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ARTICLE

The elusive relation between pension discount rates and deficits

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Abstract

The relation between defined-benefit (DB) pension discount rates and funding status is more complex than it might first appear. Existing evidence suffers from estimation biases that make precise inference unreliable. We document the biases and quantify their impact on inference in relation to corporate window-dressing of DB funding status. Our empirical evidence from the United Kingdom suggests that pension sponsors use discretion in the choice of pension discount rate not only to reduce reported deficits but also to reduce reported surpluses.

KEYWORDS

benchmark discount rate, defined benefit pension scheme, funding status, pension discount rate, window-dressing

JEL CLASSIFICATION

M49, M48, G39

1 | INTRODUCTION

Defined-benefit (DB) pension schemes have future obligations to members who are defined *ex-ante*. A key assumption in estimating their present value is the rate at which they are discounted. Accounting standards mandate using the yield on high-quality corporate bonds as a basis for the discount rate, but they leave room for discretion. There is evidence that sponsors apply this discretion opportunistically, using a higher discount rate to window-dress schemes with weaker underlying funding levels.

Empirical analysis of this form of window-dressing is complicated by two primary issues. First, we observe the *reported* funding status after any opportunistic discretion may have been employed. We are interested in the relation between the discretionary discount rate and underlying funding status *before* any discretion, by which we mean the difference between the scheme's reported discount rate and its benchmark discount rate. Given that the underlying funding status is unobservable, most existing studies proxy for it by adjusting the reported funding status. However, the adjustment methods commonly used introduce unintended bias. Second, International Accounting Standard (IAS) 19 requires that liabilities are discounted at the yield on high-quality corporate bonds with a similar currency and duration profile to the pension liabilities being valued. This is what we refer to as the benchmark discount rate. However, the standard does not prescribe a method for deriving a discount rate curve with sufficiently granular maturity intervals to facilitate exact duration matching between the liabilities and benchmark. Previous studies have tended to omit benchmark discount rates owing to this ambiguity and non-disclosure of duration before IAS 19 was introduced in 2013. Given that a scheme's benchmark rate affects both the reported discount rate and its unbiased funding status, the omission leads to an underestimation of the impact of funding status on the choice of a discount rate.

In this study, we document these biases (the scale of which are unknown a priori) and estimate their impact on inference. We do so first via a simulation exercise. This allows us to counterfactually observe the unbiased funding status and, crucially, to vary its impact on the choice of a discount rate. We analyze the performance of existing funding status adjustment methods under a range of window-dressing scenarios, scheme duration profiles and specifications, which include and omit the benchmark as a control. We propose an alternative adjustment that utilizes scheme-specific liability duration.[1] In addition, we present a method for deriving the pension discount rate curve to obtain estimates of benchmark rates. This method converts the existing Financial Times-Stock Exchnage (FTSE) dollar-denominated AA-rated pension yield curve into alternative currencies at granular maturity intervals and is thus consistent with IAS 19. We apply our adjustment method and determination of benchmark rates to investigate the use of discretionary discount rates by UK pension sponsors. The empirical sample provides the data upon which our simulations are calibrated.

Evidence from the simulations confirms that our duration-based adjustment to funding status, combined with controlling for the benchmark rate, results in direct and accurate estimation of the impact of underlying funding status on the choice of a discount rate. The superior accuracy of our adjustment method is robust to variation in both the degree of underlying relation between true funding status and the discount rate (hereafter referred to as underlying b) and scheme duration. In contrast, the biases resulting from existing estimation methods can be substantial and do not have a consistent sign. The sign depends on the underlying b, average scheme duration and whether the benchmark rate is controlled for. The largest source of bias is the downward bias stemming from the omission of the benchmark rate as a control variable. Since most existing studies omit the benchmark, our simulations suggest that scheme funding status has a larger impact on the discount rate than has previously been documented.

We apply our method to investigate the use of discretionary discount rates by UK companies over the period 2009–2018. All report under International Financial Reporting Standard (IFRS) and since the provision of DB pension schemes is well-established in the United Kingdom, we have a sizable sample of over 2500 scheme-years. We find a positive and significant relation between the discretionary discount rate and the unbiased deficit ratio (the estimated unbiased value of a scheme's pension liabilities divided by its pension assets). The relation is larger after controlling for the benchmark rate, as expected, and larger using our duration-based adjustment to the value of pension liabilities than using the two existing adjustments which dominate the literature. Across specifications that use the duration-based adjustment, the estimate of underlying b—the impact of the unbiased deficit ratio on the discount rate—ranges from 0.91 to 1.65. This is economically significant. For the average sampled scheme, this degree of manipulation reduces an increase in the scheme's reported deficit by 25% to 45% of the increase that would otherwise have

[1] Since 2013, pension sponsors reporting under IFRS have been required to report scheme duration in accordance with IAS 19. Duration is not reported under US GAAP. This has been an obstacle for researchers using US data who wish to estimate a scheme's benchmark rate.

been reported without any manipulation. The manipulation has the greatest impact on deficit ratios in the case of schemes with the most precarious funding levels.

However, a wish to understate deficits explains only a minority of sample companies' use of discretionary discount rates. We document variation in behaviour dependent on the relative health of the scheme. Nearly one quarter of scheme-years show a surplus, with the proportion increasing in recent years. In most of these cases, the discount rate is set below the benchmark rate, implying that pension liabilities are overstated. We find that underlying *b* is positive for schemes in surplus, as well as for those in deficit. This implies that companies with schemes in surplus follow a surplus-reduction policy; the understatement of the surplus increases with its size. This is a different type of window-dressing from the use of upward discretionary discount rates to understate deficits. Companies with schemes in modest deficit also tend to choose a rate below their benchmark rate. In such cases, the deficit is overstated rather than understated as would be expected if deficit reduction were the objective. Overall, a wish to understate pension liabilities only explains the behavior of companies with large deficits (specifically, with estimated unbiased deficits of at least 14% of pension assets).

The paper contributes to the literature on sponsor choice of assumptions in pension-fund valuation and reporting. The contributions are both methodological and empirical. We clarify the problems in estimating the impact of pension funding on companies' choice of a discount rate. This is worthwhile since the rate is a key variable in estimating the value of liabilities, companies have discretion in their choice of rate, and the choice has been widely studied. We present an improved methodology that can be implemented using scheme duration disclosed under IAS 19, and we show analytically and by simulation that it all but eliminates estimation error. Our method will be of benefit to other researchers interested in the window-dressing of pension schemes. We also show how to calculate a pension yield curve for the United Kingdom, which enables accurate estimation of scheme-specific benchmark rates, given scheme duration. Our technique should help analysts and regulators, who are often interested in measuring the discretionary discount rates used by sponsors in valuing individual schemes.

Our simulated evidence enables the biases from previous estimation methods to be assessed for the first time. It shows that estimates of underlying *b* can be seriously understated if the benchmark rate is omitted as a control variable, whatever method is used to de-bias reported liabilities. Our empirical finding that underlying *b* is positive is consistent with that of Billings et al. (2017) for an earlier UK sample, and with the majority of other studies, most of which are US-focused.[2] Although we arrive at the same qualitative conclusion as most prior studies, our methodological innovation enhances precision and leads to more robust inference. Our empirical results also show that a positive value of underlying *b* arises from the use by companies of discretionary rates to understate large surpluses, as well as to understate large deficits. The discretionary use of the discount rate to pursue a surplus-reduction policy is a novel contribution to the literature.

2 | THE DEFICIT-REDUCTION HYPOTHESIS AND PREVIOUS RESEARCH

A company's DB pension scheme affects its financial condition in several ways. The scheme's surplus or deficit is shown on the balance sheet under IAS 19. The value of its liabilities and funding status affect the company's annual pension expense, which is a deduction from its income, and its annual contribution to the scheme, which is a cash outflow. A stream of research examines how and why company managers might influence the above items, through exercising

[2] Papers that conclude there is a positive relation include Feldstein and Mørck (1983), Bodie et al. (1984), Thies and Sturrock (1988), Gopalakrishnan and Sugrue (1995), N. Godwin (1999), Fried and Davis-Friday (2013) and Billings et al. (2017). Papers that conclude there is no relation or a negative relation include Asthana (1999), Sweeting (2011), Jones (2013), and Bartram (2018).

discretion regarding cash contributions and key pension assumptions, including the discount rate for the scheme's liabilities, the expected rate of return on its assets and the future growth rate of employees' salaries.[3]

Managers of companies with schemes in deficit have several incentives to manipulate pension assumptions to reduce the reported deficit. A growing deficit implies increasing pension expense in the income statement, and larger cash contributions to the fund, since the company will be under pressure from the fund's trustees and regulator to reduce the deficit. Reported funding status affects the company's leverage calculated inclusive of pension assets and liabilities; a growing deficit tends toward an increasing cost of debt and a greater likelihood that debt covenants will be violated (e.g., Asthana, 1999; Bartram, 2016; Feldstein & Mørck, 1983; J. H. Godwin et al., 1996). Large funding deficits are viewed as a signal of poor managerial skills and are associated with lower management remuneration (Alderson et al., 2017).[4] For these reasons, managers might choose a higher discount rate to value future obligations. A higher rate implies a lower value, and therefore an improved funding status; it also normally reduces pension expense in the income statement.[5] The *deficit-reduction hypothesis* predicts that there is a positive relation between a company's discretionary discount rate (*DR*) and its scheme's unbiased deficit as a proportion of pension assets.[6,7]

Most of the evidence for the hypothesis comes from the United States. From the late 1990s, an increasing proportion of US DB pension schemes reported a deficit; by 2018, 93% of the 306 Standard & Poors (S&P) 500 companies with a DB scheme reported a deficit (authors' calculation). The majority of studies to date find a positive and significant relation between the discount rate and deficit as a proportion of pension assets, consistent with the hypothesis (Bodie et al., 1984; Feldstein & Mørck, 1983; N. Godwin, 1999; Gopalakrishnan & Sugrue, 1995; Thies & Sturrock, 1988; for the United Kingdom, Billings et al., 2017). Fried and Davis-Friday (2013) find that the relation became stronger after the announcement of Financial Accounting Standard (FAS) 158 in 2006, which required recognition of the pension surplus or deficit on the balance sheet, instead of in a note to the accounts. However, the relation is found to be insignificant in Asthana (1999), Jones (2013) and Sweeting (2011, for the United Kingdom), while Bartram (2018) reports a negative and significant relation.

In most of the studies, the relation is estimated after adjusting reported liabilities for the impact of the choice of the discount rate. The exceptions are Sweeting (2011), Jones (2013) and Bartram (2018). Lack of adjustment could explain the insignificant or negative relation they find since there is downward estimation bias without adjustment (shown in our simulated evidence below).

The preceding papers use regression analysis. Two recent papers deploy alternative research designs. First, Kisser et al. (2017) exploit the fact that during 1999–2005, US companies were required to report two different measures of pension liabilities. Companies had more discretion regarding reporting of "accrued liability" than "current liability." The authors find that accrued liability was lower than current liability on average, and they show that the main reason for the difference was that the assumed discount rate for accrued liability was 170 bp higher on average than the rate for current liability, which was set by regulation. The gap between the two measures of liability is positively related to pension deficit (without adjustment) as a proportion of the liability, which is consistent with the majority of results in the papers that use regression.

