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Earnings Conservatism: Panel Data Evidence from the European Union and the United States

This paper focuses on earnings conservatism, and provides new evidence based on procedures that account for variability at the firm level, drawing a comparison between the European Union and the United States. A key finding is that the estimated responsiveness of earnings to bad news is substantially higher when unobserved firm-specific effects are modelled. Furthermore, it is shown that accounting has become more conservative not only in the U.S. but also in the EU when taken as a whole, and there is little evidence of marked differences in the asymmetric timeliness of earnings between the two. Indeed, any changes in this property of earnings are likely to be attributable to a common factor that influences firms similarly in both locations, and not necessarily to the process of economic convergence that has taken place in the EU.

Key Words: Accounting; Conservatism; Earnings properties; Panel data.

This paper focuses on earnings conservatism, measured in terms of the asymmetric responsiveness of net income to the sign of market returns, and provides new evidence based on procedures that account for variability at the firm level. A comparison is drawn between two major economic blocs, the European Union and the United States. A key finding is that the estimated responsiveness of earnings to bad news (negative returns) is substantially higher when unobserved firm-specific effects are modelled. Furthermore, it is shown that accounting has become more conservative not only in the U.S. as reported elsewhere (Givoly and Hayn, 2000) but also in the EU when taken as a whole, and there is little evidence of marked differences in the asymmetric timeliness of earnings between the two. Indeed, any changes in this property of earnings are likely to be attributable to a

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common factor that influences firms similarly in both locations, and not necessarily to the process of economic convergence that has taken place in the EU.

The notion of earnings conservatism has received considerable attention recently in the empirical research literature, even though conservatism as a general term has a long tradition in accounting. Earnings conservatism can be seen as a behavioural rule for accountants, although it is far from being a precise rule. In the past, the notion of conservatism implied the lower of cost or market, as well as recording future losses but not future profits. As Solomons (1986, p. 100) points out, 'too often in the past conservatism has been an excuse for attempts to present too consistently gloomy a picture of an enterprise's financial position and earnings by, for example, undervaluing inventory or over providing for deferred maintenance'. In general, conservatism now implies cautiousness in dealing with uncertainty. In the IASB's (2002, p. F.16, para. 37) conceptual framework, the notion of conservatism is defined as 'the inclusion of a degree of caution in the exercise of the judgments needed in making the estimates required under conditions of uncertainty, such that assets or income are not overstated and liabilities and expenses are not understated. However, the exercise of prudence does not allow, for example, the creation of hidden reserves or excessive provisions.' Therefore, the notion of conservatism can be seen as a corrective to the over-optimistic behaviour of managers that has to be applied when exercising judgments.

In recent years, market-based accounting research on conservatism has focused on the timeliness with which accounting earnings reflect economic income, following the approach proposed by Basu (1997). The use of changes in prices as the benchmark against which to appreciate the degree of conservatism of accounting earnings has an intuitive appeal as long as the market captures more information than the accounting system. Since the seminal work by Ball and Brown (1968), it has been shown that prices lead earnings, or earnings lag prices, but Basu provided evidence on an asymmetry in annual accounting earnings, in that firms incorporate bad news (a fall in equity prices over the year) faster than good news (an increase in equity prices) in their earnings.

Basu's (1997) work has been extended both in theoretical and empirical terms. Empirically, the notion of earnings conservatism has been examined for companies operating in different parts of the world, while, theoretically, other studies have attempted to build a conceptual framework to describe the factors that influence asymmetry in timeliness. In summarizing the main factors that provide incentives to the managers of firms and their accountants to report bad news on a more timely basis than good news, Watts (2003) offers four explanations: contracting, litigation, taxation and regulation. To these four causes of earnings conservatism we would add another: the economic driver. Bushman and Piotroski (2006) use the term 'political economy' to imply government intervention in markets. In more advanced capital markets, however, such intervention is minimal, and we suggest that this additional cause of the asymmetric timeliness of earnings is attributable more generally to political and macroeconomic developments that might affect the microeconomics of the firm.

Often, it is assumed that the conditions underlying such determinants of corporate behaviour are dissimilar across countries, the nation state therefore providing the main level of analysis, with observations on each firm treated as the basic

units of analysis within country sub-samples. Consequently, international research in accounting is usually structured so that certain factors influencing firms are held constant within countries but are allowed to differ between countries. The European Union presents a dilemma in this respect. The longstanding member states of the EU have engaged in a process of political, economic, financial and regulatory integration over several decades. Yet in many areas of empirical research that take an international perspective, there appears to be a preference still to focus on the individual member states, or to regroup them on the basis of the historical origins of the laws and regulations in force. In our view, the only economically sensible approach would be to assess the proximity of the drivers of earnings conservatism, and in this respect we consider the EU as cohesive. For reference, however, we also report on the estimation of earnings conservatism separately by member state. At the same time, we test for earnings conservatism in the U.S. market in order to compare the two major economic zones, the EU and the U.S., which are more comparable in political and macroeconomic terms.

A further contribution of this paper is the employment of a panel data methodology in an effort to provide more reliable estimates by focusing on firm-specific heterogeneous variability within the framework outlined above. The pooled ordinary least squares estimator ignores the panel structure of the data and treats observations as being serially uncorrelated for a given firm, with homoskedastic errors across firms and time periods. Here, we show why this is an entirely untenable assumption for a dataset that comprises firm-year observations.

EARNINGS CONSERVATISM

Using a reverse regression, Basu (1997) provides the seminal assessment of earnings conservatism, providing evidence in the U.S. to support his prediction that bad news is reported on a more timely basis than good news. Pope and Walker (1999) point to differences in asymmetric timeliness between firms based in the U.K. and the U.S., but then show that these are attributable to the components of earnings that are examined.[1] In contrast, Ball *et al.* (2000) argue that earnings conservatism is likely to differ more fundamentally between certain jurisdictions, and will depend more on the political influence on accounting practices that varies in line with the type of legal system within which the firm operates. In their view, a stakeholder governance model that solves information asymmetries using insider information will weaken the demand for timely and conservative accounting earnings. These authors compare the *R*-squareds of country-specific tests of the Basu model, and conclude that tests of code law countries (with stakeholder governance) provide a lower *R*-squared than those of common law countries (with shareholder governance).[2]

[1] Pope and Walker (1999) present evidence that U.K. firms differ from their U.S. counterparts because, in the time period examined, they use 'extraordinary items' to smooth ordinary earnings.

[2] The results reported by Ball *et al.* (2000) do not completely support their initial hypothesis as, for example, the *R*-squared of the test for France (one of the three countries branded as code law) is higher than for Australia (common law) and similar to the U.K. (common law).

While the theoretical discussion appears to be valid in terms of relating the determinants of earnings conservatism with regulation, shareholder litigation and private debt, the categorization of accounting environments based on legal origins that date back to the Roman Empire or preferences for debt financing that date back to the nineteenth century seems at the least problematic.[3] For a larger set of countries, Bushman and Piotroski (2006) re-examine the influence of legal and political institutions on the timeliness of the reporting of earnings. These authors suggest that earnings conservatism is indeed higher in countries with greater judicial quality, while strong public enforcement of securities law further delays the recognition of good news. However, they once again assign countries to different legal origin subgroups and reach their conclusions based on these generalizations.

Giner and Rees (2001) ask to what extent the conclusions reached in such studies apply to three EU member states that are representative of different accounting traditions: France, Germany, and the U.K. In order to capture prior news, they incorporate in their analysis previous earnings as well as prior period news. Overall, they find that, in the three countries involved, the contemporaneous association between earnings and returns is much stronger for bad news than for good news but that the inter-country differences are not statistically significant.[4] The effect of different domicile on the level of earnings conservatism is further examined by Raonic *et al.* (2004), who show that firm-level characteristics are more important drivers, and particularly that European companies that are inter-listed tend to report losses on a much timelier basis regardless of their location. On similar grounds, Huijgen and Lubberink (2005), using a matched-pairs research design, suggest that U.K. firms cross-listed in the U.S. report their earnings in a much more conservative way than U.K. firms listed only in the U.K.

While the general conclusion of this research is simply that earnings are timelier in reflecting decreases rather than increases in prices, the results are not consistent with differences in behaviour from one location to another. Instead, there is evidence that conservatism related to news is more likely to vary at the firm level. Indeed, it should be recognized that the higher is balance sheet conservatism as proxied by the market-to-book ratio, the lower the degree of earnings conservatism as measured by asymmetry (García Lara and Mora, 2004). These two types of conservatism are referred to respectively as 'unconditional' and 'conditional' by Beaver and Ryan (2005), who provide a comprehensive model of the firm's accounting that captures the interaction between them.

Turning to the factors that seem to drive conservative accounting, as summarized by Watts (2003), it is difficult to derive clear-cut predictions of their influence at the firm level that will be systematic across the EU and the U.S., the two economic

[3] McLeay (2005, p. 730) refers to such claims as a factoid that 'has the appearance of being reliable and accurate [just] because it has been repeated so often that people assume it is true'. Whitehead (1925) described this as the fallacy of 'misplaced concreteness'.

