.94]	[2.31]	[3.59]	[1.36]	[−6.84]	[13.91]	[−8.60]	[1.84]
Model	−0.00872	0.02762	0.01974	0.00345	−0.01093	0.00945	−0.00953	0.00527
3 Pr	[−4.86]	[2.27]	[4.60]	[2.76]	[−9.39]	[10.12]	[-10.55]	[1.11]
Panel C: Window 0								
Model	0.01592	0.00154	0.01284	−0.00072	−0.00569	0.00868	0.00136	0.00485
1 Pr	[28.87]	[2.87]	[20.98]	[−1.21]	[−10.26]	[12.47]	[2.00]	[1.76]
Model	0.00487	0.00759	0.01794	0.00007	−0.00981	0.00951	−0.00567	0.00353
2 Pr	[9.09]	[14.31]	[30.25]	[0.11]	[−18.23]	[13.93]	[−8.56]	[1.07]
Model	−0.00628	0.01044	0.02552	0.00228	−0.00950	0.00950	−0.00944	0.00327
3 Pr	[−8.75]	[14.55]	[31.89]	[2.58]	[−13.17]	[10.20]	[−10.44]	[0.72]

This table presents the abnormal stock returns and *t*-statistics (in brackets) of UK firms during announcements of mandatory IAS adoption in the EU based on the Sefcik and Thompson (1986) weighted portfolio approach over the test period of 01/01/1999 to 31/12/2000. The portfolios are weighted by the counter-factual proxy for willingness to adopt IFRS based on three models. Each model is estimated using a German sample where the dependent variable is set to one for German voluntary IFRS/US-GAAP adopters and zero otherwise and the independent variables are firm characteristics, which include foreign sales, leverage, and size (Model 1), both firm characteristics and industry dummies (Model 2), and industry dummies only (Model 3). The time-series returns of the weighted portfolios are regressed over the test period on market portfolio returns proxied by FT All Shares Index and a set of event dummies representing the window period covering each of the seven sampled events. The event dummies are set to one(−1) for announcements that are favorable (unfavorable) to mandatory IAS adoption in the EU and zero otherwise. Each column shows the coefficients estimated for the corresponding events and the *t*-statistics adjusted for heteroskedasticity. The last column is based on a dummy variable that combines the values of all seven events. Panel A shows the results when a three-day event window is used (−1,0,1). Panel B shows the results when a two-day event window is used (−1,0) and panel C shows the results when only the actual event day is used (0).

associated with market reaction to these cases. By setting the corresponding event dummies to −1 over the measurement windows, the anticipated negative relationships will appear with positive signs in our analysis. Notice that portfolios weighted on the Pr values calculated from all three models yields a significantly positive relationship with the risk-adjusted abnormal returns associated with Event 3 across all three panels. The results for Event 2 are also mostly significantly positive. These findings confirm that UK firms with higher similarity to German voluntary IFRS adopters experience a market reaction that is relatively more negative (for events 2 and 3) than those with lower Pr values. Events 4 to 7 are all considered to be favorable to EU mandatory IFRS adoption. Nevertheless, their results are generally mixed as they vary across event windows and models. In particular, events 5 and 7 are generally associated with negative returns contrary to our expectations. However, it is not uncommon in studies that use public announcement dates to observe that the ability of events to capture changes in expectations tend to decrease through time and that some events are associated with returns that are counter intuitive (e.g., Armstrong et al., 2006, Table 5 panel A). This is due to the challenge of identifying events that only change expectations regarding the issue under investigation and the tendency of later announcements only to re-confirm their earlier counterparts. Therefore, we focus our primary inference on the net effect and restrict the discussion of individual events to the robustness tests in Section 5.3.3.

5.3. Robustness tests of market reaction

To enhance the robustness of our findings in Table 4 we conduct additional tests on the relationship between our counter-factual proxy for the willingness to adopt and short term market-reaction tests. First, we examine if our findings also hold in the intersection sample with IBES, which we use in Sections 5.4 and 5.5 to test long-run changes in cost of equity for H1B. Second, we apply an alternative approach to further control for possible correlation with industry membership. Third, we examine the sensitivity of the results to the identification of specific events. Finally, we regress the market reaction on each characteristic of Model 1 in order to establish whether the results are driven entirely by one of the three characteristics.

5.3.1. IBES intersection sample

As explained in Section 4.6, our tests of short-run market reaction in Table 4 are based on a full sample while tests of long-run changes in cost of equity are based on a smaller sample subject to IBES data-availability constraints. We also implement the market-reaction test using the IBES sample for two reasons. First, if the short-run market reaction (H1A) and long-run changes in cost of equity (H2B) tests are to be mutually complementary, then the findings for the former based on a larger sample should also hold in the smaller sample in which we test the latter. Second, if the market-reaction results also exist in the IBES intersection sample, this will mitigate the possibility that our findings in Table 4 are driven by smaller firms not covered by the IBES database. In Table 5 panel A, we show that the relationship between announcement returns and the counter-factual proxy (Pr) estimated from firm-specific characteristics (Model 1, Eq. (4)) holds even in the smaller IBES sample. Notice that the net risk-adjusted abnormal returns in the last column is 0.586% and statistically significant.

