On the Persistence and Dynamics of Big 4 Real Audit Fees: Evidence from the UK

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Abstract

Despite the huge audit pricing literature, there is a dearth of evidence on the temporal dynamics of audit fee adjustments and the persistence of audit fees. Based on a sample of 76,867 panel observations for a sample of UK companies audited by the Big 4 over the period 1998 to 2012, we employ consistent lagged dependent variable panel estimators to provide new evidence on the persistence and dynamics of real Big 4 audit fees. Contrary to extant research, which assumes that audit fees adjust immediately in a single period, our empirical results indicate that Big 4 real audit fees are persistent, being partly dependent on their previous realisations. We conclude that static audit fee models omit a potentially important temporal dimension of audit pricing behaviour and that further research is warranted into dynamic audit fee models across other jurisdictions.

Keywords: real audit fees, Big 4, partial adjustment, persistence, adjustment speed, dynamic panel estimates, listed and unlisted companies

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1. INTRODUCTION

Despite the huge audit pricing literature, few studies have examined whether lagged audit fees contribute to an explanation of current audit fees given other exogenous determinants. None, to our knowledge, have modelled the influence of lagged real audit fees on current real audit fees. Implicitly, standard audit fee models assume that audit fees are unrelated to fees charged in previous periods and adjust fully to their determinants in a single period.

We provide novel empirical evidence on the dynamics and persistence of Big 4 real audit fees for UK companies. We report dynamic panel estimates for both the listed and unlisted firms, which exhibit high and low[1] Big 4 concentration respectively. We focus on the Big 4 since they are of particular interest as the 'oligopolistic' suppliers of audit services in listed markets. In addition, and following Reynolds and Francis (2001) and Dhaliwal et al. (2014), who restrict their samples to Big 4/5 auditees[2], we confine our estimation to Big 4 audits to reduce the effects of cross-sectional heterogeneity and to avoid increased model complexity[3].

Although there is a voluminous literature on audit fee determinants (e.g., Hay, Knechel and Wong, 2006; and Hay, 2013) relatively few studies use panel data methods[4]. In this context, De Villiers, Hay and Zhang (2014, p. 3) stress that 'the cost behavior of audit fees, especially over time, is not well understood and the examination of audit fee behavior over time can improve our understanding of the

[1] Peel and Makepeace (2012), report that only 8.2% of the audits of a large sample of UK private companies were conducted by Big 4 auditors.
[2] Note also that prior research (Chaney, Jeter and Shivakumar, 2004; and Clatworthy, Makepeace and Peel, 2009) suggests that separate models are appropriate for Big 4 and non-Big 4 auditees, since regression-estimated slope coefficients differ significantly between the two groups.
[3] As described in Section 2, for similar reasons, we also exclude from our sample companies that had switched auditors and/or company type.
[4] Examples where standard (not dynamic) panel estimators are employed to examine audit fee determinants include the studies of Ghosh and Lustgarten (2006) who use an OLS panel estimator; Evans and Schwartz (2014), who apply fixed effects panel methods; and Oxera (2006) who employ both fixed and random effects panel estimators.

audit market'. With a few exceptions[5], extant studies ignore[6] the potential dynamics of audit pricing by assuming that fees are not persistent given the other determinants[7]. Importantly, if a lagged dependent variable is a significant[8] determinant of the dependent variable, but is omitted from the estimated panel model, then it follows that if the remaining explanatory variables are correlated to some extent with the lagged dependent variable, they will be correlated with the model error term. In consequence, biased estimates of coefficients will be observed (Nickell, 1981).

Access to unique data sources for the population of UK limited companies facilitated the collection of a comprehensive vector of explanatory variables and firm-level observations for a relatively long period. Our panel data comprises 76,867 Big 4 auditee observations over the period 1998 to 2012. Using consistent panel data estimation methods, which control for the inherent endogeneity associated with lagged dependent variables, our empirical results indicate that Big 4 real audit fees exhibit persistence (adjust dynamically) in both the listed and unlisted UK corporate markets. We conclude that static audit fee models omit a potentially important temporal dimension of audit pricing and that further research is warranted in other jurisdictions.

The rest of the paper is structured as follows. The next section describes the data and estimation

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[5] Four studies, prepared for/by regulatory authorities, employ the lagged value of audit fees in their models as an additional control variable, but do not address the dynamics of audit pricing (Oxera, 2006; OFT, 2011; PwC, 2012; and Deloitte, 2012). Xie, Cai and Ye (2010) estimate an audit fee model for Chinese listed companies with the aim of employing the residuals in a second step outcome model. They include lagged audit fees as a control variable, but estimate their model using standard OLS, which produces inconsistent (biased) estimates. Indirect evidence that audit fees might exhibit temporal dependency is provided by the research of Doogar, Sivadasan and Solomon (2015). They find that audit fee residuals exhibit persistence, in that the value of the residuals of an audit fee model estimated for the previous period are related to the audit fees charged in the current period. This finding is consistent with the hypothesis that audit fees are serially correlated, providing motivation for employing dynamic panel methods.