[4] Garcia Lara *et al.* (2005) reconsider earnings conservatism in the same three countries and relate it to earnings management practices, suggesting that French and German firms are more likely to incorporate bad news through discretionary accruals, whereas U.K. firms do not.

blocs that are juxtaposed here. On the one hand, consider the contracting explanation of conservative accounting, where published earnings are seen to signal the management's performance and thus affect their welfare either directly (salaries, bonuses) or indirectly (permanency), or both. It is notable that the average annual remuneration and bonuses of chief executive officers appear to be roughly the same across EU member states, but lower than in the U.S.[5] Thus one would expect a similar effect on the level of earnings conservatism throughout the EU, but less than in the U.S. On the other hand, consider the argument that managers might seek to calculate earnings in a manner that reduces the present value of taxes, leading to greater conservatism in financial reporting. In the EU, the interconnections between accounting and taxation, although now converging through recent efforts towards tax harmonization, started from a position of complete alignment between reported and taxable profits in many jurisdictions, unlike the U.S. Moreover, the tax burden in the EU appears to be higher than in the U.S., and, as Bushman and Piotroski (2006) suggest, steeper taxation gives rise to political costs that can influence observed reporting practices. Consequently, given these circumstances, one could expect higher conservatism in the EU than in the U.S. in this case, rather than the lower conservatism first presented. Similarly inconclusive arguments may be advanced in the case of the litigation and regulation explanations of conservatism.

However, we would also argue that an additional cause of the asymmetric timeliness of earnings comes from the impact of political and macroeconomic developments on microeconomic aspects of the firm. One of the most direct effects would be through the influence of interest rates on the firm's cost of capital. Changes in the cost of debt will affect the firm's choice between equity and debt financing, making borrowing more or less expensive. More expensive borrowing might turn the firm's management to further equity financing which in turn will augment the necessity for timelier recognition of news, and in particular for bad news to be reflected in earnings on an even timelier basis in jurisdictions with stronger shareholder protection (e.g., Ball *et al.*, 2000).

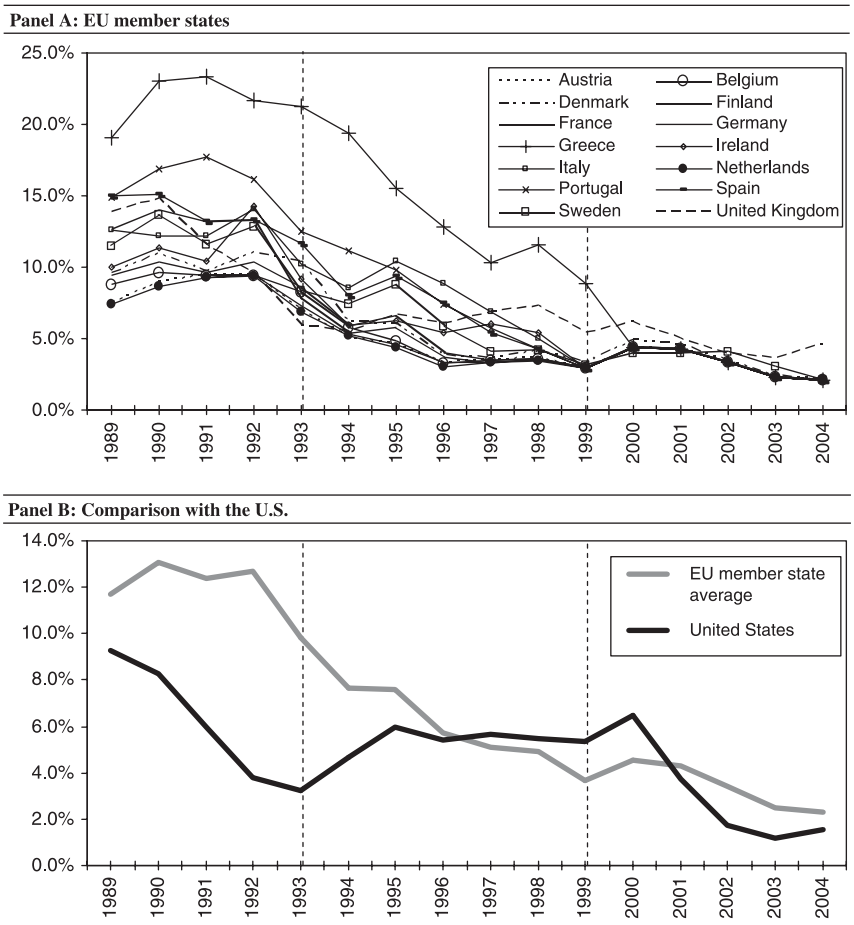
After the Maastricht Treaty in 1992, a convergence process was initiated in the EU with all member states aiming to fulfil the agreed macroeconomic criteria.[6] Figure 1, which shows the evolution of short-term interest rates, provides an illustration in Panel A of the degree of convergence that occurred in the EU prior to the locking of the euro in 1999. Such convergence would have direct effects on the cost of capital, leading to greater similarity in financial reporting behaviour in

[5] According to a Towers Perrin Report (2002), for the period 2001–02 the average annual remuneration appears to be very similar in the U.K. (668,526 U.S. dollars [USD]), Belgium (696,697 USD) and Italy (600,319 USD) and slightly lower in the rest of the EU (being in France at 519,060 USD and in Germany at 454,974 USD), while in the U.S. the relevant remuneration is triple (1,932,580 USD). In the same report, bonuses appear to be a considerably smaller part of the total CEO remuneration in EU than in U.S. companies.

[6] There are five main Maastricht convergence criteria: (a) inflation rates no more than 1.5 per cent above the three EU member states with the lowest inflation; (b) long-term interest rates no more than 2 per cent above the three EU member states with the lowest rates; (c) budget deficit under 3 per cent of the national GDP; (d) public debt ratio of less than 60 per cent of national GDP; and (e) at least two years of currency stability.

FIGURE 1

SHORT TERM INTEREST RATES



Note: Data were taken from OECD Country Statistics, www.oecd.com.

the EU. With regard to the comparison of the EU with the U.S., as shown in Panel B of Figure 1, it can be seen that the interest rate differential between the two remained substantial in the period prior to the Maastricht Treaty, with lower interest rates in the U.S. providing greater debt financing potential, which implies lower conservatism in the U.S. than the EU. However, that differential has almost disappeared since then, which would suggest that the effect on conservative reporting would no longer differ across the two economic blocs.

From the above, we can establish two main implications that should be analysed empirically in an effort to explain better the nature of conservative behaviour.

First, based on the drivers of earnings conservatism that have been identified, and due to the contradictory nature of the arguments concerning the timeliness of bad news, we would expect that the difference in earnings conservatism between U.S. firms and EU firms ought to decrease during the period examined. Second, due to the fall in interest rates and the associated increase in borrowing, we would predict a general decrease in conservatism during this time. Although we have argued that the economic integration achieved to date in the EU is consistent with the harmonization of drivers of earnings conservatism, we also recognize that other firm-specific factors should be taken into account when identifying the model to be used in the empirical analysis. This point is taken up again in the discussion of model construction that follows.

RESEARCH DESIGN

To model earnings conservatism, Basu (1997) introduced the following linear predictor with intercepts α_0 and α_1 and slopes β_0 and β_1:

$$E_{it} = \alpha_0 + \alpha_1 D_{it} + \beta_0 R_{it} + \beta_1 D_{it} R_{it} + u_{it}, \quad (1)$$

where E_{it} is earnings per share after extraordinary items scaled by the share price at the beginning of the period, R_{it} the discretely compounded annual return, D_{it} a dummy variable that takes the value of 1 if $R_{it} < 0$ (bad news) and the value of zero otherwise (good news), and $u_{it} \overset{iid}{\sim} N(0, \sigma^2)$ the residual term.

In contrast to previous research that has empirically assessed the extent of conservatism, this study uses panel data methods as they are more appropriate for the cross-sectional time series structure of the data. The pooled OLS estimator in equation (1) treats observations as being serially uncorrelated for a given firm, with homoskedastic errors across firms and time periods. Instead, the firm-specific panel specification of equation (1) for the cross-sectional unit $i = 1,2, \ldots ,n$ and firm-specific time operator $t_i = 1,2, \ldots ,T_i$ is:

$$E_{it} = \alpha_0 + \alpha_1 D_{it} + \beta_0 R_{it} + \beta_1 D_{it} R_{it} + C_i + v_{it}, \quad (2)$$

where C_i represents the firm-specific effects for firm i over T_i, and $v_{it} \overset{iid}{\sim} N(0, \sigma^2)$ the remaining normal disturbance component.[7] That is, equation (2) implies the existence of firm-specific effects in the unexplained error term u_{it} in equation (1), so that $u_{it} = C_i + v_{it}$.[8]

[7] T_i corrects the finite-sample bias that is introduced when unbalanced and unequally spaced panels are scaled by T. Thus, T_i is the length of the firm-specific time series and nT_i is the total number of observations for the entire dataset given its unbalanced panel structure.