[3] These decisions also involve the scheme's trustees and actuary, but US and UK research implies that the company has control. Naughton (2019, p.457) states that the choice of reported pension assumptions in the US is "exclusively the domain of the firm and its auditors." Anantharaman (2017) presents evidence that US companies switch actuaries in order to facilitate upward manipulation of discount rates.

[4] There is also evidence that chief executive officers (CEOs) avoid cash contributions to schemes in deficit, if their remuneration is linked to cash flows from operations (Cheng & Swenson, 2018). Since pressure to contribute is linked to the scale of the deficit, CEO pay arrangements could result in a further incentive to reduce reported deficits by manipulating pension assumptions.

[5] Pension expense = Service cost (= present value of additional benefits accrued by members) + Interest cost (= value of pension obligations × discount rate) – Expected gain on scheme assets. If the discount rate increases, the effect on service cost is unambiguously negative. The effect on interest cost is uncertain; it depends on whether the higher rate itself outweighs the reduction in the value of the liabilities due to the higher rate. The discount rate has no effect (before 2013) or a negative effect (after 2013) on the expected gain on scheme assets. Calculations by Naughton (2019) for US schemes indicate that pension expense is sensitive to the discount rate, with a negative sign for the relation.

[6] Other possible determinants of the discretionary discount rate are rehearsed in our discussion of control variables (Section 6).

[7] The existence of a deficit implies that the sponsor has reasons for not fully funding its scheme, at least in the short term. For a review of possible reasons why a scheme might be either under- or overfunded, see Sutcliffe (2016, pp.221–227).

Second, Naughton (2019) finds that there is an upward bias in reported discount rates and that the bias diminishes following the implementation of Financial Accounting Standards Board (FASB) Statement 132R in 2003, which implied there would be greater regulatory scrutiny of companies' reported rates. Bias is measured by subtracting a scheme-specific benchmark rate from its discount rate, where the benchmark rate is the yield on AA-rated corporate bonds with a duration matching the estimated duration of the scheme.[8] Naughton does not test the relation between the discretionary discount rate and funding status but rather argues that managers trade off a wish for a higher rate against a wish to avoid regulatory action. The evidence in both these studies is that discount rates are on average strongly biased upwards, suggesting that deficit mitigation is the dominant motive regarding discretion in rates, at least during times when a majority of companies' pension schemes are in deficit.[9]

3 | ESTIMATING THE IMPACT OF UNBIASED DEFICIT RATIO ON CHOICE OF DISCOUNT RATE

3.1 | Estimation assuming benchmark discount rate (*BDR*) and duration are known

This section introduces our recommended method of estimation. In 2005, companies in several countries including the United Kingdom started reporting under IFRS. IAS 19 *Employee Benefits* specifies that "the rate used to discount post-employment benefit obligations (both funded and unfunded) shall be determined by reference to market yields at the end of the reporting period on high-quality corporate bonds," matched by duration and currency (IFRS, 2018, p.1136).[10] This requirement makes it easier to measure discretion in the choice of a discount rate. It specifies a benchmark rate that is viewed by accounting regulators as unbiased and implies that the use of discretion can be measured by the difference between the discount and benchmark rates. A non-zero value for the difference does not in itself establish there has been deliberate bias in the choice of a discount rate, for example, to reduce the reported deficit. IAS 19 permits flexibility in the choice of a discount rate, implying that a non-zero value can be appropriate and indeed informative as implied by the evidence on value relevance in Hann et al. (2007).

We model the determination of the discount rate (*DR*) as the *BDR* plus manipulative and non-manipulative discretion:

$$DR_{it} = BDR_{it} + bUDefRatio_{it} + e_{it}. \tag{1}$$

where *i* is a scheme, *t* is a year and *UDefRatio* is the unbiased deficit ratio, given by the unbiased value of the pension benefit obligations (*PBO*) divided by pension assets (*PAssets*), where unbiased *PBO* is measured using the scheme's benchmark rate.[11] *e* is "error" or noise in the setting of the discount rate around its expected value of *BDR* + *bUDefRatio*. Since *e* varies independently of *UDefRatio*, a non-zero value of *e* in a given scheme-year can be viewed as a non-manipulative discretionary adjustment to *DR*. The parameter we wish to estimate is *b*, the underlying relation between *DR* and *UDefRatio* (assumed to be linear). The deficit-reduction hypothesis predicts that underlying $b > 0$. If *b* is zero, *DR* is determined by *BDR* and non-manipulative discretion.

[8] This required benchmark first appeared in the United States in 1985; in FAS 87. N. Godwin (1999) reports that the average reported discount rate was *below* the yield on AA-rated bonds during 1987–1991, an era when less than 40% of his sample firms reported at least one pension plan in deficit.

[9] Comprix and Muller (2011) identify a different case of manipulation. They argue that companies that freeze their DB plans reduce downward both the discount rate and expected rate of return on plan assets, in order to increase reported liabilities and reduce employee resistance to the freeze.

[10] The instruction quoted was part of the standard when it was first issued in 1998 and has remained unchanged since then. An alternative view is that the discount rate should be given by the expected rate of return on pension assets, a rate that would usually be higher than that on high-quality bonds and would result in a healthier reported funding status. This view is discussed in Anantharaman and Henderson (2020).

[11] Other measures of funding status are found in the literature. The key point in all cases is that pension liabilities or the pension deficit (= *PBO* – *PAssets*) is measured as a proportion of pension assets. We choose *PBO/PAssets* because it simplifies our conceptual analysis.

Consider now an estimate of b in (1) from an ordinary least squares (OLS) regression of the discount rate on the reported deficit ratio:

$$DR_{it} = \alpha + \beta DefRatio_{it} + \varepsilon_{it}, \tag{2}$$

where $DefRatio$ = reported $PBO/PAssets$. Since reported PBO is after the impact of any discretionary adjustment to DR, the value of $DefRatio$ will deviate from that of (unobservable) $UDefRatio$. Because of this problem, most papers adjust PBO to provide an estimate of unbiased PBO. But as explained below, the existing methods of adjustment result in biased estimates of b.

This being so, we propose an alternative method of adjusting the reported ratio, which can be derived as follows. Consider a scheme with a series of promised future pension payments $Y_1, Y_2 \dots Y_T$. The present value of the payments calculated using the scheme's benchmark rate is its unbiased PBO. Let its duration be n years, and let Y_n be the single payment to be made after n years such that the present value (PV) of Y_n is equal to the scheme's PBO. In other words, Y_n is the promised payoff of a zero-coupon bond with the same value and duration as the promised payoffs of the scheme. The deficit ratio is then given by:

$$DefRatio_{it} = \frac{(Y_n)_{it}}{(1 + DR_{it})^{n_{it}}} / PAssets_{it}. \tag{3}$$

Starting from a potentially biased $DefRatio$, it follows that the unbiased $UDefRatio$ is given by:

$$UDefRatio_{it} = DefRatio_{it} \times \frac{(1 + DR_{it})^{n_{it}}}{(1 + BDR_{it})^{n_{it}}} = AdjDefRatio_{1it}, \tag{4}$$

If n and BDR are known without error, $AdjDefRatio_1 = UDefRatio$. We call (4) the *duration-based* adjustment.[12]

The above adjustment is similar to one developed by Hann et al. (2007), in which the latter involves an approximation of duration. However, this adjustment cannot be used for companies reporting under IFRS because it calls for the accumulated pension-benefit obligation (ABO), which is not reported under IAS 19.[13] Our duration-based adjustment is feasible under IAS 19 because scheme duration is reported, which enables us to obtain n and BDR specific to scheme-years.

Returning to regression (2), even with an accurate estimate of $UDefRatio$, β will be biased downwards as an estimate of underlying b. This is because of variation in scheme-specific BDR, which jointly affects both DR and $UDefRatio$. Variation arises because of differences in scheme duration, variation in BDR over time and uncertainty about the exact BDR given scheme duration—in practice some discretion can be exercised by a scheme in the measurement of its BDR (e.g., see Goldman Sachs, 2020). The solution is to control for the benchmark rate. When this is done, or when the dependent variable is $DR - BDR$, the coefficient on $UDefRatio$ or $AdjDefRatio_1$ provides an unbiased estimate of b. This is shown formally in Appendix 1. However, the only paper to estimate b that includes the benchmark rate is Fried and Davis-Friday (2013). The omission of BDR in other papers is perhaps because it is hard to estimate in the absence of reported duration under US generally accepted accounting principles (GAAP)—although existing adjustments to PBO also call for a benchmark rate. The use of year or month fixed effects (FE; Billings et al., 2017) mitigates the downward bias but is an

[12] In practice n and BDR are estimated with error, which means that $AdjDefRatio_1$ will not be exactly equal to the unobserved $UDefRatio$. We explore through simulations whether estimation error in $AdjDefRatio_1$ affects estimation of underlying b, and find that the effect is negligible.

[13] Hann et al.'s adjustment to PBO is: $AdjPBO_i = \frac{PV_N[Pen_i(1+g_i)^{N_i}]}{(1+BDR_i)^{N_i}}$ where Pen is the aggregate pension of active members based on their current salaries, . N is average number of years of members to retirement (ABO is needed to infer N), g is growth in salary over the next N years and PV_N is the PV after N years of an annuity of $Pen(1+g)^N$ paid for a retirement period of 15 years, using the benchmark rate as the discount rate. A problem with their adjustment, in addition to needing ABO, is that scheme duration is implicitly given by inferred N plus assumed 15-year life expectancy on retirement. This is an approximation. Hann et al. (p.116) note that "while our estimate of N is appropriate for measuring the effects of compensation growth, it has measurement error when estimating the effects of discount rates on the PBO."

incomplete solution. In summary, our recommended regression specification is to estimate unbiased funding status by $AdjDefRatio_1$ and control for the benchmark rate.[14]

3.2 | Other methods of estimation

Two different adjustments to pension liabilities have been used in the literature. One assumes that pension benefits are perpetuity and replace the reported discount rate by the scheme's benchmark rate (Bodie et al., 1984; Feldstein & Mørck, 1983; Fried & Davis-Friday, 2013; N. Godwin, 1999; Gopalakrishnan & Sugrue, 1995; Thies & Sturrock, 1988).[15] The *perpetual-pension* adjustment, $AdjDefRatio_2$, is therefore:

$$AdjDefRatio_{2it} = DefRatio_{it} \times \frac{DR_{it}}{BDR_{it}}. \tag{5}$$

The other existing method assumes a fixed duration or fixed sensitivity of *PBO* to *DR*. Francis and Reiter (1987) and Carroll and Niehaus (1998) use a "4% rule of thumb" whereby *PBO* falls by 4 percentage points for each 0.25-point rise in *DR*, that is, a sensitivity of *PBO* to *DR* of 16 times. Billings et al. (2017) follow a suggestion by the UK Accounting Standards Board (ASB, 2007, p.17) that the sensitivity is 19 times.[16] Under the ASB's assumption, the *fixed-duration* adjustment is:

$$AdjDefRatio_{3it} = DefRatio_{it} \times [1 + 19\,(DR_{it} - BDR_{it})]\,. \tag{6}$$

The existing adjustment methods are not ideal because the coefficients β_2 on $AdjDefRatio_2$ and β_3 on $AdjDefRatio_3$ are biased as estimates underlying *b*, in both univariate regressions and controlling for the benchmark rate. The Appendix shows formally how the biases arise, for the case of $b = 0$.