Table 5
Robustness test of the connection between the similarity to German voluntary IFRS adopters and abnormal stock returns

Panel A: IBES sample

Events	1	2	3	4	5	6	7	Net
Model 1 Pr	0.00921	0.01441	0.01698	−0.00753	−0.00606	0.01011	−0.00211	0.00586
	[2.66]	[1.40]	[3.54]	[−12.70]	[−6.15]	[2.71]	[−0.53]	[1.95]

Panel B: Controlling for industry effects

Events	1	2	3	4	5	6	7	Net
Model 1 Pr	0.01386	0.01070	0.00577	0.00327	-0.00068	0.00171	0.00380	0.00593
	[5.52]	[2.08]	[3.62]	[0.99]	[−0.55]	[1.67]	[3.35]	[3.71]

Panel C: Net abnormal returns excluding individual events

Event excluded:	1	2	3	4	5	6	7
Model 1 Pr	0.00580	0.00533	0.00572	0.0086	0.00858	0.00663	0.00793
	[2.37]	[2.71]	[2.34]	[2.96]	[3.87]	[2.76]	[3.62]

Panel D: Portfolios build on individual characteristics

	MV	DTM	FS
Window -1, 0, 1	0.00077	−0.00315	0.00166
	[2.91]	[−0.75]	[1.67]
Window -1, 0	0.00061	−0.00427	0.00279
	[1.83]	[−0.85]	[3.00]
Window 0	0.00026	−0.00323	0.00299
	[0.72]	[−0.81]	[4.13]

This table presents the abnormal stock returns and *t*-statistics (in brackets) of UK firms during announcements of mandatory IAS adoption in the EU based on the Sefcik and Thompson (1986) weighted portfolio approach over the test period of 01/01/1999 to 31/12/2000. Panel A provides the results over the days −1, 0, 1 window of the seven events and their net effect for the portfolio weighted by Pr value that is estimated from Model 1 (Eq. (4)) using the IBES sample. Panel B provides the results over the days −1, 0, 1 window of the seven events and their net effect for the portfolio weighted by Pr value that is estimated from Model 1 (Eq. (4)) using the full sample. In this case, 25 industry dummies based on Worldscope classification are applied during the formation of the weighted portfolio to control for industry effect (see Section 5.3.2). Panel C provides the results over the days −1, 0, 1 window for the net effect after excluding each of the seven individual events in turn for the portfolio weighted by Pr value estimated from Model 1 (Eq. (4)) using the full sample. Panel D shows the results over days −1, 0, 1, days −1, 0, and day 0 windows of the net effect of all seven events for portfolios weighted by market capitalization (MV), leverage (DTM) and foreign sales (FS) using the full sample. In Panels A to D, the time-series daily returns of the weighted portfolios over the test period are regressed on market portfolio returns proxied by FT All Shares Index and a set of event dummies representing the test window for each of the seven sampled events. The event dummies are set to one (−1) for announcements that are favorable (unfavorable) to mandatory IFRS adoption in the EU and zero otherwise. All *t*-statistics have been adjusted for heteroskedasticity.

5.3.2. Alternative control for industry effects

The analyses in Section 5.2 indicate that the association between Pr value and announcement returns is more pronounced when the former is estimated only with firm-specific characteristics, i.e., Model 1 (Eq. (4)). This relationship deteriorates once industry membership is incorporated into the Pr estimation. As mentioned earlier, it is possible that the industry effect is capturing factors correlated with governance, which is more important in determining IFRS adoption choice in Germany and, therefore, less transferable to the UK setting. To rule out the possibility that the announcement returns associated with the Pr value that we proxied by firm-specific characteristics are driven by industry effects, we address the influence of industry membership in the construction of weighted portfolios under the Sefcik and Thompson (1986) approach. To implement this, we add 25 industry dummies based on Worldscope classification to the F matrix in Eq. (8) so that the portfolio weights generated in W'_{Pr} of Eq. (9) account for information pertaining to the Pr value based on Model 1 that is orthogonal to industry effects. Table 5 panel B presents the market-reaction tests based on this alternative industry-control approach for the days −1, 0, 1 window across the seven events individually and the net effect. The results are broadly consistent with those in Table 4. The last column, showing a net risk-adjusted abnormal return of 0.593%, is significant at the 1% level. Thus, our short-run market-reaction test for H1A remains robust even when controlling for industry effects.