[8] For a more detailed analysis of panel modelling, see Nerlove (2002), Arellano (2003), Hsiao (2003) and Baltagi (2005). Conventionally, the panel-specific effects denoted C_i include factors that are both unobserved and unobservable, which in the context of the present study would be firm-specific. Unobserved firm-specific factors include characteristics of a company that could be allowed for in the model as explanatory variables, such as the quality of management, but which are not added as they would introduce additional complexity to the modelling. Unobservable firm-specific factors represent those that are impossible or extremely difficult to measure, e.g., invisible assets such as technical know-how and brand awareness (Itami and Roehl, 1987).

If equation (2) is a meaningful and well-specified model, then it holds that:

$$\overline{E}_i = \alpha_0 + \alpha_1 \overline{D}_i + \beta_0 \overline{R}_i + \beta_1 \overline{D_i R_i} + C_i + \overline{v}_i, \tag{3}$$

where $\overline{E}_i$, $\overline{D}_i$, $\overline{D_i R_i}$, $\overline{R}_i$ and $\overline{v}_i$ indicate the respective mean values for each firm panel scaled by T_i. Equation (3) is the between-effects (BE) estimator that describes the mean variation between firms by assuming time-invariant effects.[9] To estimate equation (3), we use an analytically weighted least squares (A-WLS) estimator to account for the unequally spaced and generally unbalanced panel-specific T_i. Analytical weights $\sqrt{T_i}$ are required in least squares regressions that use averages instead of individual observations (Johnston and DiNardo, 1997), and as a result, A-WLS provides consistent results with a more stable estimation for the variance of $\sigma_C^2/T_i + \sigma_v^2$. The subtraction of (3) from (2) gives:

$$(E_{it} - \overline{E}_i) = \alpha_1(D_{it} - \overline{D}_i) + \beta_0(R_{it} - \overline{R}_i) + \beta_1(D_{it}R_{it} - \overline{D_i R_i}) + (v_{it} - \overline{v}_i). \tag{4}$$

This is the basis for an estimator that models variation within panels by assuming firm-invariant effects. Notice that the panel-specific disturbances C_i are eliminated, as is the constant term α_0, and therefore steps must be taken to allow for an intercept. Given that, following equation (2), it also holds that:

$$\overline{\overline{E}} = \alpha_0 + \alpha_1 \overline{\overline{D}} + \beta_0 \overline{\overline{R}} + \beta_1 \overline{\overline{DR}} + \overline{C} + \overline{\overline{v}}, \tag{5}$$

where $\overline{\overline{E}}$, $\overline{\overline{D}}$, $\overline{\overline{DR}}$, $\overline{\overline{R}}$ and $\overline{\overline{v}}$, are the global means for the entire sample scaled by nT_i and $\overline{C}$ the mean of firm-specific effects scaled by n, then by adding equations (4) and (5) we derive the following 'within effects' (WE) estimator:

$$\begin{aligned}(E_{it} - \overline{E}_i + \overline{\overline{E}}) = \alpha_0 &+ \alpha_1(D_{it} - \overline{D}_i + \overline{\overline{D}}) + \beta_0(R_{it} - \overline{R}_i + \overline{\overline{R}}) \\ &+ \beta_1(D_{it}R_{it} - \overline{D_i R_i} + \overline{\overline{DR}}) + (v_{it} - \overline{v}_i + \overline{\overline{v}}) + \overline{C}.\end{aligned} \tag{6}$$

To estimate equation (6), we require the restriction that $\overline{C} = 0$ and the estimated $\hat{\alpha}_0$ will reflect the average value of the within-firm effects. Therefore, the only difference between equation (6) and equation (4) is the inclusion of the intercept coefficient since the addition of global means will not alter the value of the slope, as they are constant values.

Lastly, an alternative estimator is the random effects (RE) estimator. The estimator considered here was introduced by Swamy and Arora (1972) and further developed by Baltagi and Chang (1994) for use with unbalanced panels. It is a matrix-weighted average of the between-firm and within-firm estimators, so that earnings E_{it} and its covariates $\mathbf{X}_{it}$ are transformed into:

$$E'_{it} = E_{it} - \hat{\theta}_i \overline{E}_i + \overline{\overline{E}} \quad \text{and} \quad \mathbf{X}'_{it} = \mathbf{X}_{it} - \hat{\theta}_i \overline{\mathbf{X}}_i + \overline{\overline{\mathbf{X}}}, \quad \text{with} \quad \hat{\theta}_i = 1 - \sqrt{\frac{\hat{\sigma}_v^2}{T_i \hat{\sigma}_c^2 + \hat{\sigma}_v^2}}, \tag{7}$$

where $\hat{\sigma}_v^2$ and $\hat{\sigma}_C^2$ are variances of the idiosyncratic normal error term and the panel-specific error term, respectively. This means that, if $\hat{\sigma}_C^2 = 0$, then $\hat{\theta}_i = 0$, so that

[9] For BE, we assume that there is no correlation between the covariates of the model $\overline{\mathbf{X}}_i : \{\overline{D}_i, \overline{D_i R_i}, \overline{R}_i\}$ and the panel-specific error term, $E(\overline{\mathbf{X}}_i, C_i) = 0$. In addition, in order to differentiate between the constant term $\hat{\alpha}_0$ and the firm-specific effects C_i, the constraint $\mathbf{\Sigma} C_i = 0$ is imposed over all i's.

equation (2) can be successfully estimated by OLS. On the other hand, if $\hat{\sigma}_v^2 = 0$, then $\hat{\theta}_i = 1$, which attaches full weight to the within-firm effects estimator in equation (6). Compared to the other two estimators, RE does not require the assumption that the individual effects are uncorrelated with the explanatory variables, but it is important to note that the random effects estimator treats C_i as a random normal variable $N(0,\sigma^2)$.

For the empirical part, we will estimate equations (3), (6) and (7), and compare to the pooled OLS estimator in equation (1).

SAMPLE

The study compares the EU and the U.S. for the period from 1989 to 2004.[10] Within this period, we allow for structural breaks in 1993, when the Maastricht Treaty was implemented in all EU countries, and again in 1999, the year when the Eurozone currencies were locked into the euro. These events account for the two most significant economic and financial regulatory developments within the overall time period. Hence, for the purposes of analysis, the sample frame is divided into three time periods: 1989–92, 1993–98, and 1999–2004.

The analysis is based on the closing stock prices and reported net income after extraordinary items recorded in the Worldscope database.[11] All financial firms, utilities and public organizations have been excluded in arriving at the sample due to the different nature of their products and the highly regulated market environments in which they operate, leaving 1,217 U.S. firms and 6,070 EU firms in the original sample.[12]

In related research studies, it has been usual to attempt to obtain estimates by eliminating the upper and lower 1 per cent of the frequency distributions of both earnings and annual returns. However, in our view this arbitrary univariate approach to extreme values is not justifiable as the observations that are removed are not necessarily extreme and, secondly, the remaining data may suffer from masking and swamping problems. That is to say, one extremely leveraged observation may mask the appearance of another, or a small cluster of outliers may attract the mean and inflate the variance in such a manner that some other observations will

[10] The EU saw the accession of a number of countries in 2004. This paper focuses on the fifteen pre-enlargement member states, excepting Luxembourg for which there is insufficient data. Thus it includes observations from Austria, Belgium, Denmark, Finland, France, Germany, Greece, Ireland, Italy, the Netherlands, Portugal, Spain, Sweden and the U.K.

[11] Basu (1997) and Pope and Walker (1999) use earnings both before and after extraordinary items in their assessments of the degree of earnings conservatism, and report that their results are sensitive to the definition of earnings. Others, however, conclude that their results are robust to both definitions (Garcia Lara *et al.*, 2005). For this paper, following Giner and Rees (2001), we choose to employ the 'all-inclusive' figure of earnings after extraordinary items.

[12] Our sample is drawn from the same source as Leuz *et al.* (2003), who also use the Worldscope database. Their study covers the period 1990–99, with a sample of 29,903 firm-years for the EU and 3,792 firm-years for the U.S. For the present study covering the period 1989–2004, the initial sample comprises 47,570 firm-years for the EU and 14,260 firm-years for the U.S. In the latter case, these are firms that are listed on the U.S. exchanges with the largest market capitalization for the period under consideration (NYSE, NASDAQ, AMEX and CHICAGO).