4 | EVIDENCE FROM SIMULATIONS

4.1 | Conceptual framework

Our next step is to assess various approaches to estimating underlying *b* using evidence from simulations. We estimate regressions with simulated data generated using a range of assumptions about the value of underlying *b* and average scheme duration. This exercise enables us to confirm that our recommended method is accurate and to estimate the scale of the biases that can arise from existing methods. Such an assessment is not possible using a single set of real data because underlying *b* is not known, the estimation biases vary with both *b* and average duration and might not generalize to a different dataset.

To explain the simulations, we begin with the relation between pension assets and the expected value of future obligations, E(*Y*). This is given by:

$$\mathrm{E}\ (Y_i|n_i) = PAssets_i\ \times (1 + R)^{n_i}, \tag{7}$$

[14] We note that, if higher *UDefRatio* results in discretionary *DR* being used in conjunction with other actuarial discretion that reduces *PBO* (e.g., understated salary growth), use of $AdjDefRatio_1$ to estimate underlying *b* will still result in downward bias.

[15] Thies and Sturrock (1988) assume that the benefits are an annuity and calculate a ratio of two annuity factors using the reported discount rate in the numerator and the benchmark rate in the denominator.

[16] In Asthana (1999), the assumed fixed relation between *DR* and *PBO* is estimated by the coefficient on *DR* from a regression of the reported deficit ratio on *DR* and several other explanatory variables. His adjusted deficit ratio is also adjusted for choice of actuarial-cost method and of salary growth rate.

where R is the expected rate of return on scheme assets. R is assumed to be the same for all schemes; its role is merely to set the size of deficit ratios in general.[17] Scheme duration n varies across schemes and is drawn from a normal distribution around a fixed mean $\bar{n}$. The value of future obligations also varies across schemes, with Yn drawn from a normal distribution around $\mathrm{E}(Y|n)$:

$$Y_i\ n_i = \mathrm{E}\left(Y_i|n_i\right) + x_i\sigma = \mathrm{E}\left(Y_i|n_i\right)\ \left(1 + x_i k\right) = PAssets_i (1 + R)^{n_i}\ \left(1 + x_i k\right), \tag{8}$$

where x follows a standard normal distribution across simulated schemes, and k is a fixed multiple of $\mathrm{E}(Y|n)$ that determines the size of its standard deviation σ.[18] This random element affecting *PBO* is entirely independent of the simulated choice of *DR* and *BDR*. The simulated reported deficit ratio is then:

$$DefRatio_{it} = \frac{PBO_{it}}{PAssets_i} = \frac{PAssets_i (1\ +\ R)^{n_i}\ (1\ +\ x_i k)\ / (1\ +\ DR_{it})^{n_i}}{PAssets_i} = \frac{(1\ +\ R)^{n_i}\ (1\ +\ x_i k)}{(1\ +\ DR_{it})^{n_i}}. \tag{9}$$

We see that *PAssets* is not needed for the simulations. This is because we are interested in the relation between *DR* and *PBO* given *PAssets*. The unbiased *UDefRatio* is generated using (9) and the scheme's *BDR*, which varies by year:

$$U\ DefRatio_{it} = \frac{(1\ +\ R)^{n_i}\ (1\ +\ x_i k)}{(1\ +\ BDR_{it})^{n_i}}. \tag{10}$$

Following equation (1), which gives the assumed process by which discount rates are determined, simulated *DR* is obtained by drawing from:

$$D\ R_{it} = BDR_{it}\ + bUDefRatio_{it} + e_i, \tag{11}$$

where e has a mean of zero, is normally distributed and is fixed over time for a given scheme. *BDR* is a draw from one of 10 possible values for the benchmark rate, given the scheme's simulated duration n. The values are from 10 actual yield curves for *BDR* for UK companies, one for each of the 10 years in our sample period. To simulate manipulation in the choice of a discount rate, a value of b other than 0 is inserted in equation (11), and the simulated biased *DR* is then used in (9) to obtain values of *DefRatio*.

To obtain adjusted deficit ratios, simulated deficit ratios are adjusted using equation (4), (5) or (6) for the duration-based, perpetual-pension and fixed-duration adjustment methods, respectively. The adjustments call for the benchmark rate, which in practice is estimated with error (companies do not report their benchmark rates). The error arises mainly because we estimate scheme duration with error, by means of a predictive model based on scheme and sponsor characteristics (explained in Section 5). But we know the actual reported duration for 500 scheme-years, through hand collection. In a sub-sample of 100 observations used to test our estimates of duration, we find that the relation between reported duration n and our estimate $\hat{n}$ is:

$$\hat{n}_i = 13.0 + 0.25 n_i + \epsilon_i, \tag{12}$$

with $\varepsilon \sim 0, 1.6$. To simulate error in the estimation of duration and hence *BDR*, we draw simulated $\hat{n}$ from the distribution given by (12). The benchmark rate used to calculate the adjusted deficit ratios, $\widehat{BDR}$, is the rate for the simulated

[17] Higher R implies higher $\mathrm{E}(Y|n)$n rela. tion to *PAssets*, and therefore a higher deficit ratio, for any given simulated *DR*.

[18] Both Y_i and n_i are assumed to be fixed over time in the simulations, which is why there are no year subscripts in equations (7) and (8).

year assuming duration is its estimate $\hat{n}$ rather than its true value n. The simulated adjusted deficit ratios are given by:

$$AdjDefRatio_{1it} = DefRatio_{it} \times \frac{(1+DR_{it})^{\hat{n}_i}}{\left(1+\widehat{BDR}_{it}\right)^{\hat{n}_i}}, \tag{13}$$

$$AdjDefRatio_{2it} = DefRatio_{it} \times \frac{DR_{it}}{\widehat{BDR}_{it}}, \tag{14}$$

$$AdjDefRatio_{3it} = DefRatio_{it} \times \left[1 + 19(DR_{ti} - \widehat{BDR}_{it}\right], \tag{15}$$

$AdjDefRatio_1$ is the same as *UDefRatio*, except for the simulated uncertainty about n and *BDR*. We report results for the following regressions using simulated data, for different values of underlying b in equation (11), and different measures of deficit ratio:

$$DR_{it} = \alpha + \beta(a\ measure\ of\ deficit\ ratio)_{it} + \varepsilon_{it}, \tag{16}$$

$$DR_{it} = \alpha + \beta(a\ measure\ of\ deficit\ ratio)_{i} + \gamma BDR_{it} + \varepsilon_{it}. \tag{17}$$

4.2 | Calibration of parameters

Table 1 summarizes the assumptions made about the distributions of parameter values used in the simulations. The distributions are calibrated to achieve a mean and standard deviation for each variable that is close to the values of those statistics for the variables in our sample data (after winsorization). It is important that the simulated data are similar to the sample data, in order for the β estimates using simulated data to be informative about the estimation biases likely to arise in real data.

Table 2 shows summary statistics for the sample data and simulated data for values of underlying b set between 0 and 1.0. The mean simulated duration is 17.8 years as in the sample; later, we allow the mean duration to vary. Mean-simulated estimated *BDR* ($\widehat{BDR}$) is 4.25% as in the sample, and mean-simulated actual *BDR* is 4.19%.[19] When assumed $b = 0$ (i.e., discretionary *DR* is zero), simulated $DR = BDR$. Simulated *DR* increases with assumed b, reflecting increasing upward use of discretion, whereas *BDR* and its estimate $\widehat{BDR}$ are invariant with respect to b.

The mean reported deficit ratio in the sample data is 1.22. To simulate similar levels of reported deficit ratio (which depend partly on b), we set the expected return on pension assets at 6.0% per year, well above the mean benchmark rate. This has the effect of making simulated pension obligations larger than assets. As values of underlying b increase, the mean *DefRatio* (the simulated reported deficit ratio) decreases due to increasing discretionary *DR*. Both *UDefRatio* and $AdjDefRatio_1$ are unaffected by b, by definition, but the other two adjusted ratios are affected. The mean value of $AdjDefRatio_2$ increases materially with b, while $AdjDefRatio_3$ varies much less and is close to the mean of $AdjDefRatio_1$.

[19] The small difference arises because simulated error in estimating *BDR* arises from simulated error in estimating duration (equation 12), and given a non-linear yield curve for *BDR*, mean $BDR \neq$ mean $\widehat{BDR}$

TABLE 1 Parameter values used in simulations

Parameter	Description	Calibration
b	Underlying relation between *DR* and unbiased deficit ratio, equation (10)	Not observable in sample data. A different level of underlying b is set for each set of simulations
e_i	Noise in *DR* in equation (10), in relation to its value given no bias ($b = 0$) or upward manipulation ($b > 0$)	Mean $e_i = 0$; std. dev. = 0.7%. Chosen to generate a std. dev. of simulated *DR* similar to that of sample *DR*. e_i is fixed over time for a given scheme
n_i	Scheme duration in years	n_i is assumed to be normally distributed. Its mean and std. dev. are similar to the mean and std. dev. of reported duration in sample data, based on 500 hand-collected observations
$\hat{n}_i$	Estimated scheme duration in years	Generated from $\hat{n}_i = 13.3 + 0.25n_i + \varepsilon_i, \varepsilon_i \sim 0, 1.6$. This relation and its variation fit the relation between $\widehat{n}_i$ and n_i in our hand-collected data. The std. dev. of estimated $\widehat{n}_i$ is less than that of n_i in both real and simulated data
BDR_i	For a given n_i, BDR_i is a draw from one of 10 possible values of *BDR*	Designed to replicate how sample *BDR* varies. The draw is from the actual pension yield curves for the UK for the 10 sample years (see Section 5).
$\widehat{BDR}_i$	As for BDR_i, but the draw for $\widehat{BDR}_i$ is based on $\hat{n}_i$, a duration estimated with error	Designed to simulate estimation of *BDR* with error, as in sample data. The mean and std. dev. of $\widehat{BDR}_i$ are almost the same as the mean and std. dev. of estimated *BDR* in sample data.
R	Expected rate of return on scheme assets; used to set sample deficit ratio	$R = 6.0\%$. Chosen to be sufficiently above mean *BDR* in simulated data that mean simulated and sample reported deficit ratios are similar.
k	Multiple of $\mathrm{E}(Y_i\vert n_i) \times x_i;\ x_i \sim 0, 1.0$ Sets the std. dev. of future pension obligations. $Y_i n_i$. x_i is fixed over time for a given scheme	$k = 0.45$. Chosen to generate a std. dev. of simulated deficit ratio ($PBO_i/PAssets_i$) similar to that of sample deficit ratio.

4.3 | Results of simulations

Table 3 shows results for β in regressions (16) and (17) using simulated data, for values of underlying b between 0 and 1.0 and a fixed mean duration of 17.8 years. In the univariate specification (Panel A), omitting *BDR*, the betas on *UDefRatio* and $AdjDefRatio_1$ are biased downwards—they are about 50 bp below the values of underlying b. The reason is a variation in the benchmark rate as discussed. The betas on all the other measures of deficit ratio are also substantially biased downwards. These results illustrate the material bias that can arise in estimating underlying b when omitting the benchmark rate. They imply that most previous estimates, which are from regressions that omit the benchmark rate (Section 2), are likely to be understated.[20]

Panel B shows results with *BDR* included as a control variable (the results are virtually the same in unreported univariate regressions with *DR* – *BDR* as dependent variable). Crucially, the beta coefficients on *UDefRatio* and $AdjDefRatio_1$ are very close to underlying b and to each other, as expected.[21] The downward bias in estimated betas caused by omitting *BDR* is eliminated. The simulations show that our recommended estimation method—regressing *DR* on $AdjDefRatio_1$ while controlling for *BDR*—provides a direct and accurate estimate of underlying b. The results

[20] The. extent to which previous estimates of b might be inaccurate cannot be determined precisely from our simulations, which are calibrated on our UK sample data. Estimation errors in a given study will depend both on features of the data and on the regression specification used in that study.