5.3.3. Sensitivity to individual events

The market-reaction study of this paper is similar to Comprix et al. (2003) and Armstrong et al. (2006), in that it relies on identifying events that changed the likelihood of mandatory IFRS in the EU. In this study we argue that expectations regarding mandatory IFRS adoption in the EU were formulated mainly in 1999 and 2000 based on the arguments we presented in Section 4.4. Comprix et al. (2003) use events from 2000 to 2002 but conclude that their empirical evidence shows that only those events that were early in this period had significant news content. In the intersect period of 2000 we identify the same events but we interpret the events where IOSCO endorsed IFRS conditionally (Events 2 and 3) as unfavorable, which is different in direction to Comprix et al. (2003). Our classification builds on comments in the Financial Times in the weeks around the conditional endorsement as described in Section 4.4.[19] We believe our identification and classifications of events is correct but acknowledge that some events are open to interpretation. To ensure the robustness of our results, we exclude each of the seven events in turn from our analyses. Table 5 panel C shows that regardless of which event is excluded, the net market reaction remains statistically significant to at least the 5% level. In other words, our findings in Table 4 are not driven by any individual event. Even when we exclude Events 2 and 3 simultaneously, the net market reaction remains significant at the 10% level (and at the 1% level when we control for industry effects). For brevity we do not tabulate these additional results.

5.3.4. Sensitivity to individual components

In this study we have chosen to characterize a firm by it's similarity to a German voluntary adopter by using the Pr value as a measure of similarity. Recall that the Pr estimated from Model 1 (Eq. (4)) consists of three components: i.e., market capitalization, leverage, and foreign sales. Our analyses in Table 3 show that these firm-specific characteristics are significantly associated with voluntary adoption in Germany. In using the Pr value, we essentially capture information contained in these components as well as their

[19] See Financial Times the 25th May 2000, p. 2, "Brussels' lost voice" for a more detailed discussion of the implication of events in this period.

relative weights in determining voluntary adoption. To determine whether the relationship between Pr value and announcement returns is dominated by an individual component, we construct weighted portfolios based on these firm-specific characteristics and conduct market-reaction tests using the Sefcik and Thompson (1986) approach. To implement this, we substitute the *Pr* vector in the *F* matrix of Eq. (8) with vectors of the values of market capitalization, leverage, and foreign sales. This results in three separate sets of portfolio weights in the *W* matrix of Eq. (9). Table 5 panel D, summarizes the results of the net abnormal returns across all events associated with each specific characteristic from the main model. Notice that the coefficients of market capitalization and foreign sales are positive and the coefficients for leverage are negative. Thus, the direction of the relationship between the net market reaction and each of these component variables in panel D of Table 5 is consistent with the direction of their association with voluntary non-local GAAP adoption in Table 3. This implies that all three components contribute to the overall effect captured by the Pr value in the expected direction.[20] In terms of statistical significance, the foreign-sales and market-capitalization components vary in degrees depending on the test windows. The leverage component is not significant in any test windows. The fact that the Pr value is more powerful than its individual components in capturing market reaction implies that it contains information associated with voluntary non-local GAAP adoption beyond these components and it is not driven by any of them individually.

In summary, the results of the short-term market reactions support H1A. Firms with a high willingness to adopt voluntarily experience a positive (negative) market reaction on days that increased (decreased) the likelihood of mandatory IFRS, although the strength of the evidence depends on assumptions made in connection to events that changed the likelihood of mandatory IFRS. As described earlier, the drawbacks to short-run market-reaction tests are the identification and classification of events. Thus, to increase the robustness of our conclusion, in the next section we test a parallel hypothesis, H1B, using a different methodology that does not rely on identifying specific events.

5.4. Long-run changes in the cost of equity

Table 6 presents descriptive statistics on the changes between the pre-(January 1996 to December 1998) and post-(October 2001 to September 2004) decision periods in the implied cost of equity and control variables used for the analysis of the long-term effect of mandatory IFRS.

The descriptive statistics reveals a general upward time trend in the implied cost of equity, since the changes from the pre to the post period are significantly positive when measured in both the Ohlson and Juettner-Nauroth (2005) abnormal earnings valuation model (ΔICE_{OJ}) and the Easton (2004) PEG valuation model (ΔICE_{PEG}). This time trend is consistent with the findings of Daske (2006) in a German sample, Lee, Ng, and Swaminathan (2004, Table 1) in G7 countries, and Botosan and Plumlee (2005, Fig. 1) in the US. Changes to the proxies for size (ΔMV) and profitability (ΔOPM) are insignificant. On the other hand, increases in book-to-market value (ΔBM) and debt-to-market value

[20] We also applied industry control based on 5.3.2 and obtained similar results.