TABLE 1

PANEL STRUCTURE

	Original sample		1989–92		1993–98		1999–2004	
	n	Firms	n_H	Firms_H	n_H	Firms_H	n_H	Firms_H
U.S.	14,260	1,217	2,286	616	5,051	1,004	6,479	1,149
EU	47,570	6,070	9,176	2,587	16,793	3,957	20,579	4,659
Austria	864	114	154	45	316	78	367	81
Belgium	1,157	136	207	57	373	81	495	101
Denmark	1,628	170	325	92	576	129	622	126
Finland	1,258	150	169	53	375	87	591	118
France	7,537	1,028	1,409	409	2,411	601	3,393	758
Germany	6,399	808	979	280	2,034	425	3,117	702
Greece	1,585	272	63	22	483	133	896	244
Ireland	743	88	144	43	226	48	271	61
Italy	1,935	248	452	117	578	141	806	176
Netherlands	2,138	238	446	130	768	157	844	184
Portugal	695	90	101	29	243	61	265	63
Spain	1,380	153	242	71	476	103	506	106
Sweden	2,389	371	247	79	701	184	1,339	296
U.K.	17,862	2,204	4,081	1,109	6,434	1,473	6,827	1,584
Total sample	61,830	7,286	11,474	3,205	20,945	4,681	27,161	5,825

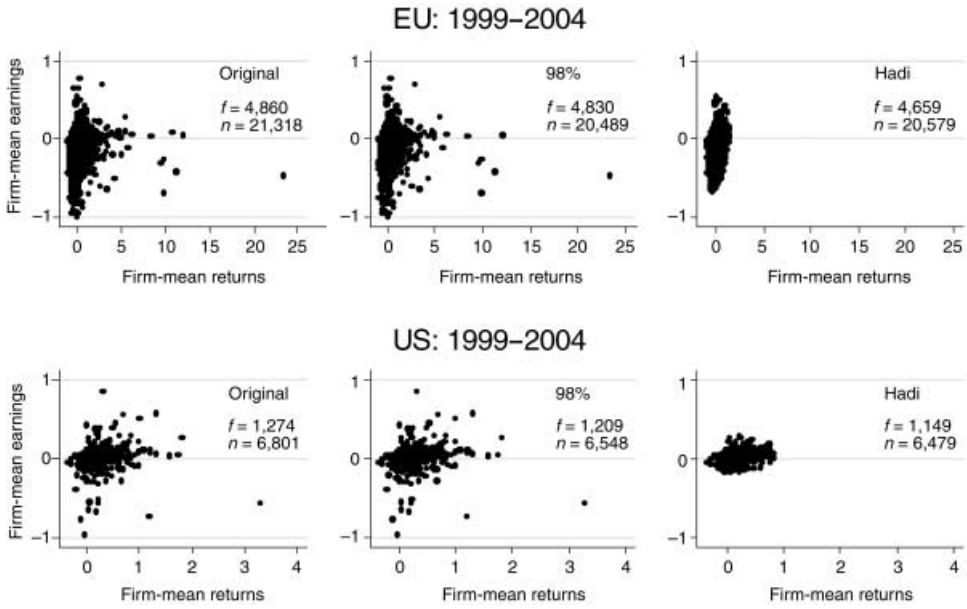
Note: n indicates the total number of cross-sectional time series (firm-year) observations before removing extreme values and Firms the number of firm-panels within the sample. n_H and Firms_H indicate the firm-year observations and the number of firms, respectively, after applying the Hadi (1992, 1994) method for detecting multiple outliers at the 5% significance level in multivariate samples prior to each estimation.

appear as outliers, when in fact they are not. In addition, the univariate approach overlooks the association between variables and cannot guarantee robust multivariate estimates. Finally, and more important for this study, heuristic univariate outlier deletion is inconsistent with panel methods of analysis, as it would lead to discontinuity in firm panels, thus restricting lagged panel estimation. Therefore, outliers are eliminated on the basis of firm-mean behaviour in the multivariate sample. That is, if the mean annual return and mean net income for firm i and time t appear as a multivariate outlier, then the entire firm-panel is excluded. The filter is adapted from Hadi (1992, 1994) and is applied prior to each estimation at the 5 per cent significance level.[13] Table 1 gives the panel data structure before

[13] For this paper we apply the algorithm proposed by Hadi (1994) which modifies the original version in Hadi (1992). The modification is also discussed in Hadi and Simonoff (1993), who consider methods for detecting multiple outliers in multivariate linear (regression) models. For a more detailed understanding for the problem of identifying outlying observations, see Barnett and Lewis (1994) and Hadi (2006).

FIGURE 2

ELIMINATION OF 1% UPPER AND LOWER UNIVARIATE VARIABILITY VS DELETION OF HADI'S BIVARIATE FIRM-MEAN OUTLIERS



Note: The scattergraphs indicate the bivariate relationship between firm-mean earnings and firm-mean returns. For comparison reasons, their plotted region is confined to $-1 \leq$ firm-mean earnings ≤ 1. The left-hand graphs plot the original samples. The central graphs show the remaining firms in the middle 98% of the distribution, i.e., following the elimination of the univariate lower and upper 1% of variation for firm-year earnings and firm-year returns. The right-hand graphs indicate the remaining firms following the outlier filter proposed by Hadi (1992, 1994) for detecting multivariate outliers as applied on the firm-means. *f* indicates the number of firms (panels) and *n* the number of firm-year observations.

and after the application of the multivariate outlier procedure. It can be seen from this table how the number of firms (and firm-year observations) varies considerably across member states within the EU, with the U.K. providing by far the largest sample. There are also differences in the number of cases in each sub-period, partially explained by their differing lengths; this is especially the case in the first sub-period, where both the number of firms and observations are considerably lower than in the other two sub-periods.

The appropriateness of the suggested procedure is illustrated in Figure 2, where the univariate approach to eliminating the extreme 1 per cent of firm-year observations is contrasted with Hadi's multivariate filter applied to firm-means. The univariate approach ignores the association between variables, whereas Hadi's filter correctly recognizes those firms that either report extreme earnings, or have extreme returns, or both. In Figure 2, average earnings and average returns for

each firm are plotted before and after the two procedures, for the U.S. and EU samples in the third sub-period from 1999 to 2004. The robustness of Hadi's filter is visibly superior.[14]

MODEL SELECTION AND EMPIRICAL RESULTS

As indicated above, panel data methods may be used to model the firm-specific component of variability. With respect to the datasets analysed here, the justification for this approach is given in Table 2, in which we report the variation of earnings and annual returns in the pooled data, together with the cross-sectional contribution to variation. It can be seen from the standard deviation of the firm averages that there is substantial cross-sectional variation, and it is this unobserved component of the pooled error that is accounted for by adding the firm-specific effects C_i in equation (2) above. It may also be noted that earnings and returns, for both the EU sample and the U.S., exhibit much higher volatility in the second period (post-Maastricht and pre-euro) than in other periods. For instance, the standard deviation of pooled earnings in the U.S. for the period 1993–98 is 3.46 and 2.85 times higher than in 1989–92 and 1999–2004 respectively, and 5.32 and 5.43 times higher at the firm level. This also substantiates the choice of three separate sub-periods for analysis.

TABLE 2

ANALYSIS OF VARIANCE

		Earnings_{ti}		Returns_{ti}	
		Pooled	Cross-sectional	Pooled	Cross-sectional
U.S.	1989–92	0.1057	0.0770	0.5250	0.2752
	1993–98	0.3660	0.4098	0.8906	0.3926
	1999–2004	0.1284	0.0754	0.5876	0.2026
EU	1989–92	0.1776	0.1231	0.4331	0.2425
	1993–98	0.7640	0.5328	0.9160	0.7637
	1999–2004	0.2990	0.1805	0.8443	0.3969

Note: The table reports standard deviations, which are biased-corrected estimates for unbalanced panels. Estimation is based on samples free of outliers, after applying the method proposed by Hadi (1992, 1994). For each sample that is tabulated, the pooled estimate is the standard deviation of y_{it} from the global mean $\bar{\bar{y}}$, where the estimate of cross-sectional variability is the standard deviation of panel means $\bar{y}_{it}$.

[14] To demonstrate the effect of Hadi's filter, consider the scattergraphs in Figure 2 for the EU sample, which initially comprised 4,860 firms and 21,318 firm-year observations. Deleting the upper and lower 1 per cent from both earnings and returns removes 829 firm-years, with 30 firms being eliminated entirely from the sample while 645 others remain in the sample with shorter or discontinuous series. Hadi's bivariate filter applied to the firm-means eliminates 201 firms and only 739 firm-years.