[6] In a recent paper, Klumpes, Komarev and Eleftheriou (2016) provide dynamic panel audit fee estimates for a sample of UK insurance companies, based on 175 firm-level observations over the period 1999–2009. However, the dependent variable is expressed as the ratio of audit fees to total assets. Hence, audit pricing dynamics (adjustments) cannot be established or interpreted in a conventional way - especially as size is the principal determinant of audit fees and does not exhibit a unit coefficient relationship in any study we are aware of.

[7] Another recent paper includes the lagged dependent variable in an audit fee model. Abdallah, Goergen and O'Sullivan (2015) examine whether cross-listed companies are charged a fee premium. They do so in the context of highlighting endogeneity issues/remedies in empirical management/business studies. They report that the log of nominal audit fees exhibits persistence, in that the coefficient of lagged audit fees is positive and statistically significant when employing appropriate estimators.

[8] Of course, given its endogenous nature, this assumes that coefficient of the lagged dependent variable would have been significant when using consistent panel estimators.

method. Section 3 presents the results of the empirical study. Section 4 provides a number of potential explanations for our empirical finding that Big 4 real audit fees exhibit persistence. Section 5 contains our concluding comments.

2. DATA AND ESTIMATION METHOD

(i) Data and Variables

Our data cover the population of UK companies and is obtained from two commercial credit reference databases which we had access to over a substantial period. Our initial sample comprises all independent (not held as subsidiary) non-financial companies audited by the Big 4, with data available over the period 1997 to 2012. This creates an unbalanced panel of 64,635 UK companies, representing 340,878 firm-year observations. Of these, 1,723 (11,716 observations) are listed auditees, with the remainder comprising unlisted (private and public) limited companies.

To avoid modelling complexity in our investigation of Big 4 audit pricing dynamics, we exclude all companies which changed auditors or company type (e.g., private to listed), together with beginning-period observations where we are unable to establish whether an auditee switched auditors/company type[9]. Following this screening, the sample decreases to 15,804 companies (88,681 observations), of which 267 (2,492 observations) are listed. This substantial reduction in the number of companies and observations stems from the fact that if a company changed auditor or company type at any point during the estimation period, then the company is removed from the sample. After deleting companies with incorrect/unusual variable values[10], the sample reduces further to 15,582 companies (86,445 observations), of which 263 companies (2,435 observations) are listed. Finally, to facilitate instrumentation, companies are required to have a minimum of 3 panel observations, leading to a final

[9] Because the focus of this note is an investigation of whether Big 4 audit fees follow a dynamic process, the deletion of such companies ensures a homogeneous sample, free from additional modelling complexities (e.g., Reynolds and Francis, 2001, p. 381).

[10] Companies are excluded if they meet one or more of the following criteria: ARTA, CATA or FORSAL outside the range of zero and unity (see Appendix for variable definitions), or pre-tax profit > sales. To avoid potential scanning errors, we also follow Clatworthy et al. (2009, p.148) and exclude companies with total assets or sales below £1,000 and/or with audit fees below £100.

sample of 76,867 observations for 10,345 companies over the period 1998 to 2012, with the figures for listed companies being 2,336 and 227 respectively.

The Appendix provides definitions, labels and summary statistics for the variables employed in the study. The dependent variable is the natural logarithm of real audit fees (LNFEE). With regard to the seminal research of Simunic (1980), and prior studies, our models include a comprehensive range of explanatory variables that focus on auditee size, complexity and risk. Unsurprisingly, corporate size has been found to be the principal determinant of audit fees (Hay et al., 2006, p. 169), though, usually, only one size (typically total assets) variable is used in empirical studies (Hay et al., 2006; and Hay, 2013). Given that company size is the principal driver of audit fees, we employ two size variables, the log of real total assets (LNTA) and the log of real sales (LNSAL). Importantly, in this context, Pong and Whittington (1994, p.1075) stress that audits have two broad dimensions, 'an audit of transactions and verification of assets. The former will be related to turnover and the latter to total assets.'

To capture audit complexity, we use five standard control variables: the ratio of accounts receivable to total assets (ARTA), the ratio of current assets to total assets (CATA), the log of the number (plus 1) of subsidiaries (LNSUB), the ratio of foreign sales to total assets (FORSAL) and whether or not a company received an audit qualification (QUAL). We also include dummy variables for private (PRIV) and unlisted public limited (PLC) companies which we expect to exhibit negative coefficients, given that listed audits are more complex/risky. In line with previous research, we include standard control variables to proxy for audit risk[11]: the ratio of total liabilities to total assets (TLTA), the ratio of net profit before tax to sales (PRSAL) and whether a company is loss-making (LOSS).