[21] The slight variation of estimated β on *UDefRatio* in relation to b arises because we winsorize simulated *DR* as in the sample data. If *DR* is not winsorized, β and b are exactly the same.

TABLE 2 Summary statistics for sample and simulated data

		Simulated data for values of *b*				
Variable	**Sample data**	**0**	**0.25**	**0.50**	**0.75**	**1.00**
DR	4.24	4.19	4.53	4.87	5.21	5.55
	(1.26)	(1.25)	(1.22)	(1.21)	(1.22)	(1.25)
BDR	Not observable	4.19	4.19	4.19	4.19	4.19
		(1.07)	(1.07)	(1.07)	(1.07)	(1.07)
$\widehat{BDR}$	4.25	4.25	4.25	4.25	4.25	4.25
	(1.10)	(1.03)	(1.03)	(1.03)	(1.03)	(1.03)
DefRatio	1.22	1.37	1.28	1.19	1.11	1.04
	(0.45)	(0.69)	(0.61)	(0.54)	(0.48)	(0.43)
UDefRatio	Not observable	1.36	1.36	1.36	1.36	1.36
		(0.66)	(0.66)	(0.66)	(0.66)	(0.66)
$AdjDefRatio_1$	1.23	1.35	1.35	1.35	1.35	1.35
	(0.56)	(0.66)	(0.66)	(0.66)	(0.66)	(0.66)
$AdjDefRatio_2$	1.24	1.32	1.37	1.41	1.43	1.45
	(0.70)	(0.67)	(0.72)	(0.75)	(0.77)	(0.78)
$AdjDefRatio_3$	1.24	1.34	1.34	1.34	1.34	1.33
	(0.67)	(0.66)	(0.66)	(0.66)	(0.64)	(0.63)
Actual duration (n)	17.75	17.75	17.75	17.75	17.75	17.75
	(3.39)	(3.99)	(3.99)	(3.99)	(3.99)	(3.99)
Estimated duration ($\hat{n}$)	17.75	17.75	17.75	17.75	17.75	17.75
	(1.94)	(1.90)	(1.90)	(1.90)	(1.90)	(1.90)

Note: The table shows means of variables in sample data and in simulated data for five values of underlying *b* used in equation (11) to generate *DR*. Standard deviations are in brackets. The variables and generation of simulated data are described in Section 4 and Table 1. Estimation of *BDR* and duration in the sample is explained in Section 5. Sample data are from 2538 scheme-years, except that data for actual duration *n* are from 500 hand-collected observations. Simulated data for each *b* are means from 5000 regressions, with 2538 scheme-years per regression.

also show that the simulated estimation errors for *n* and *BDR*, incorporated in the calculation of $AdjDefRatio_1$, do not compromise the use of $AdjDefRatio_1$ for the purpose of estimating *b*. This is reassuring since we estimate *n* and *BDR* with error in practice, and we have set the simulated estimation errors to be similar to those in the sample data.

Regarding the perpetual-pension adjustment, β_2 on $AdjDefRatio_2$ is not consistently aligned with *b*. The bias of β_2 is positive for $b = 0$, diminishes as *b* increases and turns negative for $b = 1$. But the fixed-duration adjustment performs well: β_3 on $AdjDefRatio_3$ is close to *b*, with a slight upward bias. A further point from Panel B is that the beta estimates on the deficit ratios are more precise when the benchmark rate is included.[22]

In Table 3 duration varies across schemes around a fixed average duration. However, the biases of estimated β vary with average duration, and this variation is of interest since average duration will fall in future due to the closure of many schemes to new entrants (e.g., Sutcliffe, 2016, chap.2). Figure 1 shows the biases in estimated betas allowing

[22] The *t*-statistics on the coefficients are higher than in the empirical regressions, especially for larger values of *b*. This is despite the fact that the variances of the variables (e.g., *DR*) are similar in simulated and sample data (Table 2). The reason for the lower *t*-statistics in the empirical regressions is heteroskadisticity in the error term, which is controlled for by clustered standard errors that reduce the reported *t*-statistics. The simulated data do not incorporate heteroscedasticity.

TABLE 3 Simulated relation between discount rate and deficit ratio

Panel A. Specification: $DR_{it} = \alpha + \beta(a\ measure\ of\ deficit\ ratio)_{it} + \varepsilon_{it}$					
Measure of deficit ratio	**Value of b**				
	0	**0.25**	**0.50**	**0.75**	**1.00**
UDefRatio	−0.500	−0.252	−0.006	0.239	0.485
	(−6.98)	(−3.52)	(−0.07)	(3.40)	(6.95)
DefRatio	−0.676	−0.523	−0.321	−0.069	0.239
	(−10.09)	(−6.52)	(−3.34)	(−0.60)	(1.82)
$AdjDefRatio_1$	−0.459	−0.225	0.012	0.250	0.488
	(−6.58)	(−3.19)	(0.19)	(3.54)	(6.91)
$AdjDefRatio_2$	−0.279	−0.153	−0.012	0.135	0.287
	(−4.18)	(−2.41)	(−0.19)	(2.18)	(4.52)
$AdjDefRatio_3$	−0.432	−0.215	0.007	0.236	0.478
	(−6.13)	(−3.03)	(0.10)	(3.16)	(6.11)
Panel B. Specification: $DR_{it} = \alpha + \beta(a\ measure\ of\ deficit\ ratio)_{it} + \gamma BDR_{it} + \varepsilon_{it}$					
Measure of deficit ratio	**Value of b**				
	0	**0.25**	**0.50**	**0.75**	**1.00**
UDefRatio	−0.002	0.244	0.489	0.735	0.980
	(−0.03)	(4.24)	(8.55)	(12.93)	(17.40)
DefRatio	−0.223	−0.024	0.224	0.516	0.855
	(−3.88)	(−0.35)	(2.70)	(5.27)	(7.50)
$AdjDefRatio_1$	0.060	0.296	0.536	0.773	1.010
	(0.98)	(4.83)	(8.76)	(12.66)	(16.54)
$AdjDefRatio_2$	0.228	0.395	0.570	0.750	0.937
	(3.79)	(7.07)	(10.60)	(14.04)	(17.27)
$AdjDefRatio_3$	0.107	0.323	0.549	0.783	1.030
	(1.71)	(5.18)	(8.69)	(12.01)	(15.13)

Note: The table shows the mean coefficient β (mean *t*-statistic in brackets) on five measures of deficit ratio from simulated regressions. The dependent variable is *DR*, generated using equation (11). Results are shown for five values of underlying *b* in (11) as in Table 2. In Panel A, the explanatory variable is a measure of deficit ratio only. In Panel B, *BDR* is added as a control variable. The five ratios are defined in Section 4. There are 5000 simulated regressions for each combination of *b* and a measure of deficit ratio, and each regression has 2538 simulated scheme-years. *t*-statistics are calculated using standard errors clustered at scheme level.

both *b* to vary between 0 and 1.0 as in Table 3, and average duration to vary between 10 and 20 years, controlling for *BDR*. The biases of β_1 and β_2 are superimposed in Figure 1a and of β_1 and β_3 in Figure 1b.

We see that the bias of β_1 is very small for all combinations of *b* and average duration, confirming accurate estimation of underlying *b* from the use of $AdjDefRatio_1$ (duration-based adjustment). The biases of β_3 (fixed-duration adjustment) and especially β_2 (perpetual-pension adjustment) vary and neither is the direction of bias consistently positive or negative. Although β_3 is close to β_1, for an average duration of 17.8 years as in Table 3, β_3 becomes increasingly biased upward as the duration reduces toward 10 years for low values of *b*. The graphs illustrate that in general, neither of the existing adjustment methods can be relied upon to produce accurate estimates of underlying *b*.

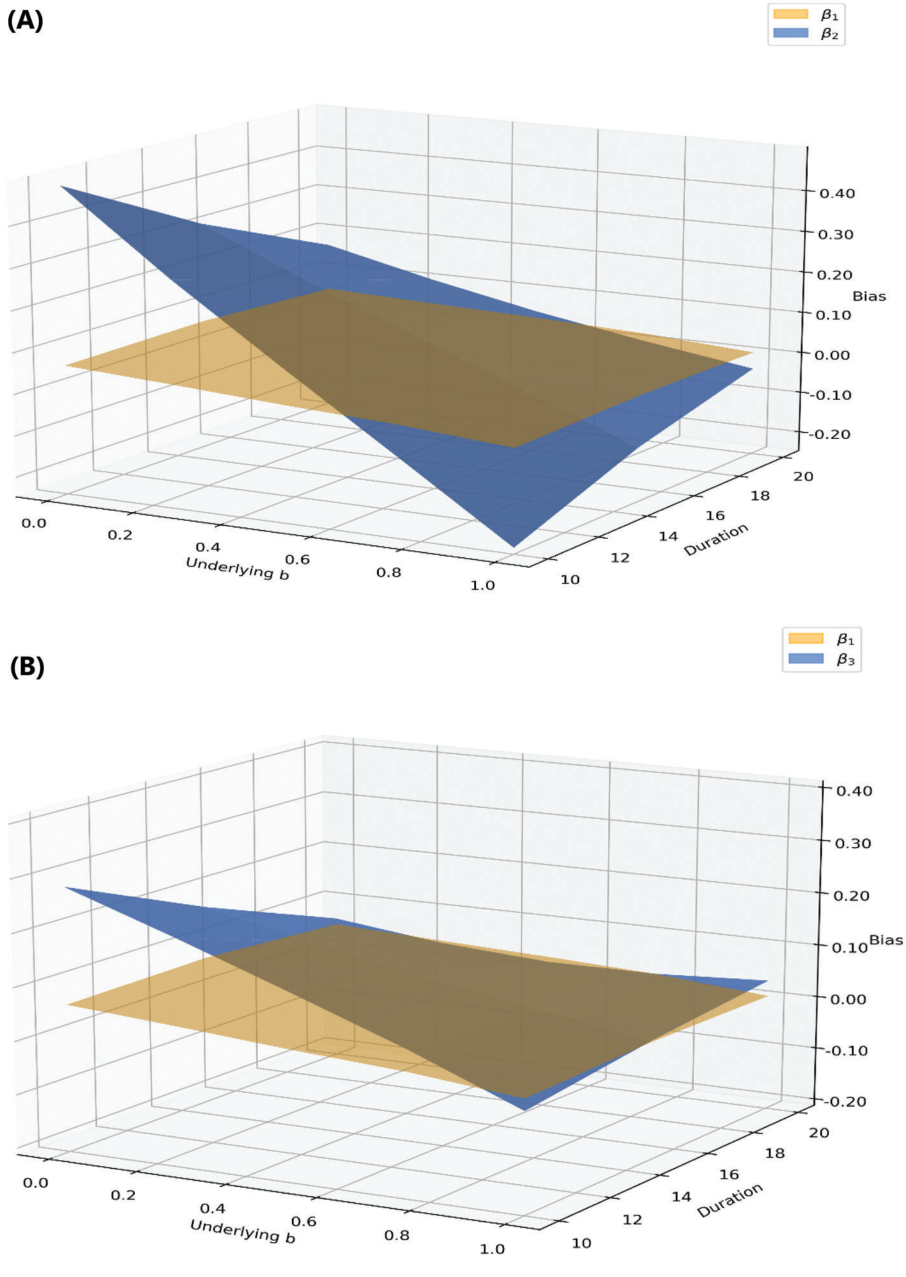

FIGURE 1 The graphs show estimated β from simulated regressions for different values of underlying b in equation (11) and average scheme duration in years. The regression specification is $DR_{it} = \alpha + \beta(\textit{a measure of deficit ratio})_{it} + \gamma BDR_{it} + \varepsilon_{it}$. Bias = $\beta - b$. (a) Biases using $AdjDefRatio_1$ and $AdjDefRatio_2$ and (b) biases using $AdjDefRatio_1$ and $AdjDefRatio_3$.