TABLE 3

MODEL SELECTION

	Estimator	1989–92			1993–98			1999–2004		
		F/χ^2	R^2	BIC/$\hat{\theta}_i$	F/χ^2	R^2	BIC/$\hat{\theta}_i$	F/χ^2	R^2	BIC/$\hat{\theta}_i$
U.S.	OLS	59.7	7.28%	–3930.0	87.2	4.93%	3958.1	103.3	4.57%	–8474.9
	WE	24.0	4.13%	–5286.1	69.4	4.90%	52.3	31.2	1.73%	–11075.6
	BE	57.3	21.92%	–1613.3	36.0	9.71%	141.7	172.6	31.01%	–3072.1
	RE	*135.0*	7.16%	*0.33*	*238.1*	4.53%	*0.71*	*220.4*	4.52%	*0.23*
EU	OLS	383.2	11.14%	–6718.9	45.6	0.81%	38516.7	459.7	6.30%	7419.1
	WE	124.8	5.38%	–11312.0	20.6	0.48%	28070.1	106.5	1.97%	–254.8
	BE	344.8	28.59%	–4613.1	29.0	2.15%	6021.4	700.5	31.09%	–4728.4
	RE	*851.8*	11.11%	*0.20*	*102.1*	0.78%	*0.39*	*1325.0*	6.28%	*0.03*

Note: The table provides diagnostics that guide model selection, between the pooled ordinary least squares estimator OLS (equation 1), the weighted least squares between-effects estimator BE (equation 3), the least squares within-effects estimator WE (equation 6) and the generalized least squares random-effects estimator RE (equation 7). Estimation is based on samples free of outliers, after applying the method proposed by Hadi (1992, 1994). *F*-tests the joint significance of all covariates for OLS, WE and BE, and, since the finite properties of RE are not known, χ^2_k with k parameters tests the joint significance for RE. R^2 is the coefficient of determination for OLS, WE and BE, and the squared correlation between y_{it} and fitted values $\hat{y}_{it}$ for RE. BIC is Schwartz's Bayesian Information Criterion which is reported for OLS, WE and BE. In the case of RE, $\hat{\theta}_i$ is the median value of the weights placed on y_{it} during transformation; when $\hat{\theta}_i = 0$ then RE is equivalent to the pooled OLS estimator, and RE converges to the WE estimator as $\hat{\theta}_i \to 1$.

Selection Criteria and Tests for Panel Specification

To identify the most appropriate panel method, we compare the pooled OLS estimator of equation (1) to the BE estimator of equation (3), the WE estimator of equation (6) and the RE estimator of equation (7). The diagnostics are given in Table 3, which include an *F*-test of overall model significance, the *R*-squared coefficient of determination and the Schwartz Bayesian Information Criterion (BIC). Note that, in the case of RE, a chi-squared test of overall model significance is reported instead of the *F*-test, together with $\hat{\theta}_i$, the median of the weights that are placed on the variable transformation.

The reported *F*-test for the OLS, BE and WE models and the corresponding chi-squared test for RE find all estimators to be significant in all three periods for the EU and the U.S. A key finding in Table 3 is the high level of variability explained by the fitted BE model. The *R*-squareds from the BE estimator are considerably higher than from OLS, WE or RE estimations. This implies the existence of highly significant firm-specific effects.[15] On this basis, the OLS estimator is

[15] The presence of firm-specific effects is verified through additional tests not reported here. The null hypothesis that all C_i are equal to zero is rejected by a further *F*-test for all samples. Then, the modified Wald test proposed by Baum (2000), which examines the null of a constant variance across panels ($\sigma_i = \sigma$ for all $i = 1,2, \ldots ,n$), gives results that are overwhelmingly in favour of group-wise heteroskedasticity in the residuals.

rejected since its assumption of homoskedastic residuals for the pooled sample of firm-years clearly does not hold.

To examine the appropriateness of a random effects estimator, $\hat{\theta}_i = 0$ would point to an inadequate specification, in the sense that RE would offer nothing more than the simple pooled OLS estimator. On the other hand, as $\hat{\theta}_i \to 1$, the RE estimator converges to the WE estimator. However, although the results reported here show that $\hat{\theta}_i$ is non-zero, the estimates of the transformation parameter are relatively low and additional tests allow us to reject a random effects specification.[16] Thus, the choice remains between BE and WE. The tests yield inconclusive evidence in this respect, as the BIC, although favouring WE, does not do so consistently. Moreover, the WE estimator yields mostly insignificant parameters, which do not assist the investigation of the null hypothesis of earnings conservatism.[17] This is not surprising, since the WE estimator attempts to fit the firm-specific one-way error-component model for three relatively short-term periods. Thus, the initial observation that the *R*-squareds from the BE estimator are considerably and consistently higher leads us finally to the conclusion that neither the RE nor the WE estimators can outperform the BE estimator. Thus, in the next section we proceed with our analysis by obtaining estimates solely from the BE estimation.

Testing for Structural Breaks in the Selected Model

As mentioned earlier, we assume that there may be structural breaks in the earnings-returns relation in 1993 (the implementation of the Maastricht Treaty) and 1999 (the locking of the euro), which imply dissimilar sets of parameters in the three sub-periods 1989–92, 1993–98 and 1999–2004. To investigate the validity of this hypothesis, we consider three separate specifications of equation (3), the BE estimator, one for each sub-period $s = 1, 2, 3$, as follows:

$$\overline{E}_{is} = d_{is}(\alpha_{0s} + \alpha_{1s}\overline{D}_{is} + \beta_{0s}\overline{R}_{is} + \beta_{1s}\overline{D_{is}R_{is}} + C_{is} + \overline{v}_{is}), \tag{8}$$

where $\overline{E}_{is}$, $\overline{D}_{is}$, $\overline{R}_{is}$ and $\overline{D_{is}R_{is}}$ are averaged with respect to s, and d_{is} is a dummy variable that is equal to 1 for sub-period s and 0 otherwise. To test the null hypothesis of no structural breaks, we examine whether the equalities $\alpha_{01} = \alpha_{02} = \alpha_{03}$, $\alpha_{11} = \alpha_{12} = \alpha_{13}$, $\beta_{01} = \beta_{02} = \beta_{03}$ and $\beta_{11} = \beta_{12} = \beta_{13}$ hold at the same time, and to do that, we first bring together the three specifications described by equation (8), and then estimate the pooled model by A-WLS so that we allow for its unbalanced T_{is} structure.[18] Thereafter, we may apply a conventional linear Wald test W to

[16] At the 5 per cent significance level, the RE estimator converges to the pooled OLS estimator in fifteen EU member state samples only. For the remaining samples (all 3 U.S., 3 EU samples, and at the EU member state level, the remaining 27 samples), the Hausman (1978) specification test finds RE to be inconsistent in 3 EU samples, 2 U.S. samples and 18 EU member state samples.

[17] The WE estimator finds only 19 significant β_1 and 38 significant α_0 coefficients in contrast to the BE which finds 41 significant β_1 and 48 α_0 coefficients, out of a total 51 sample estimations.

[18] The datasets are unbalanced both with respect to the different sub-period lengths (the first sub-period covers four years, and the second and third six years each) and, as mentioned previously with regard to equation (3), to variation in panel length.

TABLE 4

ESTIMATION OF EARNINGS CONSERVATISM IN THE U.S. AND EU

	Period	Firms	α_0	α_1	β_0	β_1	R^2
U.S.	1989–92	2,286	0.069***	0.021	0.040***	0.446***	21.92%
	1993–98	5,051	0.100***	0.177***	–0.044	1.381***	9.71%
	1999–2004	6,479	0.075***	0.039**	0.066***	0.592***	31.01%
EU	1989–92	9,176	0.096***	0.038***	0.062***	0.509***	28.59%
	1993–98	16,793	0.049**	0.147**	0.043*	0.919***	2.15%
	1999–2004	20,579	0.099***	–0.015	0.069***	0.626***	31.09%

Note: Firms indicate the number of firm-panels within the sample. α_0, α_1, β_0 and β_1 are the estimated coefficients from equation (3) and R^2 is the coefficient of determination. *** indicates significance at the 0.1%, ** at the 1% and * at the 5% levels.

examine the joint hypothesis of equal parameter coefficients across sub-periods.[19] As predicted, the results are strongly in favour of the inclusion of these particular structural breaks. For both the EU and the U.S. samples, *W* rejects the null with $p < 0.0001$ (*F*-statistics of 37.12 and 9.62, respectively).