We also include a variable (LATE), which denotes that a company filed its accounts after the statutory filing time deadline, thereby incurring penalties. Such companies are expected to be associated with more complex/risky audits (Evans and Schwartz, 2014). Two standard variables indicate whether a company's year-end falls in December or March, referred to as the 'busy' audit period (BUSY), and

[11] Following prior research, to mitigate the impact of outliers, TLTA and RETSAL are winsorized at their 1% and 99% percentiles. We do not winsorize ARTA, CATA or FORSAL since they lie naturally between zero and one.

whether a company is located in London (LOND) to account for the associated higher cost of living. We control for industry differentials (INDUSTRY) via 33 two-digit SIC dummy variables and also include time dummies (TIME) in our model specifications, together with auditor indicator variables, with PwC being the base case.

Finally, since higher industry market share may be associated with higher fees, we follow Evans and Schwartz (2014) and compute the log of the Herfindahl–Hirschman Index (LNHHI), with reference to audit fees. More specifically:

$$HHI_{i,t} = \sum_{j=1}^{N} \left(Market\ share_{i,t,j} * 100\right)^2 \quad (1)$$

Where *i* is an industry sector, *t* is the panel year, *j* is an auditor and *N* is the number of auditors.[12]

(ii) Estimation Method

Due to correlated errors, it is well known that including Y_{it-1} (the lagged dependent variable) in models estimated with standard panel methods (e.g., the studies of Oxera, 2006; and OFT, 2011) result in biased parameter estimates (Nickell, 1981). In this study we employ the Arellano-Bover/Blundell-Bond (ABBB) generalised method of moments (GMM) system estimator for unbalanced panels (see Arellano and Bover, 1995; Blundell and Bond, 1998; and Blundell and Bond, 2000). Though the outcome model is estimated in levels, the ABBB GMM method jointly estimates an elegant system of equations where Y is specified in both differences and levels[13].

As well as employing standard (X) instruments in the difference equation, we include lagged levels of Y as instruments, whereas for the levels equation, we include lagged first differences of Y as instruments[14]. Put simply, the system estimator combines the moment conditions in both levels and

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[12] We employ 33 industrial sectors with reference to two-digit SIC codes and use all the available (11,745) auditors (Big 4 and non-Big 4) on our database to establish yearly audit fee market shares. Separate HHI calculations are made for the pooled (listed and unlisted company) sample, and for the sub-samples of listed and unlisted auditees.

[13] An issue with GMM panel estimators is the proliferation of instruments, which increase in number as a quartic of the available panel time periods (Roodman, 2009), leading to potential overfitting of endogenous lagged dependent variables, together with misleading specification tests. In consequence, we use the method of collapsed instruments as described by Roodman (2009). This results in a smaller instrument matrix. It is implemented using Roodman's (2009) *xtabond2* Stata package.

[14] Examples of the use of dynamic panel estimators in business finance research include the studies of Athanasoglou, Brissimis and Delis (2008), Ozkan (2001), Haynes, Thompson and Wright (2007), Garcia-Teruel and

differences to provide efficient and consistent estimates of panel models that include lagged dependent variables[15].

We report two standard dynamic panel model specification tests in our empirical study. The Sargan (1958) over-identifying restrictions test, tests whether instruments are valid with regard to an absence of correlation between instruments and model errors. Statistically insignificant Sargan statistics are consistent with the instruments being appropriate. In a similar vein, we report tests of whether there is evidence of autocorrelation of model residuals (AR). Serially correlated residuals lead to biased parameter estimates. For GMM models which contain Y_{it-1}, the appropriate test is whether there is evidence of second order, AR(2), serial correlation (Arellano and Bover, 1995). As with the Sargan test, statistically insignificant AR(2) statistics are consistent with an absence of serially correlated errors.

As shown in the Appendix, to account for inflation, and following McMeeking, Peasnell and Pope (2007) and Evans and Schwartz (2014), we express audit fees in real terms (constant 2012 prices). Similarly, the size variables included in our models (sales and total assets) are also computed in real terms, since auditors are expected to adjust real audit fees in response to real changes in auditee size (Evans and Schwartz, 2014). In this context, note that inflation effects disappear in the ratio of two nominal variables, assuming the same variable (deflator) is employed as the numerator for both ratios. However, this is not the case for panel data estimators where variables are in levels. More specifically, if we regressed the level of nominal audit fees on the lagged level of nominal audit fees, *ceteris paribus*, a significant relationship could be observed between the two nominal audit fee variables (even in the limiting case when real audit fees are constant) if inflation is non-zero. We would simply be capturing

Martinez-Solano (2008) and Garcia-Teruel and Martinez-Solano (2010), who examine the dynamics of bank profits, capital structure, executive remuneration, cash holdings and accounts payable respectively.