Table 6
Descriptive statistics on changes in implied cost of capital and control variables

	ΔICE_{OJ}	ΔICE_{PEG}	ΔMV	ΔBM	ΔDTM	ΔSG	ΔOPM
Mean	0.028	0.036	0.004	0.304	0.156	-23.359	1.951
StDev	0.173	0.116	0.095	0.546	0.426	84.263	81.891
t-statistics (Mean)	3.505	6.721	0.912	12.058	7.931	-6.003	0.516
Observations	469						

This table present descriptive statistics on changes in the calculated implied cost of equity according to the Ohlson and Juettner-Nauroth (2005) abnormal earnings growth model (ΔICE_{OJ}) and the Easton (2004) PEG model (ΔICE_{PEG}). ΔMV is relative change in the natural logarithm of market value. ΔBM is the change book-to-market value. ΔDTM is the change in long-term debt to market value. ΔSG is the change in sales in growth. ΔOPM is the change to the operational margin. The changes are calculated as the difference in a 36-month median between the pre-announcement period (January 1996 to December 1998) and post-announcement period (October 2001 to September 2004).

(ΔDM) and decreases in sales growth (ΔSG) are significant, which indicate a decline in growth and a rise in borrowing between the pre- and post- decision periods.

Table 7 presents the results of the analysis on long-run changes in implied cost of equity subsequent to the decision to impose mandatory IFRS adoption across the EU.

We apply cross-sectional regressions (Eq. (15)) of changes in implied cost of equity capital (ΔICE_{OJ} or ΔICE_{PEG}) on the Pr, which is the counter-factual proxy for willingness to adopt IFRS controlled by relative changes in market value, changes in book-to-market value (ΔBM), changes in debt-to-market value (ΔDTM), changes in sales growth (ΔSG), and changes in operating profit margin (ΔOPM). The dependent variables and control variables are calculated as their difference in three-year median value between pre- and post-decision periods. The Pr value is calculated based on Models 1 and 2, where the former is based only on firm characteristics and industry dummies are added to the latter. Model 3 is excluded for brevity since its performance is weak as shown in Table 4. Panels A and B presents the results based on ΔICE_{OJ} and ΔICE_{PEG} respectively as the dependent variable.

In both panels A and B of Table 7, the intercepts of all regressions are significantly positive, which confirms the background upward trend of implied cost of equity capital (e.g., Daske, 2006). In panel A, the long-run changes in the implied cost of equity capital obtained from the Ohlson and Juettner-Nauroth (2005) abnormal earnings growth model have a significantly negative relationship with the Pr value estimated from Model 1, which is evidence in support of H1B. This relationship is significant both before and after the addition of control variables, indicating the robustness of our findings. Panel B shows that the coefficients for the Pr value in Model 1 are also significantly negative when using the Easton (2004) PEG valuation model. This result is also significant in both univariate and multivariate regressions, indicating our findings are robust even to changes in proxies of implied cost of equity. In general, the results in Table 7 demonstrate that UK firms that share similarities in foreign sales, leverage, and size with the German early adopters have significantly lower increases in the cost of capital between the pre- and post-decision periods. The results are not subsumed by control variables that correlate with the implied cost of equity. Since two of these control variables, leverage and size, are components of the Model 1 itself, this implies that the significant relationship between Pr and changes in implied cost of equity are not simply driven by changes in the values of these two

Table 7
Changes in implied cost of equity capital of UK firms conditional on their degree of similarity to German voluntary IFRS adopters following the decision to impose mandatory IFRS adoption in EU

	Intercept	Pr	ΔMV	ΔBM	ΔDTM	ΔSG	ΔOPM
Panel A: ΔICE_{OJ}							
Model 1	0.07903	−0.10642					
	[4.04]	[−3.01]					
	0.07382	−0.09826	−0.52616	−0.01809	0.03086	−0.00000	−0.00008
	[3.69]	[−3.08]	[−5.00]	[−0.87]	[2.74]	[−0.01]	[−0.93]
Model 2	0.03436	−0.01297					
	[2.78]	[−0.63]					
	0.03921	−0.00952	−0.52285	−0.01660	0.03436	0.00000	−0.00007
	[2.78]	[−0.52]	[−4.91]	[−0.78]	[2.96]	[0.01]	[−0.79]
Panel B: ΔICE_{PEG}							
Model 1	0.06867	−0.06867					
	[3.79]	[−2.10]					
	0.06272	−0.05876	−0.56048	−0.01881	0.02250	−0.00004	0.00004
	[3.49]	[−2.10]	[−5.66]	[−0.93]	[2.25]	[−0.49]	[0.47]
Model 2	0.03384	0.00368					
	[3.02]	[0.19]					
	0.03036	0.00787	−0.56029	−0.01862	0.02484	−0.00004	0.00004
	[2.90]	[0.48]	[−5.64]	[−0.91]	[2.42]	[−0.48]	[0.55]