Empirical Results

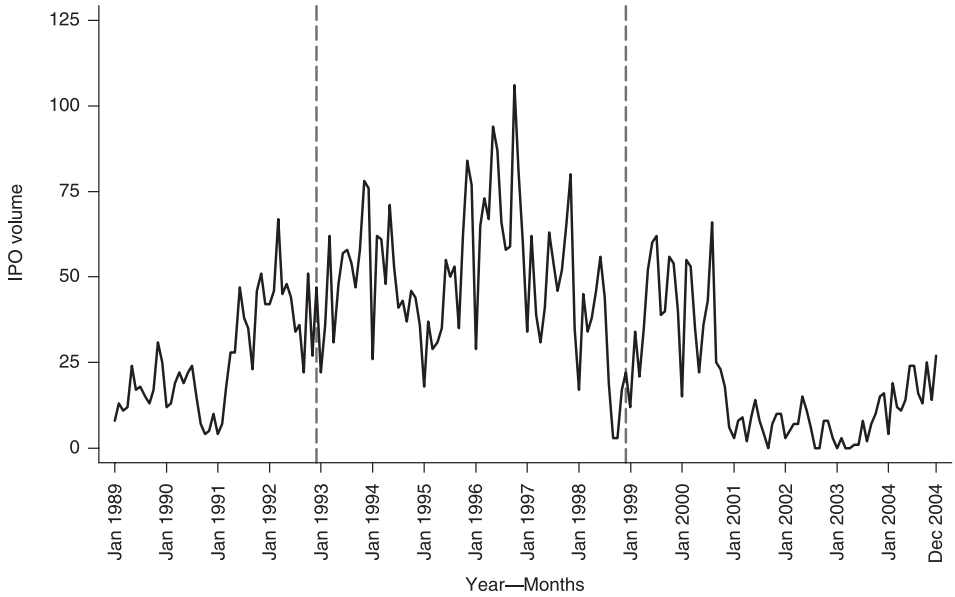
Table 4 presents the estimated coefficients of equation (3) for the period 1989–2004 by allowing for structural breaks in 1993 and 1999. The model describes the manner in which reported earnings incorporate the economic gains and economic losses of the firm during the period, using contemporaneous changes in shareholder value as a proxy, reflecting either the good news of a share price increase or the bad news of a share price decrease.

As can be observed in Table 4, the β_1 coefficients on the incremental effect of bad news are significant and high in all cases. In the EU, β_1 increases from 0.509 in the pre-Maastricht period to 0.919, and then drops to 0.626 after the locking of the euro. By comparison, the respective increase in the U.S. is from 0.446 to 1.381, then falling to 0.592. In general, it appears that firms in the EU and the U.S. operate under similar levels of earnings conservatism once we allow for firm-level variability. The relatively low values of β_0 in both the EU and the U.S. are to be expected as firms tend to be cautious when reacting to the appearance of good news about economic gains, not recognizing these in accounting income until they are eventually realized. At the same time, the coefficient α_0, which, as Pope and Walker (1999) explain, captures a proportion of permanent earnings, is positive and highly significant in each sample. This should not be surprising, since conservative

[19] The generalized Wald test *W* is detailed in Greene (2005, p. 486) and has an *F* distribution with $(k, n - 2k)$ degrees of freedom, with *k* parameters and *n* total observations for the stacked model. The use of a *W* type of test is justified, since the underlying distribution of the test-statistic remains unaltered even in the presence of structural breaks (Andrews, 1993).

FIGURE 3

VOLUME OF INITIAL PUBLIC OFFERINGS IN THE U.S. (1989–2004)



Note: The graph presents the number of IPOs per month for the U.S. The data are collected by Ibbotson *et al.* (1994) and from Jay Ritter's website: http://bear.cba.ufl.edu/ritter.

accounting points to a significantly positive α_0 that is proportional to the good news valued by the market in the past, the recognition of which was deferred to future periods in a gradual yet permanent manner.

The high levels of earnings conservatism in the second period may be related to the boom in stock prices during those years. Equity prices reached high levels and afterwards were considered to be overvalued both in the U.S. (see, e.g., Jensen, 2004) and in European countries (see, e.g., De Bie and De Haan, 2004). This situation had led many companies to raise more capital through the equity markets, as documented by Pastor and Veronesi (2005), and the increased level of equity financing is evident in the high volume of initial public offerings (IPOs) in the U.S., as shown in Figure 3. A broader equity base of this kind is likely to create more demand for timelier bad news recognition, especially in an economic zone with relatively strong shareholder protection, such as the U.S. Although the estimates of asymmetric timeliness fall from 1999 onwards, after the stock bubble, it is nevertheless the case that the conservatism level in the last sub-period examined is higher than in the initial sub-period, in both the U.S. and the EU.

Results at the member state level within the EU are provided in an Appendix, together with a comparison with previously published estimates obtained using

OLS methodology. It can be seen that, by allowing for firm-specific effects, the estimation with our sample gives generally larger β_1 coefficients than reported in the prior literature.[20] For instance, the significant β_1 OLS coefficients for the French sub-sample, which vary between 0.070 and 0.251 depending on the study that is cited, are noticeably lower than the between-effects estimates, which vary between 0.390 and 0.555 over the three sub-periods involved in our own study. On another issue, the comparison with previously published estimates in Panel B of the Appendix also shows how the range of returns (from minimum to maximum) differs considerably from one sample to another, which is consistent with the view advanced earlier that a univariate outlier deletion procedure that discards 1 per cent from each tail lacks robustness. Also, by comparison with other published results, which invariably have reported on OLS estimations, the *R*-squared from the between-effects estimator appears to be considerably higher.[21] Therefore, it may be concluded that the higher *R*-squareds from the BE estimation point to a better specified model, reflecting more of the variation in mean-returns contemporaneously when these are negative. An important conclusion of this study is that, once unobserved and unobservable firm-specific effects are accounted for, the responsiveness of earnings to bad news from the stock market is much higher in comparison to the aggregated pooled firm-year estimate.

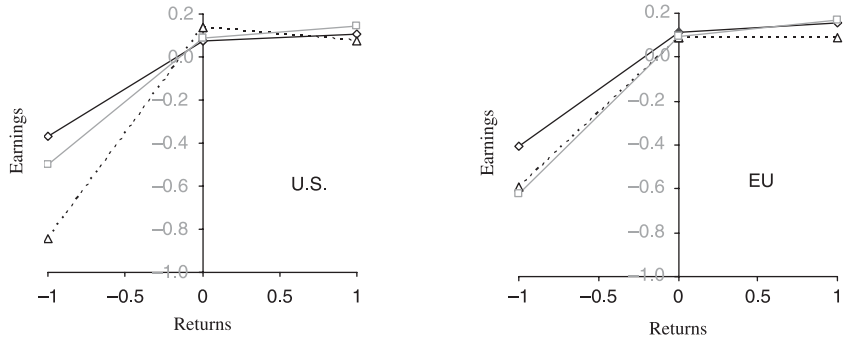
Finally, as for the α_1 coefficient, zero is shown in Panel A of the Appendix to be a plausible estimate across the member states of the EU or, in cases where the estimate is statistically significant, it is usually very low. This coefficient may be interpreted as the reversal of prior year market information in the light of current value changes, and for this reason we would expect it to be very close to zero. Support for this view can be seen in Raonic *et al.* (2004, p. 133), who find insignificant α_1, and attribute their result to the fact that 'the lack of timeliness in European financial reporting is manifested primarily in a tendency to delay the recognition of good news'. Similarly Giner and Rees (2001) do not find this coefficient significant in their analysis at the EU level. In a more analytical setting, Beaver and Ryan (2005) demonstrate that the reverse regression of earnings on current returns is strongly non-linear and passes through zero, proposing a polynomial approach to the problem that discards the use of dummy variables that are dichotomous on zero. Their suggestion implies the exclusion of those parameters that account for past market good news conveyed in current earnings (α_0) and the incremental past information that is contained in current bad news (α_1). However, while we would expect that $\alpha_1 = 0$, our results show that α_0 contains significant information.

[20] The Appendix shows details of the samples used in eight prior studies, together with the *R*-squareds and regression coefficients from OLS estimations. To assist comparison, OLS estimates are also given for the data used in the present study for the full period 1989–2004, deleting the upper and lower 1 per cent of the distribution in this case to be consistent.

[21] Although the between-effects *R*-squareds reported for the EU in Table 3 are relatively low in the second sub-period, they are still substantially higher than the OLS fit, the *R*-squared for each sub-period being as follows: 1989–1992 BE 28.59%, OLS 11.14%; 1993–1998 BE 2.15%, OLS 0.81%; and 1999–2004 BE 31.09%, OLS 6.30%.

FIGURE 4

A SIMPLIFIED REPRESENTATION OF EARNINGS CONSERVATISM IN THE U.S. AND EU



	Period	Firms	α_0	β_0	β_1	R^2
U.S.	1989–92	2,286	0.0747***	0.0333**	0.4099***	21.70%
	1993–98	5,051	0.1400***	–0.0644**	1.0463***	8.68%
	1999–2004	6,479	0.0868***	0.0543***	0.5351***	30.57%
EU	1989–92	9,176	0.1124***	0.0436***	0.4749***	28.27%
	1993–98	16,793	0.0903***	0.0301	0.6827***	1.91%
	1999–2004	20,579	0.0937***	0.0724***	0.6412***	31.08%

Note: Firms indicates the number of firm-panels within the sample. The solid black line indicates the period 1989–92, the dashed black line the period 1993–98 and the solid grey line the period 1999–2004. All insignificance coefficients (less than the 5% level) are set to zero in the graph. R^2 is the coefficient of determination *and* α_0, β_0 and β_1 are the coefficients from the between-effects estimation of the restricted model $\overline{E}_i = \alpha_0 + \beta_0 \overline{R}_i + \beta_1 \overline{D_i R_i} + C_i + \overline{v}_i$. *** indicates significance at the 0.1%, ** at the 1% and * at the 5% levels.