[15] As a precursor to estimation, we tested for unit root in the logarithm of audit fees employing Harris-Tzavalis tests (Harris and Tzavalis, 1999). We conducted tests on companies with the maximum of 15 time series observations; that is a balanced panel with 1,195 companies. The Harris-Tzavalis tests rejected the null of a unit root (at $p < 0.01$) when drift or trend were included and with/without cross-sectional demeaning (to reduce the impact of possible cross-sectional dependence). We also carried out the same tests on each of the continuous explanatory variables. Again, they rejected the null of a unit root in all series (at $p < 0.01$ in all cases). *A priori*, if audit fees are characterised by a unit root, perhaps the only candidate variable that would cointegrate with audit fees would be a non-stationary corporate size variable. However, as stated, for the current data, the statistical tests reject the hypothesis of a unit root in the size variables employed in the current study.

the persistence in the level of the price index[16].

3. EMPIRICAL RESULTS

To illustrate how our models perform in a standard setting, and to facilitate comparison with extant/future studies, Table 1 reports standard (non-dynamic) fixed effects and OLS panel estimates[17] of LNFEE. As shown in the table, we report models for the pooled (listed and unlisted auditee) sample, together with separate ones for listed and unlisted companies. The table shows that, in general, the primary control variables exhibit their expected signs and that, other than for the listed sample, are highly significant for the OLS specifications. When individual company-specific time invariant effects are controlled for with the fixed effects specifications, some variables lose statistical significance, with the model R^2s also being smaller than those of the OLS models.

Tables 1 to 3 about here

Although models 3 and 4 for listed companies reveal that fewer variables are statistically significant than for their unlisted counterparts (models 4 and 5), or for the pooled specifications (models 1 and 2), the R^2s of the listed models are substantially higher in all cases[18]. Also noteworthy is that, for the OLS estimates, LNHHI is positively and significantly associated with LNFEE in both the listed and unlisted company samples. However, for the pooled sample, LNHHI exhibits a negative and significant coefficient in the fixed and random effects specifications (models 1 and 2). Importantly, and as expected, the coefficients of the corporate size variables (LNTA and LNSAL), are highly significant in all models.

Table 2 presents similar OLS and fixed effects models to those shown in Table 1, but includes LNFEE(t-1) as an additional explanatory variable. We do this for two reasons. Firstly to illustrate the

[16] Note, however, that if the logarithms of nominal audit fees are employed, then the coefficients will be estimated correctly, but only if time dummies are included to allow the intercept to have a different value in each period.
[17] For completeness, we also estimated random effects panel specifications. In terms of coefficients signs and significance levels, the results are broadly in line with the OLS panel estimates. As noted by a reviewer, random effects models require strong assumptions with regard to interpreting estimated parameters.
[18] Chaney et al. (2004, p. 64) also report a relatively low adjusted R^2 (0.57) for their audit fee model of UK private companies; whereas McMeeking, Peasnell and Pope (2006, p. 217) report a substantially higher R^2 (0.82) for their model of UK listed companies. As shown in Table 1, these compare to R^2 s of 0.62 (0.84) for the unlisted (listed) models in the current study.

persistence of audit fees in a standard panel modelling framework; and secondly, to provide useful coefficient bounds with which to gauge the dynamic panel coefficient estimates. As Table 2 reveals, in general, the same patterns of significant explanatory variables as those reported are in Table 1 are repeated, though with the addition of LNFEE(t-1), the model R^2s are substantially higher than their counterparts in Table 1. For the pooled (unlisted) samples, the OLS LNFEE(t-1) coefficients of 0.849 (0.851) indicate that real audit fees are highly persistent, with the coefficient (0.659) for the listed model being considerably lower, but still consistent with substantial persistence in audit fees. The same pattern is evident for the fixed effects models, but the coefficients of LNFEE(t-1) are much smaller.

Although the OLS and fixed effects coefficients of LNFEE(t-1) are biased, they provide useful bounds for interpretation purposes. Specifically, relative to dynamic panel estimates, Blundell and Bond (2000) and Baum (2013) demonstrate that, due to the correlated error structures, the coefficient of a lagged dependent variable is biased downward (upward) for the fixed effects (OLS) panel estimates. As commented by Baum (2013, p. 24), 'given the opposite directions of bias present in these estimates, consistent estimates should lie between these values'. As reported below, all our dynamic panel coefficient estimates of LNFEE(t-1) lie between those of the fixed effects and OLS models, and are therefore consistent on this basis.