This table presents the coefficient and *t*-statistics (in brackets) of cross-sectional regressions of changes in implied cost of equity on the degree of similarity to German voluntary adopters (Pr) controlled by changes in market value (ΔMV), changes in book-to-market value (ΔBM), changes in debt-to-market value (ΔDTM), changes in sales growth (ΔSG), and changes in operating profit margin (ΔOPM). The implied cost of equity is calculated based on the Ohlson and Juettner-Nauroth (2005) abnormal earnings valuation model (ΔICE_{OJ}) in panel A and the Easton (2004) PEG valuation model (ΔICE_{PEG}) in Panel B. The changes in implied cost of equity and control variables are calculated as the difference in a 36-month median between the pre-announcement period (January 1996 to December 1998) and post-announcement period (October 2001 to September 2004). The Pr value is calculated based on two models. Each model is estimated using a German sample where the dependent variable is set to one for German voluntary IFRS/US-GAAP adopters and zero otherwise and the independent variables are firm characteristics, which includes foreign sales, leverage, and size (Model 1) and both firm characteristics and industry dummies (Model 2). The *t*-statistics are adjusted for heteroskedasticity.

components between the pre- and post-decision periods.[21] Although Daske (2006) and Cuijpers and Buijink (2005) did not confirm a significant decrease in cost of equity subsequent to adoption, our study shows that UK firms sharing similar characteristics to German voluntary adopters experienced a significantly lower increase in their cost of equity subsequent to the decision of mandatory IFRS adoption across the EU. Higher Pr firms in the UK are, therefore, associated with higher economic benefits from this regulatory decision relative to their lower Pr counterparts. The elimination of self-selection bias in the

[21] As a further test of the robustness we also include changes in foreign sales to total sales between pre- and post-decision period as a control variable. This does not change the results. The *t*-statistics on the Pr coefficients are – 3.12 and – 2.41 when estimating the cost of capital according to the abnormal earnings growth and PEG models respectively. It is left out of the tabulated results as there is no theoretical or empirical evidence for a connection between foreign sales and the cost of capital.

mandatory setting we study could account for this difference with previous studies of voluntary accounting-policy-choice setting. Finally, in both panels A and B of Table 7, the Pr values based on Model 2 exhibit an insignificant relationship with long-run changes in implied cost of equity. This is consistent with our interpretation of Table 4 that the addition of industry dummies only adds noise to the analyses of the UK sample. It also strengthens the argument that industry effects may be more important in capturing the first-mover advantage in Germany than in the UK, perhaps because it is correlated with corporate governance structure that is more important in explaining IFRS adoption choice in Germany than the UK.

5.5. Robustness tests of long-run changes in cost of equity

To enhance the robustness of our findings in Table 7 we conduct additional robustness tests on the relationship between our counter-factual proxy for the willingness to adopt and the long-run changes in the cost of equity. First, we include industry dummies in our tests to control for possible correlation with industry membership. Second, we control for other disclosures than those presented in the annual and interim reports. If our results are robust we should find the strongest relationship among firms with the least other disclosures, as IFRS adoption is only related to the annual and interim reports. Third, we test whether the relationship is less pronounced among ADR firms that already indirectly comply with an international accounting regime. Finally, we test if this relationship only exists around the decision period (1999 and 2000) by reperforming the test just before this period.

5.5.1. Controlling for industry effects

Although we control for changes in factors generally known to be correlated with risk, it remains uncertain whether our regression model sufficiently controls for cross-sectional differences in firm characteristics. Many firm characteristics are correlated with industry membership and thus controlling for industry fixed effects enables us to test whether the relationship is independent of these. Table 8 panel A includes 25 industry dummies based on the major industry classification in Worldscope (the coefficients on industry dummies are not reported). The results show that the relationship is slightly stronger after controlling for the industry effect. Thus, the lack of control for industry effects is not responsible for the relationship between the implied cost of equity and the counter-factual proxy for willingness to adopt.

5.5.2. Disclosures unrelated to IFRS

IFRS is limited to disclosures in the annual and interim reports. Although these statements are among the most important a firm makes it is not the only way to communicate with the market. If other disclosures are substantial the annual and interim reports would become relatively less important. Therefore, the economic consequences of mandatory IFRS are expected to be less pronounced. Following Leuz (2003) and Lang and Lundholm (1996) we use analyst following as a proxy for the level of other disclosures. We partition the sample into two sub-samples, following Botosan (1997), and define low following as below the median of the sample (3.5 analysts) and high following as above the sample median. The number of analysts following the firm is defined as the average yearly following from 1998 to 2002 obtained from IBES.