The simulation results point to two conclusions. First, they show that, in empirical estimation, the benchmark discount rate should be included as a control variable, or the dependent variable should be the discount rate minus the benchmark. Second, β_1 on $AdjDefRatio_1$ is always extremely close to underlying b, whatever the assumed values of b and average duration if the benchmark rate is controlled for. The values of β_3 are only slightly less accurate, given a fixed average duration of 17.8 years as found in our sample (and similar to the assumed duration of 19 years in the fixed-duration adjustment). More generally, both β_2 and β_3 show variable bias that depends on underlying *b* and duration. The simulations confirm that the best practice for estimation is to estimate an unbiased deficit ratio using the duration-based adjustment and to control for the benchmark rate.

5 | REGRESSION SPECIFICATION AND DATA

We now estimate underlying *b* using real data from a sample of UK companies reporting under IFRS. The sample period is 2009–2018. The first year is 2009 because the FTSE pension yield curve to estimate benchmark rates is not available before then. We estimate several regression models, the fullest being:

$$DR_{it} = \alpha + \beta(a\ measure\ of\ deficit\ ratio)_{it} + \gamma BDR_{it} + \delta X_{it} + YearFE + \varepsilon_{it}, \tag{18}$$

where X is a vector of control variables (in addition to BDR), related to either the scheme or the sponsoring company, and *YearFE* is year FE.

A scheme's duration is needed to estimate its benchmark rate. Following amendments to IAS 19 in 2011, from 2013, nearly all schemes in our sample either report duration explicitly, or it can be inferred from information on the sensitivity of pension liabilities to a small change in the reported discount rate (e.g., $\pm$ 0.25 percentage points). To estimate duration, we first hand-collect duration data for 500 scheme-years. We use 400 of these observations to train a variety of machine learning models. The remaining 100 are used to test the resulting out-of-sample duration forecasts. The explanatory variables (features in machine learning parlance) used to fit the models are a combination of scheme and sponsor characteristics.[23]

Machine learning tools are superior in this context to more traditional econometric methods such as OLS regression as they impose no functional form on the relation between duration and the explanatory variables. As a consequence, they allow for higher order interactions between variables, without a requirement to specify these a priori.[24] We fit the training data using OLS as well as three machine learning variants—a random forest, a support vector machine and a gradient boosted regression. The random forest model produces the lowest root-mean-squared error (RMSE) consistently across random shuffles of the training and test data. The random forest model is an ensemble learning technique, which aggregates the forecasts of individual decision trees (Breiman, 2001). It exploits a principle known as "bagging," under which each tree is a fit on a random subset of the data. In addition, each tree uses a random subset of the features to partition the data. Feature and subset randomization helps minimize the correlation in errors across trees and has been shown to increase forecast accuracy in ensemble methods.[25] The RMSE of the forecasts is 2.9 years. Our evidence using simulated duration data with the same degree of estimation error indicates that it should make little difference to the coefficients on deficit ratios.

The next step is to estimate the benchmark rate for each scheme-year, using hand-collected or estimated duration together with the yield curve for the relevant year. There is a published pension yield curve, the FTSE (previously Citi) pension curve, composed of individual zero-coupon interest rates for durations of up to 30 years at half-yearly intervals, derived from market yields on AA-rated US bonds. We convert from a dollar to a sterling yield by inferring the difference between spot and forward foreign exchange rates from the relevant zero-coupon yield curves derived from

[23] The scheme-related characteristics are pension liabilities/firm assets, net pension expense/firm sales, net pension interest expense/firm sales, discount rate, deficit ratio, deficit/firm market value and pension service cost/pension interest expense. The company characteristics are age, size, cash holding, earnings before interest and tax (EBIT), return on assets, leverage, interest coverage ratio, *Z*-score and industry.

[24] Machine learning has been increasingly used in accounting and corporate finance studies to uncover complex patterns in financial data and make accurate out-of-sample predictions (e.g., Bao et al., 2020; Bertomeu et al., 2020; Li et al., 2020).

[25] The only other paper we know of that estimates duration-matched benchmark rates is Naughton (2019). His predictive model for duration differs somewhat from ours; it includes certain scheme features not available for our sample, and it is not estimated using machaine learning.

long-term government bonds, under the same no-arbitrage argument that results in the interest rate parity relation:

$$(1 + BDR_n)^n = \left(\frac{1 + Yield_n}{1 + Yield_{n,\$}}\right)^n \times (1 + BDR_{n,\$})^n,$$

$$\text{or } BD\,R_n = \frac{(1 + Yield_n)(1 + BDR_{n,\$})}{1 + Yield_{n,\$}} - 1, \tag{19}$$

where $BDR_{n,\$}$ = pension discount rate for duration n in \$, as at the relevant financial year-end for the scheme-year; BDR_n = inferred pension discount rate in £; $Yield_{n,\$}$ ($Yield_n$) = zero-coupon yield of US government bonds (UK government bonds) with duration n as at the relevant financial year-end.[26]

A method of estimating scheme-specific *BDR*, given scheme duration, is an important step forward, with benefits beyond the estimation of underlying *b*. It allows us to disentangle permitted discretion, based on scheme duration and currency, from that which appears more opportunistic in nature. It, therefore, enables the scheme-specific discretionary element of the discount rate to be estimated, and thus the funding status after removing the impact of discretion. This is useful for several stakeholders. It allows the analyst community (on the buy side, sell side, in credit rating agencies and the banking sector) to assess the relative discount rate aggression being applied by pension sponsors in the valuation of their schemes. It, therefore, enables analysts to estimate funding status after removing the impact of the opportunistic discretion. The approach also allows regulatory agencies (e.g., the Pensions Regulator and Pension Protection Fund) to appraise the fidelity of pension disclosures in their assessment of both pension risk and pension governance.

We include several other control variables that could potentially have explanatory power, in view of previous evidence on the determinants of pension discount rates. Our scheme-related controls are duration and measures of deficit in relation to the sponsor's market value. Fried and Davis-Friday (2013) find a negative relation between discretionary *DR* and a rough proxy for the duration (service cost/(service cost + interest cost)). They view this as consistent with the use of discretion to reduce reported pension liabilities because a shorter duration implies that a larger increase in discretionary *DR* is needed to achieve a given reduction in PBO. On the other hand, a higher discount rate increases rather than reduces interest costs in the income statement when durations are short (e.g., 10 years; Fried et al., 2014), reducing the benefit. The expected sign on the relation between discretionary *DR* and duration is therefore uncertain.

The incentive to reduce the reported deficit should increase with the size of the unbiased deficit in relation to company value or assets (Anantharaman, 2017; Billings et al., 2017. In fact, Anantharaman's measure of funding status is *AdjDeficit/CompanyAssets* rather than *AdjDeficit/PAssets*). We measure deficit in relation to company value by *AdjDeficit/MV* = (*AdjPBO* – *PAssets*)/*Market value*. The expected sign on *AdjDeficit/MV* is positive.

Our company-related controls with their expected signs are company size (*RelSize*, +), profitability (*ROA*, –), default risk (*Z-score*, –), *Leverage* (+), and *Current ratio* (–). Large firms might have more incentive to manipulate the discount rate because they are more visible and because they might be able to exert more pressure on actuaries or auditors to sanction a high rate (Anantharaman, 2017). So we expect a positive sign on *RelSize*. Firms with low profitability, low *Z*-score (high default risk), high leverage or low current ratio, have more incentive to increase their discount rate and reduce their reported deficit because they are more likely to be in or close to financial distress (e.g., Asthana, 1999) or to be seeking to reduce debt or increase liquidity. A lower deficit reduces pressure on the firm from the scheme's

[26] We use a dollar-denominated pension yield curve because of a relative paucity of sterling AA-rated bonds. For a method of estimating a pension yield curve directly from sterling-denominated bonds, see Skinner and Ioannides (2005).

TABLE 4 Summary statistics for scheme and company variables used in empirical regressions

	Mean	Std dev	Min	P25	P50	P75	Max	N
DR (%)	4.24	1.26	0.70	3.25	4.32	5.30	8.25	2538
BDR (%)	4.25	1.10	1.78	3.27	4.27	5.03	6.19	2538
DefRatio	1.22	0.45	0.82	1.03	1.14	1.27	4.94	2538
$AdjDefRatio_1$	1.23	0.56	0.77	1.00	1.12	1.26	5.25	2538
$AdjDefRatio_2$	1.24	0.70	0.64	0.98	1.11	1.26	6.31	2538
$AdjDefRatio_3$	1.24	0.67	0.74	1.00	1.11	1.25	6.19	2538
Duration (years)	17.75	1.94	8.00	16.80	17.72	18.76	30.00	2538
Deficit/MV	0.12	0.26	−0.21	0.00	0.03	0.12	1.50	2538
$AdjDeficit_1/MV$	0.13	0.65	−2.46	−0.00	0.05	0.18	3.78	2538
$AdjDeficit_2/MV$	0.10	0.35	−1.10	−0.00	0.02	0.11	2.22	2538
$AdjDeficit_3/MV$	0.10	0.30	−1.07	−0.00	0.02	0.11	1.75	2538
RelSize	2.65	1.10	1.00	2.00	3.00	4.00	4.00	2538
Leverage	0.22	0.16	0.00	0.09	0.21	0.31	0.77	2538
ROA	0.04	0.08	−0.33	0.02	0.05	0.08	0.24	2538
Z-score	1.18	0.76	0.16	0.65	1.00	1.52	4.16	2536
Current ratio	1.06	0.68	0.19	0.68	0.94	1.24	4.86	2522

Note: The sample period is 2009–2018. Scheme variables: Estimation of *BDR* is explained in Section 5; *DefRatio* = *PBO*/pension assets; in $AdjDefRatio_{1(2)(3)}$, *PBO* is adjusted as in equations (4) to (6); *Duration* is hand-collected for 500 scheme-years and estimated by means of a predictive model for the remainder (explained in Section 5); *Deficit/MV* = *PBO* minus pension assets, divided by market value of company's equity as at its financial year-end for the relevant scheme-year; in $AdjDeficit_{1(2)(3)}/MV$, *PBO* is adjusted as in equations (4) to (6). Company variables: *RelSize* = variable indicating the quartile a company is in by a ranking of the assets of sample companies in the same country and calendar year (Q4 = largest); *ROA* = return on assets = profit after tax divided by total assets; *Z-score* = *Z*-score calculated following Altman (2000); *Leverage* = total debt divided by total assets; *Current ratio* = current assets divided by current liabilities. Values are shown after winsorization at the 1st and 99th percentile. *N* = scheme-years.

trustees or regulator to increase cash contributions to the scheme. A lower deficit also reduces leverage inclusive of pension obligations, which implies greater creditworthiness and could avoid breach of loan covenants.[27]

The data are from Worldscope.[28] We include all current and former listed companies that are DB sponsors in the sample period for at least 3 consecutive years. Financial-sector companies are excluded because their financial characteristics are not comparable with those of non-financial companies. We also exclude scheme-years with missing values for the discount rate. The final sample consists of 2538 scheme-years from 389 listed companies of all sizes. Continuous variables including the discount rate are winsorized at the 1st and 99th percentiles.