To further illustrate the persistence of real audit fees, Table 3 presents simple univariate reduced form panel (autoregressive) regression results where LNFEE is regressed on LNFEE(t-1). For the OLS specification, it reveals a very high degree of persistence, with the LNFEE(t-1) coefficient being similar (around 0.95) across all models, and with the model R^2s also being very high. Unsurprisingly, the table shows that when the dynamic system panel system method is employed - where the lagged value of audit fees is instrumented with its prior values - the model coefficients reduce substantially, though they still exhibit high persistence.

Table 4 about here

Table 4 reports dynamic system panel estimates for the pooled, listed and unlisted samples.

We present two model specifications. As discussed in the next section, models 1, 3 and 5, which include lagged Y and the vector of X_{it} variables, can be interpreted in terms of a Koyck partial adjustment process. Models 2, 4 and 6 extend the Koyck specification to include X_{it-1} variables and correspond to the linear error correction (LEC) model.

As shown in the table, for all models, the AR(2) serial correlation and over-identification of instruments tests are statistically insignificant, implying that the models are well-specified in terms of appropriate instruments and the absence of serially correlated model errors.

Table 4 reveals that, in general, the X_{it} variables exhibit similar signs and significance levels as their counterparts reported in tables 1 and 2. Importantly, the LNFEE(t-1) coefficients are highly significant and positive in all models, indicating that real audit fees are temporally persistent, implying that current real audit fees partially adjust on the basis of their past realisations. We discuss potential explanations for this empirical finding in Section 4.

Interestingly, in terms of goodness of fit, difference between model chi-square tests (Werner and Schermelleh-Engel, 2010), indicate that LEC specifications (models 2, 4 and 6) have significantly higher model chi-squares (at $p < 0.001$ in all cases) than their Koyck counterparts[19] (models 1, 3 and 5). However, whereas the coefficient estimates of LNFEE(t-1) for the Koyck and LEC models are similar for the pooled (models 1 and 2) and unlisted (models 5 and 6) samples, for the listed sample, the LNFEE(t-1) coefficient for the LEC specification (0.419) is larger than that of the Koyck one (0.340). Furthermore, statistical tests for the difference between model coefficients (Clogg, Petkova and Haritou, 1995) indicate that the coefficients of LNFEE(t-1) for the unlisted Koyck (0.546) and LEC (0.545) specifications are significantly larger (at the 5% level in both cases[20]) than those of their listed model counterparts.

4. POTENTIAL EXLANATIONS FOR PERSISTENCE

[19] As shown in Table 4, this is consistent with a number of X_{it-1} variables being statistically significant. In particular, the LEC specification for listed companies (Model 4) reveals that both size variables - LNTA(t-1) and LNSAL(t-1) - exhibit significant negative coefficients. This result is not counterintuitive. If current real audit fees are positively related to size in the long-run, it follows that the coefficient on the lagged size variable in estimates of the LEC equation (see Section 4, equation 5) can be negative.

[20] The p-values for the difference in coefficients values are 0.028 and 0.033 for the Koyck and LEC models respectively.

In this section we outline potential explanations for our empirical finding that Big 4 real audit fees exhibit persistence. However, since our dynamic model estimates are in reduced form we cannot discriminate between them[21].

(i) Two-period Adjustment Model

This model provides the simplest explanation for the persistence of real audit fees. It is based on the premise that lagged audit fees are used as an initial base to determine current ones, modified with reference to the current values of the vector of X variables. There is some qualitative evidence to support this interpretation. The Oxera report prepared for the then Department for Trade and Industry and Financial Reporting Council (Oxera, 2006 p.139) states that 'in the course of the interviews, Oxera has learned that the determination of audit fees for any given year (current year) is often closely related to the agreed audit fee for the last year, and then amended for any new factors'.

Consistent with this, a report on audit pricing submitted by Deloitte LLP to the Competition Commission (Deloitte, 2012, p.12) notes that 'the previous year's audit fee is likely to be the starting point for any negotiation over audit fees'. A similar report by PricewaterhouseCoopers LLP (PwC, 2012, p. 16) comments that 'there are good theoretical reasons to include the previous year's audit fee in our model, not least because last year's fee is often taken as the starting point for discussing and agreeing the following year's fee'. As reported in models 3 (5) in Table 4, the LNFEE(t-1) coefficients imply that 0.340 (0.546) of current real audit fees are determined by their prior values for listed (unlisted) companies.

(ii) Partial Adjustment and Error Correction Models

We next outline two econometric models which can be employed to represent dynamic adjustment processes (persistence) in annual real audit fees; that is, where real audit fees are conjectured to adjust over a number of periods.