Table 8
Robustness test of the association between the similarity to German voluntary IFRS adopters and the change to the implied cost of equity

	Intercept	Pr	Pr*ADR	ADR	ΔMV	ΔBM	ΔDTM	ΔSG	ΔOPM
Panel A: Including industry dummies									
Ohlson-Juettner model	0.12845	−0.13398			−0.51721	−0.02923	0.04082	0.00000	−0.00012
	[3.58]	[−3.59]			[−5.20]	[−1.42]	[3.13]	[0.05]	[−1.50]
PEG model	0.09178	−0.09286			−0.55763	−0.02952	0.03293	−0.00003	−0.00001
	[2.97]	[−2.84]			[−6.01]	[−1.47]	[2.79]	[−0.40]	[−0.19]
Panel B: High/low analyst following									
High analyst following (>3.5)									
Ohlson-Juettner model	0.05008	−0.05405			−0.70606	−0.03000	0.01290	−0.00025	0.00000
	[2.56]	[−1.72]			[−4.29]	[−1.11]	[1.24]	[−1.10]	[−0.01]
PEG model	0.04039	−0.03242			−0.68270	−0.03092	0.00835	−0.00051	−0.00027
	[2.38]	[−1.18]			[−4.46]	[−1.26]	[0.84]	[−2.79]	[−2.20]
Low analyst following (<3.5)									
Ohlson-Juettner model	0.08961	−0.14093			−0.42238	−0.00871	0.53605	0.00005	−0.00009
	[2.97]	[−2.77]			[−3.59]	[−0.30]	[2.22]	[0.43]	[−0.92]
PEG model	0.07295	−0.07641			−0.47291	−0.01539	0.04545	0.00002	0.00007
	[2.70]	[−1.79]			[−4.30]	[−0.52]	[2.09]	[0.28]	[0.94]
Panel C: Including ADR-listed firms									
Ohlson-Juettner model	0.07204	−0.09770	0.14883	−0.06532	−0.51255	−0.01409	0.03210	0.00000	−0.00009
	[3.65]	[−3.07]	[2.23]	[−1.96]	[−5.14]	[−0.74]	[3.06]	[−0.09]	[−1.33]

PEG model	0.06113	−0.05840	0.08346	−0.03484	−0.54636	−0.01438	0.02280	−0.00004	0.00003
	[3.44]	[−2.09]	[1.33]	[−1.09]	[−5.82]	[−0.78]	[2.45]	[−0.57]	[0.42]
Panel D: Alternative test period (PEG model)									
Main test	−0.01252	0.01993			−0.23026	−0.00337	0.02816	0.00005	−0.00214
	[−1.28]	[1.11]			[−4.96]	[−0.39]	[2.93]	[0.26]	[−2.06]
Including ADR-firms	−0.01199	0.01918	−0.04130	0.01764	−0.23389	−0.00326	0.02493	0.00006	−0.00208
	[−1.22]	[1.07]	[−1.54]	[1.28]	[−5.81]	[−0.47]	[3.32]	[0.29]	[−2.18]
High analyst following (>3.5)	−0.01533	0.02096			−0.18632	−0.01000	0.01940	0.00012	−0.00242
	[−1.87]	[1.38]			[−4.55]	[−1.11]	[1.79]	[0.91]	[−4.43]
Low analyst following (<3.5)	−0.00608	0.00475			−0.31891	0.00119	0.04178	−0.00012	−0.00187
	[0.14]	[0.14]			[−2.86]	[0.06]	[2.73]	[−0.26]	[−1.35]

This table presents the coefficient and *t*-statistics (in brackets) of cross-sectional regressions of changes in implied cost of equity capital on the degree of similarity to German voluntary adopters (Pr) controlled by changes in market capitalization (ΔMV), changes in book-to-market value (ΔBM), changes in debt-to-market value (ΔDTM), changes in sales growth (ΔSG), and changes in operating profit margin (ΔOPM). The implied cost of equity capital is calculated based on the Ohlson and Juettner-Nauroth (2003) abnormal earnings valuation model and the Easton (2004) PEG valuation model. The changes in implied cost of equity capital ΔICE_{OJ} and ΔICE_{PEG} are calculated as the difference in a 36-month median between the pre-announcement period (January 1996 to December 1998) and post-announcement period (October 2001 to September 2004). The Pr value is calculated based on Model 1 described in Table 3. The control variables are calculated as their difference in a three-year median value between pre-announcement and post-announcement periods. The *t*-statistics are adjusted for heteroskedasticity. Panel A further includes 25 industry dummies as classified by Worldscope's major industry classification. Panel B partitions the sample into a low analyst-following and high analyst-following group. High analyst following is defined as above the median of average yearly following from 1998–2002 and low following is the remaining firms. Panel C includes 43 firms cross-listed in the US. The cross-listed firms are assigned a dummy variable taking the value one if the firm is cross-listed in the US and zero otherwise (ADR). Pr*ADR is the interaction between Pr and ADR. Panel D presents the results of a regression of changes in the cost of capital on the Pr value and a set of control variables. The change in the cost of capital is calculated according to the PEG model as the difference between the 36-month median from January 1990 to December 1992 and the 36 months median from October 1995 to September 1998. The change in the control variables are calculated over the same period as the change in the cost of capital.