Table 4 provides summary statistics for *DR* and the explanatory variables. There is much heterogeneity in *DR*, *BDR* and each of the measures of deficit ratio, even after winsorizing. We note that the mean *DR* and *BDR* are almost identical; in fact, *DR* is below *BDR* in 65% of scheme-years (not shown). Hence, *DR* is biased *downward* more often than upward. This is despite the fact that the mean reported deficit ratio is 1.22 (median = 1.14) and that there is a reported deficit in 82% of scheme-years (not shown). It is only for scheme-years with values of $AdjDefRatio_1$ above 1.14 that a

[27] Bartram (2018) finds a highly significant positive relation between the discount rate and expected return on pension assets. But this relation is likely to reflect simultaneous manipulation of pension assumptions to make schemes "look better," as Bartram notes (p.342), and is therefore unlikely to be causal. Hence, we do not include expected return on pension assets as an explanatory variable.

[28] Pension data are also available from Thomson Reuters Eikon and Capital IQ. Worldscope's data are more comprehensive. If a company has more than one pension scheme, the assets and liabilities are consolidated. If the schemes use different discount rates, Worldscope reports the mid-point of the rates.

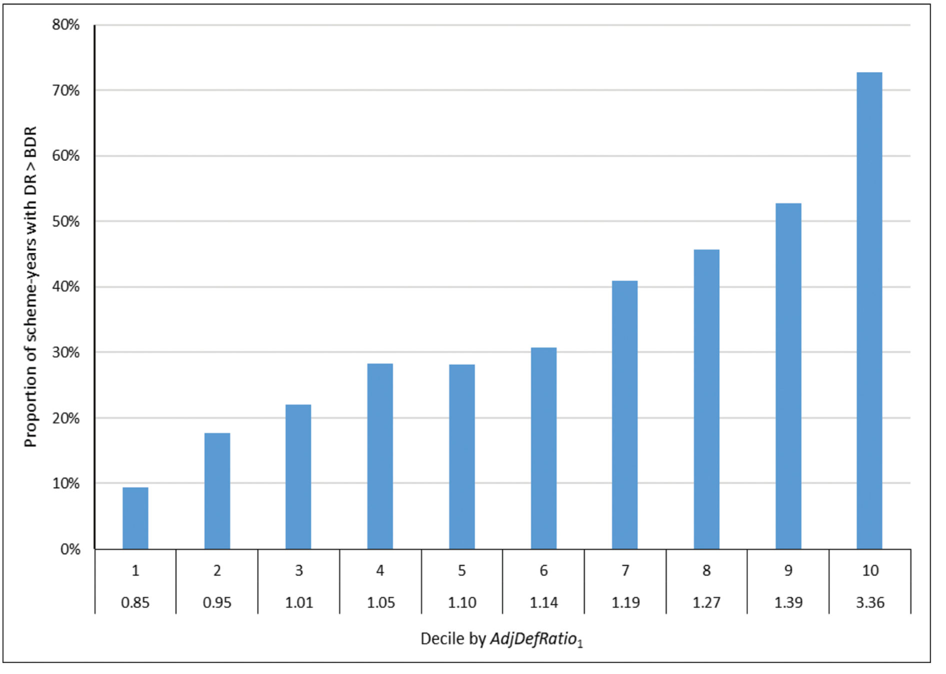

FIGURE 2 Discretionary discount rate by funding status. The graph shows the proportion of scheme-years with *DR* > *BDR*, by decile of $AdjDefRatio_1$, ranked from scheme-years with the largest surplus (Decile 1) to largest deficit (Decile 10). The median value of $AdjDefRatio_1$ for each decile is shown beneath its rank.

majority of sponsors choose *DR* above *BDR*. Hence, for smaller deficits, the reported deficit ratio tends to be larger than the estimated unbiased ratio.

The absence of upward bias in *DR* on average contrasts with sizable upward bias found for US companies (see Section 2). Naughton (2019, fig. 2) estimates that mean *DR* – *BDR* is 30 bp in 2004–2005, with *BDR* estimated using scheme-specific duration. Kisser et al. (2017) using a different approach estimate that mean *DR* – *BDR* is as high as 170 bp during 1999–2007.

6 | RESULTS

6.1 | Estimates of underlying *b*

Table 5 shows the results of regressions that have the same specifications as the simulated regressions in Table 3. Panel A shows univariate regressions. The results suggest that the discount rate is positively related to the unbiased deficit ratio, even without controlling for benchmark rates. The coefficient on $AdjDefRatio_1$ is 0.807 ($t = 4.52$). Panel B shows results with *BDR* included.[29] These regressions are better specified, with a much higher R^2. The coefficients on *BDR* are between 0.84 and 0.90 and are highly significant, indicating that discount rates track their estimated benchmark rates closely. The coefficients on the three adjusted deficit ratios are larger and more significant than in the univariate regressions, as in the simulations, due to the removal of downward bias caused by the omission of *BDR* in Panel A. We have seen using simulated data that the coefficient on $AdjDefRatio_1$ provides an accurate estimate of underlying *b*, controlling for *BDR*. Using our sample data, the coefficient on $AdjDefRatio_1$ is 0.961 ($t = 4.82$), implying that *b* is close to 1.0. The coefficient on $AdjDefRatio_3$ is 0.847 ($t = 5.40$), implying a slightly lower estimate of *b*. The similarity of the two

[29] We do not include any other control variables or FE in Table 5 to preserve direct comparability with the results of the simulated regressions.

TABLE 5 Relation between discount rate and deficit ratio

Panel A. Specification: $DR_{it} = \alpha + \beta(\text{a measure of deficit ratio})_{it} + \varepsilon_{it}$				
	(1)	(2)	(3)	(4)
DefRatio	0.535**			
	(2.46)			
$AdjDefRatio_1$		0.807***		
		(4.52)		
$AdjDefRatio_2$			0.709***	
			(5.94)	
$AdjDefRatio_3$				0.695***
				(4.98)
Constant	3.587***	3.248***	3.358***	3.376***
	(14.16)	(15.87)	(24.50)	(21.10)
N	2538	2538	2538	2538
R^2	0.037	0.128	0.154	0.134
Panel B. Specification: $DR_{it} = \alpha + \beta(\text{a measure of deficit ratio})_{it} + \gamma BDR_{it} + \varepsilon_{it}$				
	(1)	(2)	(3)	(4)
DefRatio	0.562**			
	(2.44)			
$AdjDefRatio_1$		0.961***		
		(4.82)		
$AdjDefRatio_2$			0.853***	
			(6.35)	
$AdjDefRatio_3$				0.847***
				(5.40)
BDR	0.843***	0.884***	0.896***	0.893***
	(37.36)	(55.48)	(62.23)	(58.02)
Constant	−0.025	−0.693***	−0.624***	−0.606***
	(−0.10)	(−2.77)	(−3.50)	(−2.95)
N	2538	2538	2538	2538
R^2	0.574	0.713	0.755	0.731

Note: Regression results using sample data. The variables are defined in Table 4. The dependent variable is *DR*. In Panel A, the explanatory variable is a measure of deficit ratio. In Panel B, *BDR* is added as a control variable. *t*-statistics in brackets are calculated using standard errors clustered at scheme level.
***, ** and * represent significance at the 1%, 5% and 10% levels.

coefficients is as expected, in view of the results from simulated data generated assuming the same average duration as in our sample.

Table 6 shows the results of multivariate regressions, including *BDR* and the other control variables, and year FE. Pension-related variables are in the first four rows, and company variables are in the remainder. The coefficients on the three adjusted deficit ratios are all close to their values in Table 5, Panel B, and they remain significant at the 1% level. For example, the coefficient on $AdjDefRatio_1$ is 0.908 ($t = 4.48$) in Table 6, compared with 0.961 ($t = 4.82$) in Table 5. These results show that our estimates of underlying *b* are little affected by the inclusion of further control variables.

TABLE 6 Relation between discount rate and deficit ratio, with all control variables

	(1)	(2)	(3)	(4)
DefRatio	0.461**			
	(2.15)			
AdjDefRatio$_1$		0.908***		
		(4.48)		
AdjDefRatio$_2$			0.820***	
			(5.85)	
AdjDefRatio$_3$				0.806***
				(4.98)
BDR	0.472***	0.626***	0.696***	0.648***
	(6.97)	(10.64)	(11.94)	(11.40)
Duration	−0.067***	−0.025	−0.013	−0.020
	(−3.29)	(−1.47)	(−0.85)	(−1.25)
Deficit/MV	0.012			
	(0.13)			
AdjDeficit$_1$*/MV*		−0.025		
		(−0.48)		
AdjDeficit$_2$*/MV*			0.039	
			(0.38)	
AdjDeficit$_3$*/MV*				0.006
				(0.05)
RelSize	0.066	0.061*	0.051	0.055
	(1.52)	(1.77)	(1.56)	(1.61)
Leverage	0.489*	0.258	0.175	0.197
	(1.81)	(1.25)	(0.96)	(0.99)
ROA	0.451	0.272	0.237	0.232
	(1.31)	(0.95)	(0.82)	(0.74)
Z-score	0.034	0.027	0.018	0.022
	(1.12)	(1.01)	(0.75)	(0.86)
Current ratio	0.081	0.013	−0.011	−0.001
	(0.99)	(0.20)	(−0.21)	(−0.02)
Constant	2.942***	0.919	0.461	0.880
	(4.78)	(1.43)	(0.81)	(1.49)
Year fixed effects (FE)	Yes	Yes	Yes	Yes
N	2520	2520	2520	2520
R^2	0.618	0.735	0.772	0.751

Note: Regression results using sample data, including the full set of control variables. The variables are defined in Table 4. *t*-statistics in brackets are calculated using standard errors clustered at scheme level.
***, ** and * represent significance at the 1%, 5% and 10% levels.

TABLE 7 Relation between discount rate and deficit ratio, with firm FE

	(1)	(2)	(3)	(4)
DefRatio	0.221			
	(0.61)			
$AdjDefRatio_1$		1.649***		
		(6.40)		
$AdjDefRatio_2$			1.195***	
			(5.80)	
$AdjDefRatio_3$				1.367***
				(5.69)
BDR	0.559***	0.785***	0.815***	0.788***
	(14.43)	(16.60)	(16.30)	(16.55)
Duration	−0.047***	−0.026**	−0.018*	−0.021**
	(−3.79)	(−2.56)	(−1.93)	(−2.30)
Controls	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
N	2520	2520	2520	2520
R^2	0.793	0.882	0.885	0.885

Note: Regression results using sample data, including the full set of control variables and firm FE. Results for most of the controls are omitted to save space. *t*-statistics in brackets are calculated using standard errors clustered at scheme level.
***, ** and * represent significance at the 1%, 5% and 10% levels.