[21] This would require a structural model based on specifications of the demand and supply for/of real audit fees and optimal pricing based on their determinants. For a discussion of demand and supply factors in the context of competition in the audit market, see Gerakos and Syverson (2017).

(a) The Koyck Partial Adjustment Model

In this model, real audit fees exhibit a desired, planned or target value, Y_{it}^*, that depends on the explanatory (X_{it}) variables[22]. Adjustment to the desired value is given by the Koyck partial adjustment mechanism

$$Y_{it} - Y_{it-1} = \lambda(Y_{it}^* - Y_{it-1}) \tag{2}$$

where λ is the constant speed of adjustment with restrictions $0 < \lambda \leq 1$.

Consequently we estimate the reduced form equation

$$Y_{it} = (1 - \lambda)Y_{it-1} + \lambda X_{it} + \varepsilon_{it} \tag{3}$$

Where ε_{it} is the error term and where t and i denote time and company observations, Y is the log of real annual audit fees, and X represents the vector of explanatory variables.

From inspection of (3) it is clear that higher adjustment speeds towards desired real audit fees are associated with smaller estimated coefficients for lagged real audit fees. More specifically, $\lambda = 1 - \beta$.

Hence, at the limit, a coefficient of 0 for lagged audit fees ($\lambda = 1 - 0 = 1$) implies the absence of a dynamic (persistent) relationship, with real audit fees adjusting immediately in a single period, being unrelated to their lagged values (as assumed by static audit fee models). In contrast a coefficient of 1 ($\lambda = 1 - 1 = 0$) indicates that real audit fees never adjust into steady state, always being comprised of a constant mark-up of their lagged values[23]. Hence, coefficients of 0.9 (0.1) would, for example, be associated with slow (fast) adjustments speeds. The LNFEE(t-1) coefficients of 0.340 (0.546) for models 3(5) reported in Table 4 for listed (unlisted) companies, imply that adjustment is quicker for listed auditees ($\lambda = 0.660$) than it is for their unlisted counterparts ($\lambda = 0.454$).

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[22] This may occur in response to adjustment costs, imperfect market information (see De Villiers et al., 2014, p. 5), or strategic pricing.

[23] Note that the short-run impact of an X variable is λ times its long-run coefficient. For example, if the short-run coefficient on X in the Koyck formulation is 0.2 and $\lambda = 0.5$, then the long-run coefficient on X is 0.4. We also note that estimates of long-run coefficients can exhibit severe bias. For instance, small biases in estimates of the coefficient on the lagged dependent variable can have large effects on the estimates of the long-run coefficients. This is because the distributions of the estimates of the long-run coefficients are heavy-tailed and complex (see Reed and Zhu, 2015).

(b) The Linear Error Correction Model

The error correction specification is a form of the autoregressive distributed-lag (ARDL) model of type ARDL (1,1). In the linear error correction (LEC) model, annual real audit fees exhibit a value, Y_{it}^{e}, that depends solely on the explanatory variables, X_{it}. Adjustment to this value is given by a linear error correction process where changes in the real audit fees adjust in response to deviations from Y_{it}^{e} in the past period, together with changes in the X_{it} variables (see Alogoskoufis and Smith, 1991). Letting $Y_{it}^{e} = X_{it}$, the LEC is given by

$$Y_{it} - Y_{it-1} = -\mu(Y_{it-1} - X_{it-1}) + \delta(X_{it} - X_{it-1}) + u_{it} \tag{4}$$

The estimated reduced form is therefore given by

$$Y_{it} = \alpha_0 Y_{it-1} + \alpha_1 X_{it} + \alpha_2 X_{it-1} + u_{it} \tag{5}$$

Where u_{it} is the error term and α_i are constant coefficients.

Comparing (3) and (5) we observe that the Koyck adjustment mechanism is nested within the LEC specification[24]. More specifically, if the X_{it} variables are statistically significant, but the X_{it-1} variables are not, then we have evidence of a Koyck (rather than a LEC) adjustment process. The long-run coefficients for the impact of the variables X on Y are given by $\frac{\alpha_1+\alpha_2}{1-\alpha_0}$ or $\frac{\alpha_1+\alpha_2}{\mu}$. The short-run impact of X_{it} is given by α_1. As with the Koyck partial adjustment model, the speed of adjustment is computed as $1-\alpha_0$.

(iii) Spurious State Dependence

It is important to note that a further potential source of endogeneity (heterogeneity) is spurious state dependence. This may arise in the current study if slow moving variables[25] are omitted from the model and are correlated with both the dependent and lagged dependent variables. If this source of endogeneity

[24] We should note that there are other processes that can give rise to lagged dependent variable specifications, though the interpretation would differ. For example, if one or more of the X variables are expectations of future or past values of X - given by (say) an adaptive expectations scheme - we would by substitution obtain a lagged dependent variable, but would also induce a serially correlated error term.