Table 8 panel B reports the findings. The absolute size of the coefficient on the counter-factual proxy for willingness to adopt is almost three times larger in the low analyst-following group than in the high analyst-following group using the Ohlson and Juettner-Nauroth (2005) abnormal earnings growth model to estimate the implied cost of equity, and twice as large using the Easton (2004) PEG model. In the low analyst-following group the relationship is significant at the 1% level (*p*-value 0.006), whereas it is only significant at the 10% level (*p*-value 0.087) in the high analyst-following group when estimating the implied cost of equity using the abnormal earnings growth model (the difference is significant at the 10% level). When the PEG model is used, the relationship is marginally significant (*p*-value 0.075) in the low analyst-following group and not significant (*p*-value 0.238) in the high following group (the difference is not significant at conventional levels). These results are consistent with the findings in Botosan (1997). The additional disclosure imposed by IFRS mainly benefits firms with low analyst following. The results are consistent with the relationship between long-run changes in the cost of capital and the counter-factual proxy for willingness to adopt being caused by changes to the annual and interim reports, which is associated with IFRS, and not with an omitted correlated variable.

5.5.3. Including a control group

Until now we have excluded cross-listed firms from the sample. The logic behind this approach is that firms that are cross-listed are under more scrutiny from foreign investors and analysts and therefore are not on a comparable basis with the rest of the UK sample. In the UK, the majority of cross-listed firms have ADR-listings in the US. These firms generally produce reconciliations of income and book value of equity to US-GAAP. That is to say, these firms are indirectly already complying with an international accounting regime. Leuz (2003) provides empirical evidence that the difference in economic consequences of voluntary adoption of IFRS or US-GAAP is statistically insignificant.

As a robustness check, we test if the connection between our counter-factual proxy and the change in the implied cost of capital is indeed less pronounced among ADR-listed firms. If the relationship is equally significant among the ADR-listed firms, which already disclose earnings and book value of equity under an international accounting regime, this could imply that the connection arises by chance or is driven by some unobserved correlated variable and not mandatory adoption of IFRS. On the other hand, the observation that this relationship only exists among firms that do not have an ADR listing would mitigate these concerns. We identified 43 firms that are ADR-listed in the US stock market, and meet all data requirements described in Section 4.6. The number of ADR-listed firms is too small to partition the sample based on this variable but large enough for us to add them to the main sample with a dummy variable (ADR) and an interaction term between ADR and the counter-factual proxy for willingness to adopt (Pr*ADR). Table 8 panel C reports the results. The coefficient on Pr*ADR is positive both when estimating the implied cost of equity, using the abnormal-earnings growth and PEG models, but only significant in the former case. Regardless of estimation model, the relationship between the change in the cost of equity and the counter-factual proxy for ADR-listed firms (the coefficient on Pr and Pr*ADR together) is positive and insignificant (*p*-value 0.377 and 0.647, respectively). These results imply that the connection is not present in the ADR-listed control group and are consistent with the counter-factual proxy for willingness to adopt only explaining the change in the implied cost of equity when the firm does not already comply with an international accounting regime.

5.5.4. Non-decision making period

If the relationship between our counter-factual proxy of the UK firms' willingness to adopt and long-run changes in cost of equity exist by default and is not specific to the decision period we examine, then we would wrongly infer the results in Table 7 as a support for H1B. To mitigate this concern, we measure the long-run changes in cost of equity around a non-decision-making period and replicate the tests in Section 5.4. We use the latest possible period before the decision period (1999 and 2000) and restrict our tests to PEG estimates of the cost of equity to limit the loss of observations due to data availability. The results presented in Table 8 panel D show that the coefficient on the Pr value, contrary to expectations, is positive but insignificant (p-value 0.27). Including ADR firms as a control group does not change this result and the interaction term between the ADR dummy and the Pr value is insignificant (*p*-value 0.13). Finally, the coefficient on the Pr value is higher in absolute terms for firms with a high analyst following (although insignificant), inconsistent with the relationship being driven by changes to the annual and interim reports. Thus, replication of the test just prior to the decision period does not suggest that the relationship between Pr and changes to the cost of capital exists by default.

5.5.5. The technology bubble

Our overall study period spans from 1996 to 2004 during which the technology bubble occurs. Since this phenomenon is likely to be sector-specific, the observations that our results for the short-run market-reaction analyses (Table 5 panel B) and for the long-run changes in cost of equity-capital analyses (Table 8 panel A) remain robust after controlling for industry effect indicate that they are not driven by the bubble. In addition, the long-run change in cost of equity capital is computed from the pre-announcement period (January 1996 to December 1998) to the post-announcement period (October 2001 to September 2004), which is likely to exclude the peak period of this bubble. Finally, we also applied ADR-listed firms as a control group in the robustness check for the analyses of long-run changes in cost of equity capital (Table 8 panel C). Since the ADR control group is also likely to be affected by the bubble, it is unlikely that our findings would be driven by this particular phenomenon.

5.6. Implications

The inferences from both methodologies used in this study are consistent with significant differences across the population of UK quoted firms in the perceived net benefit of mandatory IFRS adoption.