Our evidence indicates that there is a strong positive relation between the discount rate and the unbiased deficit ratio. We can quantify the effect of this relation on differences in reported deficits (previous studies do not take this step). A value of *b* of 91 bp implies that an increase in the unbiased deficit ratio of +10 percentage points is associated with a *DR* that is larger by $0.10 \times 0.91 = 9.1$ bp. Consider a scheme with an unbiased *PBO* of £123 m and *PAssets* of £100 m. Its deficit is £23 m, and its unbiased deficit ratio is 123/100 = 1.23, the same as the sample average. Let its duration be 17.8 years and its *DR* be 4.25%; these values are also the sample averages. Suppose that the scheme's assets fall to £92.48 m, which means that its unbiased deficit ratio increases by 10 percentage points to £123 m/£92.48 m = 1.33. The scheme will now discount its obligations by 4.25% + discretionary *DR* of 10 points × 0.91 bp = 4.34%. This higher rate reduces its reported *PBO* by $1 - [(1.0425)/(1.0434)]^{17.8} = 1.54\%$ of £123 m = £1.89 m. As a result, the reported increase in the deficit will be £1.89 m/£7.52 m = 25% lower than the unbiased increase.

For schemes that are further in deficit, the reduction in a 10-percentage-point increase in the deficit ratio is larger. Suppose we start with an unbiased *PBO* of £150 m, *PAssets* of £100 m and *DR* of 4.25%. In this case, a fall in *PAssets* of £6.25 to £93.75 m is sufficient to increase the unbiased deficit ratio by 10 percentage points. Reported *PBO* falls by £150 m × 1.54% = £2.31 m, which reduces the increase in the reported deficit by £2.31 m/£6.25 m = 37% of the unbiased increase.

Turning to the control variables, none apart from *BDR* is reliably significant. Most are insignificant at the 10% level in every specification. Hence, we find limited evidence for other potential determinants of the discount rate, beyond the benchmark rate and funding status. Duration has a negative coefficient, as expected, but is only significant (at the 5% level) when firm FE are included (Table 7). Pension deficit in relation to company value is not significant, for any measure *AdjDeficit/MV*, and even when we exclude measures of adjusted deficit ratio, the measures of *AdjDeficit/MV* remain insignificant (not shown). Hence, British companies' choice of discount rate appears to be sensitive to the unbiased pension liabilities in relation to pension assets, but not to the unbiased deficit in relation to company value. This implies that the propensity to manipulate the discount rate is related to the funding status of the scheme rather than how that status relates to the size of the company. Company size has a positive sign but is not significant, which fails to support the idea that larger companies exert more pressure on their actuaries to adopt less conservative assumptions (Anantharaman, 2017).

We find no evidence that upward manipulation of the discount rate is associated with poor financial health. For example, *ROA* and *Z-score* have insignificant coefficients. Our evidence differs from that of several older papers that examine US data, which find that choice of discount rate is linked to the company's financial health (Bodie et al., 1984; Asthana, 1999; Thies & Sturrock, 1988). Our evidence is however consistent with recent results in Billings et al. (2017) for the United Kingdom and Bartram (2018, tab. 4) for the United States that do not show a clear link between the discount rate and the sponsor's health. A possible explanation is that the funding of many schemes was worse by the time of our sample (2009–2018), compared with in the 1980s and 1990s, and that by the 2010s, the health of the scheme itself had become the most pressing consideration when choosing the discount rate. In addition, UK regulation of pensions became stricter following the collapse of the Maxwell Group in 1992 leaving a severely underfunded DB scheme and stricter still following the creation of the Pensions Regulator in 2005. As a result, sample companies in poor financial health probably had less leeway than in earlier years to manipulate pension assumptions.

The results so far are from pooled OLS regressions that estimate both the cross-section and within-scheme relation between *DR* and deficit ratio. Table 7 reports abbreviated results for regressions that include firm FE. Their inclusion means that the coefficients on deficit ratio are determined only by the within-scheme relation between yearly changes in *DR* and deficit ratio. Hence, this specification captures firms' response, in terms of *DR*, to changes over time in their funding status. The pure within-scheme relation is of interest since a natural prediction from the deficit-reduction hypothesis is that if a scheme's unbiased deficit ratio worsens (increases) in a given year, the sponsor will respond with an upward shift in discretionary *DR*.

We find strong support for this prediction. The coefficients on the three adjusted ratios are larger and more significant than in the regressions without firm FE. The coefficient on $AdjDefRatio_1$ is 1.649 ($t = 6.40$), compared with 0.908 ($t = 4.48$) in Table 6, and again, it is the largest of the coefficients on the adjusted ratios. The coefficients on *BDR* are also larger and more significant that in Table 6.

An estimated value of underlying *b* of 165 bp implies that an increase in unbiased deficit ratio of 10 percentage points is associated with an increase in *DR* of 16.5 bp. Consider again a scheme with a duration of 17.8 years, *DR* of 4.25%, unbiased *PBO* of £123 m and *PAssets* of £100 m. An increase in the unbiased ratio of 10 percentage points implies that reported *PBO* will be lower by $1 - [(1.0425)/(1.0442)]^{17.8} = 2.78\%$, a reduction in *PBO* of £3.42 m or 45% of the increase in the deficit that would otherwise have been reported. If we start instead with an unbiased ratio of 1.50 that increases by 10 points, the reduction in reported *PBO* is 67% of the would-be increase in the deficit.

Our result that underlying *b* is positive is consistent with that of Billings et al. (2017) for the United Kingdom, for the period 2005–2009 (and for companies drawn from the largest 350 only). They find a significant negative relation between *DR* and funding status measured by $(PAssets - AdjPBO_3)/AdjPBO_3$, consistent with a positive relation between *DR* and our deficit ratio. Their variables are measured in terms of first differences, which has a similar effect as including firm FE. They do not control for *BDR* but do include month FE, which partly controls for differences in *BDR*. Our evidence regarding deficit in relation to company size is not consistent with Billings et al. They find a significant positive relation between *DR* and measured by $AdjDeficit_3/MV$, whereas we find no significant relation between *DR* and any measure of *AdjDeficit/MV*.

Finally, we investigate whether schemes with an unbiased pension surplus have a tendency to window-dress. We estimate the underlying *b* for such schemes, and thereby ascertain the direction of manipulation of the discount rate. Twenty-four percent of scheme-years are estimated to have an unbiased surplus ($AdjDefRatio_1 < 1$), and the proportion increases from 17% in 2009 to 37% in 2018. Since 85% of schemes in surplus have a negative value for discretionary *DR* (i.e., $DR < BDR$), companies with schemes in surplus tend to overstate pension liabilities and understate the surplus.

TABLE 8 Relation between discount rate and deficit ratio, split by surplus and deficit

	Surplus (1)	Deficit (2)	Surplus (3)	Deficit (4)
AdjDefRatio$_1$	2.006***	0.886***	1.751***	1.640***
	(3.72)	(4.10)	(5.33)	(6.05)
BDR	0.601***	0.639***	0.718***	0.824***
	(7.02)	(9.36)	(14.93)	(14.33)
Duration	−0.003	−0.025	−0.021**	−0.033**
	(−0.20)	(−1.14)	(−2.31)	(−2.05)
Controls	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Firm FE	No	No	Yes	Yes
N	609	1911	609	1911
R^2	0.870	0.683	0.969	0.853

Note: Regression results for the samples of schemes in unbiased surplus ($AdjDefRatio_1 < 1.0$) and deficit ($AdjDefRatio_1 > 1.0$), including the full set of control variables. Results for most of the controls are omitted to save space. *t*-statistics in brackets are calculated using standard errors clustered at scheme level.
***, ** and * represent significance at the 1%, 5% and 10% levels.

But the relation between the size of the surplus and discretionary *DR* is unknown a priori. A positive *b* would imply that companies with a higher surplus operate a surplus-reduction policy, whereby they choose a lower discount rate in order to understate the surplus.

Table 8 shows results for samples of schemes estimated to have either an unbiased surplus ($AdjDefRatio_1 < 1$) or an unbiased deficit ($AdjDefRatio_1 > 1$). The coefficients on $AdjDefRatio_1$ for schemes in surplus are positive and significant and are larger than for schemes in deficit. We infer that companies with schemes in surplus do indeed follow a surplus-reduction policy. This is a different type of window-dressing from using higher discretionary discount rates to understate deficits. Perhaps understatement of the rate when the scheme is in surplus is seen as prudent by executives because it leaves them room to mitigate possible future deterioration in funding status by nudging the discretionary discount rate upward. Another possibility is that a growing surplus is undesirable because it gives rise to "visibility costs" as discussed by Asthana (1999, p.48). He cites evidence that overfunded schemes are more likely to be acquisition targets, and that trustees and members of schemes in surplus are more likely to engage in negotiations to improve member benefits, which can increase the long-term costs of the scheme. In addition, there is evidence that a surplus increases the sponsor's credit rating and reduces its credit spread by less than an equivalent deficit reduces the rating and widens the spread (Cardinale, 2007; Carroll & Niehaus, 1998). This evidence implies that a higher surplus is less beneficial than a lower deficit, in terms of the cost of debt.

6.2 | Companies' use of discretionary discount rates

Our evidence shows that discretionary discount rates are used to reduce reported deficits but also that deficit reduction does not fully account for companies' behavior. First, nearly one quarter of scheme-years are estimated to be in surplus. In most of these cases, the company chooses *DR* below *BDR*, and we find further that *b* is positive for schemes in surplus. Second, most companies with schemes in modest deficit likewise choose *DR* below *BDR*. A majority of companies chooses *DR* above *BDR* only if their scheme has a large deficit—to be precise, if $AdjDefRatio_1$ is above 1.14 (Section 4).

Figure 2 summarizes the relation between the discretionary discount rate and funding status. It shows the proportion of scheme-years with *DR* above *BDR* by decile of $AdjDefRatio_1$. We see that the proportion increases almost monotonically, from 9% for the decile of schemes most in surplus (91% of the surpluses are understated) to 73% for the decile most in deficit (73% of the deficits are understated).

Overall, our evidence on the use of discretionary discount rates shows the following features. (i) Companies with schemes in surplus follow a surplus-reduction policy. They tend to set *DR* below *BDR*, which means that the scheme's pension liabilities are overstated and its surplus is understated. The fact that *b* is positive implies that as unbiased surplus increases, understatement also increases. (ii) Companies with schemes in modest deficit also tend to set *DR* below *BDR*. This means that their deficit is overstated, contrary to the expected use of discretionary *DR* to reduce reported deficits. But a positive *b* for schemes in deficit implies that as deficits increase, overstatement of the deficit reduces (if *DR* is below *BDR*) or understatement increases (if *DR* is above *BDR*). (iii) Schemes with large deficits tend to set *DR* above *BDR*. Hence, their deficits are understated, and positive *b* implies that deficits become increasingly understated as the unbiased deficit increases. The behavior of schemes with large deficits is therefore consistent with the deficit-reduction hypothesis.