Given the above, the analysis proceeds by restricting the reverse regression of equation (3) so that it includes the constraint $\alpha_1 = 0$ and without constraint on α_0. In Figure 4 we show how the predicted change in earnings yield is affected by a unit stock price change, with linear predictors that are not discontinuous at zero (the estimated coefficients are tabulated below the graphs). The differences between the U.S. and the EU are seen to be relatively small. Indeed, by comparing the first and last periods, it can be seen that conservatism in the EU increased at a similar rate to the U.S.[22] What is more, at the firm-level, the β_1 coefficients appear more plausible than in previous studies on conservatism that have used OLS regressions (see Appendix). As firms are expected to react immediately to bad news, it is reasonable to expect $\beta_0 + \beta_1 \to 1$. This seems to be borne out in the findings, as $\beta_0 + \beta_1$ now tends towards the value of 1, especially in the second period in the U.S. ($\beta_0 + \beta_1 = 0.9819$).

[22] The reaction of U.S. firms to bad news for the first period in relation to the third period is equal to $(\beta_0 + \beta_1)_{1989–1992}/(\beta_0 + \beta_1)_{1999–2004} = 0.75$, which is a similar rate to the respective reaction of EU firms (about 0.73).

Finally, the α_0 coefficient remains positive and highly significant in the restricted estimation. More importantly, α_0 now appears to reflect the cost of capital, as modelled by Pope and Walker (1999). In particular, the evolution of the intercept tracks the movements in interest rates that were depicted earlier in Figure 1, suggesting an association with the cost of debt. It is evident that short-term interest rates were lower in the U.S. than in the EU for the first period, while for the second period the U.S. short-term interest rate moves to a higher level than in the EU, and both fall substantially in the third period to less than 6 per cent. Likewise, the component of earnings yield captured by α_0 rises from 7.47 per cent to 14.0 per cent and then falls to 8.68 per cent for the U.S. firms, and falls from 11.24 per cent initially in the EU to about 9 per cent thereafter.

CONCLUSION

This paper focuses on earnings conservatism and provides new evidence in the EU and the U.S., by using new procedures that account for variability at the firm level. Although we argue that the economic integration achieved to date in the EU is consistent with the harmonization of the drivers of earnings conservatism, we consider that other factors that are either unobservable or unobserved are by no means homogeneous, and they should be taken into account when establishing the empirical model to be used in the analysis. A key finding is that the estimated responsiveness of earnings to bad news is substantially higher when those firm-specific effects are modelled in comparison to the aggregated pooled firm-year estimate.

The analysis covers a period (1989–2004) in which several important economic events took place in the EU, from the signing of the Maastricht Treaty and the establishment of economic convergence criteria to the locking of most currencies into the euro. It is shown that accounting has become more conservative in the EU when taken as a whole, but this is also the case in the U.S., and furthermore there is little evidence now of marked differences in the asymmetric timeliness of earnings between the two. In our view, any change in this property of earnings is likely to be attributable to common economic factors that influence firms similarly in both locations.

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APPENDIX

Panel A: Between-effects estimation by EU member state using a multivariate outlier filter applied to firm-means

	1989–92					1993–98					1999–2004				
	α_0	α_1	β_0	β_1	R^2	α_0	α_1	β_0	β_1	R^2	α_0	α_1	β_0	β_1	R^2
Austria	0.082**	–0.036	0.025	0.196	0.271	0.071**	–0.007	0.074	0.161	0.121	0.163***	–0.135**	–0.024	0.406***	0.499
Belgium	0.080*	–0.018	0.127	0.079	0.196	0.109***	–0.052	–0.016	0.912***	0.276	0.104***	0.032	0.090*	0.588***	0.398
Denmark	0.070**	0.005	0.148**	0.264*	0.432	0.091***	0.007	0.046	0.434***	0.257	0.142***	–0.028	0.107*	0.741***	0.579
Finland	0.133*	–0.05	0.049	0.461**	0.368	0.106***	–0.024	0.028	0.108	0.121	0.121***	–0.059*	0.01	0.276***	0.424
France	0.111***	–0.001	0.062	0.390***	0.195	0.049**	0.031	0.118***	0.555***	0.229	0.079***	–0.02	0.166***	0.506***	0.287
Germany	0.040**	0.059*	0.044	0.497***	0.245	0.062***	–0.011	0.072**	0.529***	0.287	0.104***	–0.029	0.069*	0.618***	0.265
Greece	0.103**	–0.04	0.059	0.074	0.478	0.096***	–0.023	–0.013	0.224	0.072	0.077***	–0.022	0.001	0.113***	0.143
Ireland	0.147***	0.035	0.04	0.761***	0.642	0.097**	0.024	0.016	0.626***	0.418	0.113**	–0.022	–0.05	0.493**	0.271
Italy	0.104**	–0.005	0.078	0.409*	0.203	0.061	0.065	0.064	1.092***	0.303	0.091***	–0.034	0.102	0.484***	0.375
Netherlands	0.090***	0.033	0.139**	0.188	0.259	0.092***	0.003	0.058*	0.405**	0.193	0.141***	–0.092*	–0.03	0.415***	0.324
Portugal	0.173***	–0.087	0.017	0.219	0.363	0.145***	–0.053	–0.138*	0.756***	0.269	0.122	0.03	0.292*	1.324***	0.373
Spain	0.065	0.029	–0.061	0.395	0.024	0.110***	–0.032	0.013	0.665**	0.193	0.077**	0.016	0.087	0.279	0.098
Sweden	0.094***	0.006	–0.019	0.356***	0.374	0.074***	–0.048	0.101***	0.062	0.189	0.072***	–0.023	0.141***	0.560***	0.609
U.K.	0.092***	0.063***	0.095***	0.504***	0.376	0.086***	0.021	–0.004	0.588***	0.215	0.108***	–0.032	0.025	0.706***	0.341

Note: α_0, α_1, β_0 and β_1 are the estimated coefficients from the between-effects estimator of equation (3). R^2 is the coefficient of determination. *** indicates significance at the 0.1%, ** at the 1% and * at the 5% levels.

APPENDIX

(CONTINUED)

Panel B: OLS estimation with elimination of upper and lower 1% compared with prior published research

	Database	Period	Earnings	β_0	β_1	R^2	n	Earnings				Returns			
								Mean	*SD*	Min	Max	Mean	*SD*	Min	Max
Austria	**WS**	**89-04**	**A**	**0.019**	**0.159*****	**0.112**	**829**	**0.04**	**0.12**	**–0.57**	**0.38**	**0.05**	**0.43**	**–0.77**	**2.30**
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.069	–0.093	0.030	361								
Belgium	**WS**	**89-04**	**A**	**0.056*****	**0.213*****	**0.160**	**1,113**	**0.04**	**0.14**	**–0.82**	**0.47**	**0.06**	**0.39**	**–0.75**	**1.89**
García Lara and Mora (2004)	EX	87-04	A	0.081***	0.144	0.078	378	0.07	0.12	–0.54	0.83	0.14	0.36	–0.53	1.88
Raonic *et al.* (2004)	DS	87-99	B	0.054*	0.223**	0.091	200	0.09	0.12	–0.30	1.19	0.19	0.36	–0.67	1.58
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.029	0.181***	0.123	337								
Denmark	**WS**	**89-04**	**A**	**0.052*****	**0.225*****	**0.185**	**1,568**	**0.04**	**0.14**	**–0.81**	**0.54**	**0.08**	**0.43**	**–0.75**	**2.02**
Raonic *et al.* (2004)	DS	87-99	B	0.051***	0.439***	0.426	119	0.06	0.09	–0.29	0.30	0.17	0.45	–0.50	3.03
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.039	0.244***	0.154	498								
Finland	**WS**	**89-04**	**A**	**0.015**	**0.220*****	**0.136**	**1,210**	**0.05**	**0.14**	**–0.71**	**0.52**	**0.13**	**0.50**	**–0.80**	**2.59**
Raonic *et al.* (2004)	DS	87-99	B	0.018	0.135*	0.051	146	0.11	0.15	–0.25	0.89	0.30	0.70	–0.63	3.69
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.006	0.110*	0.072	428								
France	**WS**	**89-04**	**A**	**0.026*****	**0.219*****	**0.100**	**7,246**	**0.02**	**0.17**	**–1.20**	**0.48**	**0.08**	**0.47**	**–0.82**	**2.63**
Ball *et al.* (2000)	GV	85-95	B	0.080***	0.070*	0.126	1,054	0.06	0.11			0.15	0.43		
Giner and Rees (2001)	EX	90-98	A	0.044**	0.251***	0.134	1,425	0.04	0.12	–0.59	0.47	0.11	0.36	–0.61	1.65
García Lara and Mora (2004)	EX	87-04	A	0.052***	0.190***	0.132	1,313	0.05	0.13	–0.76	0.51	0.12	0.43	–0.67	1.98
Raonic *et al.* (2004)	DS	87-99	B	0.019**	0.031	0.063	555	0.06	0.08	–0.43	0.46	0.21	0.47	–0.85	3.56
García Lara *et al.* (2005)	DS	90-00	B	0.020*	0.160***	0.120	1,637	0.04	0.10	–0.95	0.36	0.08	0.43	–0.80	2.38
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.046***	0.040	0.064	2,439								