These results highlight another dimension of the implication of IFRS adoption not explored in existing literature, which suggest either a reduced cost of capital (Leuz & Verrecchia, 2000) or no effect (Daske, 2006; Cuijpers & Buijink, 2005). Contrary to these views, we conclude that IFRS adoption has resulted in winners and losers. Rather than portraying IFRS as a uniformly good thing or a uniformly bad thing, it is important to recognize that some firms gain and some firms lose from complex, mandatory-accounting changes such as IFRS. This seems to us to make sense, because if all UK firms would have benefited from a regime like IFRS then it would have been adopted by the Accounting Standards Board (ASB) years ago.

Although our results do not imply that the cost of capital has been reduced or increased as a consequence of mandatory IFRS in general, the knowledge that significant differences exist among firms is important when considering costs and benefits to society. Implementing mandatory IFRS has the potential to redistribute wealth among agents in society through changes to the cost of capital. If the sole aim of the policy is to reduce the cost of capital, the best solution might be optional compliance as opposed to mandatory compliance. Optional compliance would allow firms to assess their own net benefits before committing to IFRS.

This paper also makes a methodological contribution by showing that the economic consequences of mandatory-accounting regulation in one economy may be partially predictable by using information contained in the accounting-policy commitments in a similar economy. This is a particularly interesting finding in the setting we use, because Germany and the UK differ in their approach to accounting regulation. Germany is generally classified as a code-law country and the UK is generally classified as a common-law country (Nobes, 2006). The fact that the same factors are significant determinants of benefits to international accounting harmonization suggests that benefits are driven by firm-specific characteristics rather than or maybe in combination with the quality of the national legal frameworks firms are departing from. Prior studies have focused on the quality of national regulation (Comprix et al., 2003). Future research in this area could examine how differences and links between country-specific and firm-specific factors affect the cost of capital changes caused by regulatory changes like mandatory IFRS adoption.

Finally, the finding that UK firms with a higher willingness to adopt the IFRS if given a chance are also those that would benefit from additional disclosure leads to several interesting questions. First, what is the disclosure quality of these firms? Are these firms that supply lower-quality information by choice or are these firms with increased demand for information due to growing investor interest? Second, if such firms knew that improvements in disclosure will benefit them, then what action did they take when voluntary adoption of IFRS was not allowed in the UK? To what extent are they allowed to voluntarily disclose more information under existing UK-GAAP? Could they have compensated with high-quality auditors? Finally, if UK-GAAP is assumed to be close to IFRS, then why should the decision of mandatory IFRS adoption incur economic consequences? Could our empirical evidence imply an overestimation of the degree of similarity between the two sets of standards? Alternatively, could any given differences between UK-GAAP and IFRS (although small relative to the difference between HGB and IFRS) have varying implications across UK firms? These questions are worthy of further study.[22]

6. Summary

In Germany, firms have had the option to comply with an international accounting regime (IFRS or US-GAAP) instead of domestic standards since 1998 and voluntary adoption is widespread. In the UK firms have not had this option and compliance with an international

[22] We thank Willem Buijink for suggesting them.

accounting regime is, therefore, very limited and only as a supplement to UK-GAAP. From 2005, IFRS is mandatory in both Germany and the UK as a consequence of EU regulation. We use this unique setting to create a counter-factual proxy for UK firms' willingness to adopt IFRS based on German firms' actual accounting-standard choices, and show that this proxy can predict cross-sectional variations in the economic consequences of mandatory IFRS adoption in the UK.

Using an event-study methodology, we find evidence that the stock-price reaction of UK firms to announcements favorable (unfavorable) to mandatory IFRS adoption is positively (negatively) related to our proxy for UK firms' willingness to adopt IFRS. To increase robustness we also study the long-run changes to the implied cost of equity of UK firms after the decision to mandate IFRS. We find that the change to the implied cost of equity is negatively related to our proxy for UK firms' willingness to adopt IFRS. Based on these two methodologies, we infer that cross-sectional variations in the economic consequences of mandatory IFRS adoption by UK firms can be predicted by their willingness to adopt IFRS proxied by the degree of similarity in characteristics with German voluntary IFRS and US-GAAP adopters.

Thus, mandatory IFRS has a different effect on the cost of capital depending on firm characteristics. Firms with similar characteristics to German voluntary adopters have greater benefits from international accounting harmonization and in particular from mandatory IFRS adoption.

This study also provides evidence on the information contained in firms' accounting policy commitments. We show that commitments made in one country can be used to predict the economic consequences of mandatory regulation in another country. Of course, some determinants may be less transferable from Germany to the UK, given the fact that the two countries investigated differ in their approach to accounting regulation, with the UK's common-law regulation being more similar to IFRS than Germany's code-law regulation. Thus, whereas the prior literature generally argues that relative reductions in cost of capital is related to the quality improvements in the legal framework, this study suggests that relative benefits are at least partly explained by firm-specific factors.

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