7 | CONCLUSION

The valuation of pension liabilities is complex, and a function of a myriad of economic and demographic assumptions. One of the most prominent is the rate used to discount the stream of projected liabilities. Small variations in the discount rate can have a sizeable impact on the reported funding status and perceived health of the pension scheme. It has been claimed that sponsors of schemes with weaker funding status opportunistically exercise discretion to window-dress scheme health. There are two primary empirical complications in the examination of this issue. First, we observe funding status after any opportunistic discretion has been employed, but we are interested in the impact of the unbiased funding status on the discount rate. Second, while the relevant international pension accounting standard IAS 19 appears prescriptive, requiring schemes to use the yield on high-quality corporate bonds of matched duration and currency as the liabilities, there is ambiguity in how to construct the pension discount curve, and thus ambiguity about the true benchmark discount rate.

In response to the first empirical issue, researchers have applied a variety of adjustment methods to proxy for the unbiased funding status. However, these adjustment measures introduce unintended estimation bias that has not previously been documented. We conduct a simulation exercise that allows us to quantify the bias induced by these measures and to document how they vary both with the underlying relationship between unbiased funding status and discount rate (i.e., the underlying *b*) and as a function of the duration profile of sampled schemes. We find that the bias is considerable, inconsistent in sign and dependent both on the underlying *b* and duration. We propose a new adjustment measure that leverages scheme-specific duration estimates that are now in the public domain owing to the disclosure requirements of IAS 19.

In response to the second empirical issue, we develop a method for deriving the currency-specific pension discount curve, with enough granularity for a benchmark to be inferred for any scheme duration. Most of the existing literature omits the benchmark rate as a control variable, and this leads to a downward bias in the estimated impact of unbiased funding status on the discount rate. When we combine the use of our duration-based adjustment measure and inclusion of the benchmark discount rate as a control, our simulations show that we largely eliminate bias in the estimation of underlying *b*, that is, in the estimate of how unbiased funding status affects the choice of the discount rate. This finding is robust to variation in the underlying *b* and scheme duration profiles.

Having documented the superiority of this estimation framework, we conduct an empirical analysis, investigating the choices of the rate of UK pension sponsors across 2500 scheme-years from 2009 to 2018. Across specifications that use the duration-based adjustment, the estimate of underlying *b* ranges from 0.91 to 1.65. For the average sampled scheme, this degree of manipulation reduces an increase in the scheme's reported deficit by 25% to 45%

of the increase that would otherwise have been reported without any manipulation. This is economically significant and implies that sponsors of schemes with weaker funding status upwardly manipulate their discount rates to window-dress the extent of the funding malaise. A wish to understate deficits, however, explains only a minority of sample companies' use of discretionary discount rates. We find that underlying b is positive for schemes in surplus (around 25% of scheme-years), as well as for those in deficit. This implies that companies with schemes in surplus follow a surplus-reduction policy; the understatement of the surplus increases with its size. This is a different type of window-dressing from the use of upward discretionary discount rates to understate deficits.

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APPENDIX

1. Relation between *DR* and $AdjDefRatio_1$ (duration-based adjustment)

Appendix 1 shows (i) that the coefficient β on *UDefRatio* is a downward biased estimate of underlying *b* in a univariate regression without controlling for *BDR*, and (ii) that $\beta = b$ controlling for *BDR*. Since $AdjDefRatio_1 = UDefRatio$ if *BDR* and duration are known, the proofs also apply to $AdjDefRatio_1$.

Consider first the regression of *DR* on unbiased deficit ratio:

$$DR_{it} = \alpha + \beta UDefRatio_{it} + \varepsilon_{it}. \tag{A1}$$

As in equation (11), *DR* is assumed to be determined by:

$$DR_{it} = BDR_{it} + bUDefRatio_{it} + e_i, \tag{A2}$$

where *BDR* and *e* vary independently of *UDefRatio*. Estimated β is then given by:

$$\beta = \frac{\text{cov}\,(DR_{it},\ UDefRatio_{it})}{\text{var}\,(UDefRatio_{it})} = b + \frac{\text{cov}\,(BDR_{it},\ UDefRatio_{it})}{\text{var}\,(UDefRatio_{it})}. \tag{A3}$$

Since *BDR* and *UDefRatio* are negatively correlated (see equation 10), $\beta < b$. It is clear from (A3) that the downward bias is invariant with respect to *b*.

If *BDR* is included as a control in (A1), or if the dependent variable is *DR* – *BDR*, then cov(*BDR*, *UDefRatio*) = 0, and β = *b*. It is easier to show this for the case of *DR* – *BDR* as the dependent variable. In this case, β in (A1) is given by:

$$\beta = \frac{\text{cov}\left(DR_{it} - BDR_{it},\ UDefRatio_{it}\right)}{\text{var}\left(UDefRatio_{it}\right)} = \frac{\text{cov}\left(bUDefRatio_{it} + e_i,\ UDefRatio_{it}\right)}{\text{var}\left(UDefRatio_{it}\right)} = b. \tag{A4}$$

2. Relation between *DR* and $AdjDefRatio_2$ (perpetual-pension adjustment)

Appendix 2 explains the bias in the coefficient on $AdjDefRatio_2$ as an estimate of underlying *b*, if *b* = 0. Starting with the univariate regression (A1), and using $AdjDefRatio_2$ as the explanatory variable, the OLS estimator of β_2 is:

$$\beta_2 = \frac{\text{cov}\left(DR_{it},\ AdjDefRatio_{2it}\right)}{\text{var}\left(AdjDefRatio_{2it}\right)} = \frac{\text{cov}\left(BDR_{it},\ AdjDefRatio_{2it}\right) + \text{cov}\left(e_i,\ AdjDefRatio_{2it}\right)}{\text{var}\left(AdjDefRatio_{2it}\right)}, \tag{A5}$$

and from equations (5) and (10):

$$Adj\ DefRatio_{2it} = DefRatio_{it} \times \frac{DR_{it}}{BDR_{it}} = \frac{(1 + R)^{n_i}\left(1 + x_i k\right) \times \left(BDR_{it} + e_i\right)}{\left(1 + BDR_{it} + e_i\right)^{n_i} \times BDR_{it}}. \tag{A6}$$

First let *BDR* be constant, so cov(*BDR*, $AdjDefRatio_2$) = 0 in (A5). For convenience, let $(1 + R)^{n_i}(1 + x_i k) = a$, and $BDR_{it} = d$. $AdjDefRatio_2$ can then be written as a function of *e*:

$$AdjDefRatio_2 = f\ (e) = \frac{a\,(d + e)}{d(1 + d + e)^n}. \tag{A7}$$

Differentiating with respect to *e* gives:

$$f'\ (e) = \frac{ad^{-1}}{(1 + d + e)^n} - \frac{na\,(d + e)\,d^{-1}}{(1 + d + e)^{n+1}} = \frac{a\left[d^{-1} + 1 - n + d^{-1}(1 - n)\,e\right]}{(1 + d + e)^{n+1}}. \tag{A8}$$

Since *a* and (1 + *d* + *e*) are positive, $f'(e) > 0$. Hence, *f*(*e*) is an increasing function if:

$$d^{-1} + 1 - n + d^{-1}\,(1 - n)\,e > 0, \tag{A9}$$

or if:

$$e_i < \frac{1}{n_i - 1} - BDR_{it}, \tag{A10}$$

and $n > 1$. If condition (A10) is satisfied, cov(*e*, $AdjDefRatio_2$) > 0, and therefore $\beta_2 > 0$ even if *DR* is chosen independently of *UDefRatio* as we have assumed.

Now, let *BDR* be a random variable that is independent of *UDefRatio* and *e*. The upward bias if condition (A10) is satisfied is unaffected. But cov(*BDR*, $AdjDefRatio_2$) is now non-zero. It can be shown that the sign of this covariance could be positive or negative, depending on the size of the variance of *e* in relation to that of *BDR* and also on the value of *n*.

If *BDR* is added as a control in regression (A1), cov(*BDR*, $AdjDefRatio_2$) = 0 in (A5), and β_2 will be upward biased if (A10) is satisfied. In our simulations, the assumed average *BDR* is 4.25%, and values of average duration *n* are between 20 and 10 years. For *n* = 20 (10), (A10) will be satisfied if most values of "non-manipulative discretion" *e* are below 1.01% (6.86%). The term *e* is unobservable, but its standard deviation is 0.70% (Table 1), implying that (A10) will be satisfied for most scheme-years, for all values of assumed duration. We, therefore, expect β_2 to be biased upward, for assumed *b* = 0, which is what we find (Figure 1a).

3. Relation between *DR* and $AdjDefRatio_3$ (fixed-duration adjustment)

Appendix 3 explains the bias in the coefficient on $AdjDefRatio_3$, if $b = 0$. From equations (6) and (9):

$$AdjDefRatio_{3it} = \frac{(1 + R)^{n_i}(1 + x_i k)}{(1 + DR_{it})^{n_i}}\left[1 + 19\left(DR_{it} - BDR_{it}\right)\right] = \frac{(1 + R)^{n_i}(1 + x_i k)}{(1 + BDR_{it} + e_i)^{n_i}}\left[1 + 19e_i\right]. \quad (A11)$$

The OLS estimator of β_3 is as in (A5), except that the deficit ratio is now $AdjDefRatio_3$. Using the same notation as above, this can be written as

$$AdjDefRatio_3 = f(e) = \frac{a(1 + 19e)}{(1 + d + e)^n}. \quad (A12)$$

Let BDR be constant. Differentiating with respect to e gives:

$$f'(e) = \frac{19a}{(1 + d + e)^n} - \frac{an(1 + 19e)}{(1 + d + e)^{n+1}} = \frac{-a\left[19(n - 1)e + n - 19(d + 1)\right]}{(1 + d + e)^{n+1}}. \quad (A13)$$

Since a and $(1 + d + e)$ are positive, $f'(e) > 0$. Hence, $f(e)$ is an increasing function if:

$$-a\left[19(n - 1)e + n - 19(d + 1)\right] > 0, \quad (A14)$$

or if:

$$e_i < \frac{BDR_{it} + 1 - \frac{n_i}{19}}{n_i - 1}, \quad (A15)$$

and $n > 1$. This shows that cov(e, $AdjDefRatio_3$) will be positive if condition (A15) is satisfied.

Now let BDR be a random variable that is independent of $UDefRatio$ and e. The proof regarding cov(e, $AdjDefRatio_3$) will not be affected. $AdjDefRatio_3$ can be written as a function of BDR:

$$AdjDefRatio_3 = f(d) = \frac{a(1 + 19e)}{(1 + d + e)^n}. \quad (A16)$$

Differentiating with respect to d gives:

$$f'(d) = -\frac{(19e + 1)an}{(d + e + 1)^{n+1}} \quad (A17)$$

Since $(d + e + 1)^{n+1}$, and a and n are positive, $f'(d)$ will be negative as long as $19e + 1 > 0$, or $e > -5.3\%$. This implies that cov(BDR, $AdjDefRatio_3$) is negative since it is almost certain in practice that e will exceed -5.3%. Therefore, the sign on β_3 depends on whether the positive cov(e, $AdjDefRatio_3$) or the negative cov(BDR, $AdjDefRatio_3$) dominates.

If BDR is added as a control variable, cov(BDR, $AdjDefRatio_3$) = 0, and β_3 will be upward biased if (A15) is satisfied. For $BDR = 4.25\%$ and duration $n = 20$ (10) years, (A15) will be satisfied if most values of e are below -0.05% (5.74%). Hence, (A15) will clearly be satisfied for $n = 10$ years but not quite for $n = 20$ years. This implies upward bias in β_3 for smaller values of duration, for assumed $b = 0$, which is what we find (Figure 1b).