[25] We are grateful to the reviewer for drawing this issue to our attention, especially with regard to slow moving variables.

is not purged via the GMM instrumentation process, it will result in biased estimates of the lagged dependent variable (see Heckman and Borjas, 1980; and Dube, Hitsch and Rossi, 2010).

In this context, Dimitrakopoulos and Kolossiatis (2016, p.1066) provide a helpful typology of sources of persistence in panel data. With regard to the current study, they can be interpreted as follows: (a) true state dependency, which implies that past audit fees are related to current fees; (b) spurious state dependence, which implies that latent heterogeneity is at least partly responsible for the persistence of audit fees; and (c) the dynamic effects of any shocks impacting on audit fees may lead to serial dependence (serial error correlation). In addition, as discussed in Dube et al. (2010), loyalty or relationship commitment can lead to price persistence (true state dependence) via an optimal pricing policy (for a discussion of the impact of commitment/loyalty on auditor-client relationships, together with their association with audit fees, see Levinthal and Fichman, 1988; De Ruyter and Wetzels, 1999; and Farag and Elias, 2011).

5. CONCLUSION

In this paper we extend extant research by modelling Big 4 real audit fees as part of a dynamic adjustment process, where audit fees are linked temporally to their lagged values. Using appropriate dynamic panel methods, we find strong evidence that Big 4 real audit fees of listed and unlisted UK companies are persistent, adjusting temporally on the basis of prior audit fee realisations. We provide a number of potential explanations for this persistence, but acknowledge that our reduced form models cannot differentiate between them. This would require a structural modelling approach that gives cognisance to both supply and demand factors associated with audit fee determinants.

We conclude that static audit fee models omit a potentially important temporal dimension of audit pricing and that further research is warranted across other jurisdictions. Natural extensions to the current study include the impact of auditor switching on audit fees, and the pricing of non-Big 4 audits, in a dynamic framework.

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APPENDIX Variable definitions and summary statistics

Variables†	**Description**	**Mean**	**Median**
Audit fees (£)	Audit fees, RPI adjusted, constant December 2012 prices	85,283	21,950
Total assets (£000)	Total assets, RPI adjusted, constant December 2012 prices	443,825	21,440
Sales (£000)	Turnover, RPI adjusted, constant December 2012 prices	248,825	20,202
LNFEE (£)	Natural logarithm of audit fees, RPI adjusted, constant December 2012 prices	10.074	9.997
LNTA (£)	Natural logarithm of total assets, RPI adjusted, constant December 2012 prices	16.979	16.881
LNSAL (£)	Natural logarithm of sales, RPI adjusted, constant December 2012 prices	16.828	16.821
ARTA	Account receivables to total assets	0.164	0.094
CATA	Current assets to total assets	0.638	0.721
LNSUB	Natural Logarithm (1+ number of subsidiaries)	0.214	0.000
FORSAL	Sales outside UK divided by total sales	0.096	0.000
QUAL	1 if received an audit qualification	0.033	0.000
LISTED	1 if listed company (base case)	0.030	0.000
PLC	1 if unquoted public limited company	0.002	0.000
PRIV	1 if private limited company	0.967	1.000
TLTA	Total liabilities to total assets	0.821	0.743
PRSAL	Profit before tax to sales	-0.037	0.034
LOSS	1 if company reported a loss	0.284	0.000
LATE	1 if company filed accounts after statutory deadline	0.110	0.000
BUSY	1 if company year-end is in December or March	0.722	1.000
DEL	1 if audited by Deloitte	0.198	0.000
EY	1 if audited by Ernst and Young	0.195	0.000
KPMG	1 if audited by KPMG	0.333	0.000
PwC	1 if audited by PwC (base case)	0.275	0.000
LOND	1 if company has registered address in London	0.346	0.000
LNHHI	Logarithm of Hirschman-Herfindahl index	7.106	7.271
INDUSTRY	33 two-digit SIC industry dummy variables	-	-
TIME	Time dummies for each year	-	-

Notes:
The number of observations = 76,867
† For binary variables, zero is coded for remaining observations.