Germany	**WS**	**89-04**	**A**	**0.002**	**0.335*** **	**0.116**	**6,150**	**–0.01**	**0.20**	**–1.43**	**0.61**	**0.02**	**0.46**	**–0.88**	**2.38**
Ball *et al.* (2000)	GV	85-95	B	0.050***	0.100**	0.054	1,245	0.04	0.09			0.09	0.32		
Giner and Rees (2001)	EX	90-98	A	0.033	0.207***	0.090	1,391	0.02	0.10	–0.56	0.32	0.04	0.29	–0.60	1.24
García Lara and Mora (2004)	EX	87-04	A	0.032**	0.134***	0.079	1,702	0.03	0.10	–0.50	0.38	0.09	0.33	–0.55	1.64
Raonic *et al.* (2004)	DS	87-99	B	0.006	0.166*	0.056	297	0.06	0.06	–0.44	0.65	0.15	0.53	–0.62	7.26
García Lara *et al.* (2005)	DS	90-00	B	0.020	0.220***	0.090	3,245	0.02	0.12	–0.93	0.35	0.02	0.35	–0.81	2.36
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.035*	0.220***	0.108	3,056								
Greece	**WS**	**89-04**	**A**	**0.001**	**0.070*** **	**0.134**	**1,522**	**0.04**	**0.07**	**–0.39**	**0.28**	**0.48**	**2.09**	**–0.85**	**18.30**
Bushman and Piotroski (2006)	CS/GV	92-01	B	–0.002	0.087*	0.156	163								
Ireland	**WS**	**89-04**	**A**	**–0.015**	**0.229*** **	**0.086**	**713**	**0.03**	**0.14**	**–0.92**	**0.30**	**0.14**	**0.54**	**–0.77**	**3.64**
Raonic *et al.* (2004)	DS	87-99	B	–0.003	0.382*	0.277	136	0.08	0.09	–0.49	0.32	0.16	0.43	–0.62	1.90
Bushman and Piotroski (2006)	CS/GV	92-01	B	–0.031	0.495***	0.206	260								
Italy	**WS**	**89-04**	**A**	**0.039****	**0.199*** **	**0.101**	**1,857**	**0.01**	**0.15**	**–0.94**	**0.28**	**0.02**	**0.37**	**–0.70**	**1.79**
García Lara and Mora (2004)	EX	87-04	A	0.074**	0.081	0.067	624	0.04	0.15	–0.73	0.54	0.05	0.38	–0.62	2.03
Raonic *et al.* (2004)	DS	87-99	B	0.026	0.213***	0.061	269	0.06	0.13	–0.49	1.09	0.21	0.61	–0.60	5.13
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.005	0.135***	0.046	582								
Netherlands	**WS**	**89-04**	**A**	**–0.008**	**0.251*** **	**0.147**	**2,056**	**0.06**	**0.11**	**–0.61**	**0.29**	**0.07**	**0.39**	**–0.78**	**1.83**
García Lara and Mora (2004)	EX	87-04	A	0.054**	0.146**	0.157	680	0.09	0.10	–0.38	0.61	0.16	0.37	–0.60	1.47
Raonic *et al.* (2004)	DS	87-99	B	0.011	0.153**	0.103	397	0.09	0.08	–0.37	0.76	0.21	0.42	–0.71	2.17
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.009	0.177***	0.161	916								
Portugal	**WS**	**89-04**	**A**	**–0.106****	**0.425*** **	**0.054**	**668**	**0.00**	**0.26**	**–1.95**	**0.60**	**0.09**	**0.46**	**–0.71**	**2.50**
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.018	0.263	0.027	163								
Sweden	**WS**	**89-04**	**A**	**0.016***	**0.268*** **	**0.174**	**2,305**	**0.01**	**0.16**	**–0.83**	**0.39**	**0.11**	**0.58**	**–0.85**	**3.75**
Raonic *et al.* (2004)	DS	87-99	B	–0.042	0.180	0.061	187	0.16	0.22	–0.44	1.30	0.29	0.59	–0.59	2.78
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.043	0.486***	0.304	671								
Spain	**WS**	**89-04**	**A**	**0.042***	**0.255*** **	**0.052**	**1,330**	**0.04**	**0.20**	**–1.82**	**0.50**	**0.11**	**0.44**	**–0.69**	**2.40**
García Lara and Mora (2004)	EX	87-04	A	0.038	0.518**	0.077	419	0.02	0.35	–2.93	1.34	0.14	0.57	–0.83	0.57
Raonic *et al.* (2004)	DS	87-99	B	0.041***	0.091*	0.167	275	0.08	0.07	–0.26	0.46	0.27	0.61	–0.63	4.49
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.056**	0.314***	0.114	555								

APPENDIX

(CONTINUED)

	Database	Period	Earnings	β_0	β_1	R^2	n	Earnings				Returns			
								Mean	*SD*	Min	Max	Mean	*SD*	Min	Max
U.K.	**WS**	**89-04**	**A**	**–0.016*****	**0.298*****	**0.119**	**17,181**	**0.01**	**0.17**	**–1.25**	**0.32**	**0.05**	**0.50**	**–0.86**	**2.96**
Pope and Walker (1999)	CS	76-96	B	0.081***	0.165***	0.172	7,189	0.09	0.13	–1.31	1.18	0.16	0.47	–0.86	5.40
Pope and Walker (1999)	CS	76-96	A	0.031***	0.276***	0.104	7,189	0.08	0.16	–1.57	1.24				
Ball *et al.* (2000)	GV	85-95	B	0.040***	0.150***	0.138	5,758	0.07	0.09			0.14	0.38		
Giner and Rees (2001)	EX	90-98	A	0.027***	0.289***	0.151	6,022	0.04	0.14	–1.08	0.36	0.14	0.44	–0.75	2.29
García Lara and Mora (2004)	EX	87-04	A	0.014	0.243***	0.135	6,285	0.05	0.13	–0.90	0.38	0.08	0.39	–0.70	1.61
Raonic *et al.* (2004)	DS	87-99	B	0.010	0.111***	0.074	800	0.07	0.06	–0.43	0.45	0.17	0.43	–0.89	5.72
García Lara *et al.* (2005)	DS	90-00	B	0.020***	0.240***	0.140	10,131	0.04	0.14	–0.97	0.37	0.06	0.45	–0.82	2.39
Bushman and Piotroski (2006)	CS/GV	92-01	B	0.002	0.276***	0.105	7,596								
U.S.A.	**WS**	**89-04**	**A**	**0.012*****	**0.114*****	**0.057**	**13,739**	**0.05**	**0.08**	**–0.58**	**0.38**	**0.16**	**0.44**	**–0.71**	**2.32**
Basu (1997)	CS/CRSP	63-90	B	0.075***	0.166***	0.115	25,665								
Pope and Walker (1999)	CS	76-96	B	0.026***	0.282***	0.112	18,380	0.07	0.13	–2.00	0.64	0.14	0.39	–0.88	4.17
Pope and Walker (1999)	CS	76-96	A	0.077***	0.242***	0.118	18,380	0.07	0.14	–2.00	0.82				
Ball *et al.* (2000)	GV	85-95	B	0.030***	0.290***	0.147	21,225	0.03	0.14			0.13	0.43		
Bushman and Piotroski (2006)	CS/GV	92-01	B	–0.005	0.312***	0.114	26,776								
EU	**WS**	**89-04**	**A**	**0.003**	**0.263*****	**0.116**	**45,754**	**0.02**	**0.16**	**–1.17**	**0.41**	**0.07**	**0.50**	**–0.85**	**3.03**

Note: All studies included report OLS estimates, and eliminate the upper and lower 1% of variability from both earnings and returns. For comparison, we also provide in bold the OLS estimates obtained with the samples used in this study. Commercial databases: WS Worldscope, CS Compustat, GV Global Vantage, EX Extel, DS Datastream. Earnings definitions: A is after extraordinary items, and B is before extraordinary items. β_0 and β_1 are the estimated coefficients from the pooled OLS estimation of equation (1). R^2 is the coefficient of determination, and n is the number of firm-year observations in the sample. Basu (1997) and Bushman and Piotroski (2006) do not provide country-specific descriptive statistics. *** indicates significance at the 0.1%, ** at the 1% and * at the 5% levels.