TABLE 1 Fixed effects and OLS panel estimates of audit fees

	All companies		Listed companies		Unlisted companies	
Variables	**(1) Fixed effects**	**(2) OLS**	**(3) Fixed effects**	**(4) OLS**	**(5) Fixed effects**	**(6) OLS**
LNTA	0.117***	0.168***	0.322***	0.309***	0.111***	0.165***
LNSAL	0.190***	0.333***	0.285***	0.277***	0.188***	0.333***
ARTA	0.0494*	0.0988**	0.242	0.203	0.0474*	0.0970**
CATA	-0.0311	0.122***	0.0502	0.0487	-0.0357	0.121***
LNSUB	-0.0200	0.129***	0.0233	0.0551*	-0.0164	0.141***
FORSAL	0.0259	0.234***	0.161***	0.604***	0.00458	0.222***
QUAL	0.0194	0.0899***	0.184***	0.216*	0.0159	0.0840***
PLC	N/A	-1.358***			N/A	-0.555***
PRIV	N/A	-0.812***				
TLTA	0.0416***	0.0257***	0.0625	-0.0371	0.0403***	0.0255***
PRSAL	-0.0535***	-0.126***	-0.0862***	-0.152***	-0.0522***	-0.125***
LOSS	0.0405***	0.123***	0.0260	0.0873	0.0399***	0.124***
LATE	0.0304***	0.193***	-0.00489	0.00214	0.0307***	0.205***
BUSY	0.0286	0.0428***	0.0294	0.131*	0.0318	0.0372**
DEL	N/A	-0.0199	N/A	-0.157*	N/A	-0.0135
EY	N/A	0.0742***	N/A	-0.0501	N/A	0.0803***
KPMG	N/A	-0.107***	N/A	-0.178*	N/A	-0.105***
LOND	N/A	0.177***	N/A	0.422***	N/A	0.170***
LNHHI	-0.0474**	0.0205	0.122	0.208**	0.0370	0.115***
Constant	5.092***	2.054***	0.322***	-0.723	4.570***	0.631**
TIME	✓	✓	✓	✓	✓	✓
INDUSTRY	✓	✓	✓	✓	✓	✓
R^2	0.593	0.660	0.696	0.836	0.565	0.618
N	76,867	76,867	2,336	2,336	74,531	74,531

Notes:

Variables are defined in the Appendix.

Estimated equation:

$LNFEE = \beta_0 + \beta_1 LNTA + \beta_2 LNSAL + \beta_3 ARTA + \beta_4 CATA + \beta_5 LNSUB + \beta_6 FORSAL + \beta_7 QUAL + \beta_8 PLC + \beta_9 PRIV + \beta_{10} TLTA + \beta_{11} PRSAL + \beta_{12} LOSS + \beta_{13} LATE + \beta_{14} BUSY + \beta_{15} DEL + \beta_{16} EY + \beta_{17} KPMG + \beta_{18} LOND + \beta_{19} LNHHI + u$

Time and industry dummy variables are included.

N/A Indicates that a time invariant coefficient cannot be estimated due to perfect collinearity with fixed effects.

***, **,* indicate coefficients are significant at the 1%, 5% and 10% levels respectively (two-tailed tests). Clustered standard errors.

TABLE 2 Fixed effects and OLS panel estimates of audit fees including lagged dependent variable

	All companies		**Listed companies**		**Unlisted companies**	
Variables	**(1) Fixed effects**	**(2) OLS**	**(3) Fixed effects**	**(4) OLS**	**(5) Fixed effects**	**(6) OLS**
LNFEE(t-1)	0.434***	0.849***	0.301***	0.659***	0.437***	0.851***
LNTA	0.0645***	0.0197***	0.284***	0.140***	0.0590***	0.0179***
LNSAL	0.138***	0.0578***	0.201***	0.0715**	0.136***	0.0577***
ARTA	0.0527**	0.0492***	0.270	0.202*	0.0500**	0.0483***
CATA	-0.0200	0.0128*	0.0463	-0.00280	-0.0238	0.0110
LNSUB	-0.000137	0.00595	0.0278	0.0172	0.00443	0.00331
FORSAL	0.0422***	0.0520***	0.167***	0.302***	0.0248*	0.0458***
QUAL	0.00447	0.00778	0.158***	0.0954*	0.00116	0.00695
PLC	N/A	-0.208***			N/A	-0.0561
PRIV	N/A	-0.149***				
TLTA	0.0235***	0.000355	0.0432	-0.00428	0.0222***	-0.000137
PRSAL	-0.0396***	-0.0186***	-0.0686***	-0.0553***	-0.0385***	-0.0175***
LOSS	0.0231***	0.0130***	0.0158	0.00103	0.0229***	0.0139***
LATE	0.0344***	0.0563***	0.000699	0.0189	0.0350***	0.0582***
BUSY	0.0151	0.00426	0.0358	0.0297	0.0169	0.00264
DEL	N/A	0.000857	N/A	-0.0332	N/A	0.00155
EY	N/A	0.0187***	N/A	-0.0309	N/A	0.0207***
KPMG	N/A	-0.0196***	N/A	-0.0613*	N/A	-0.0180***
LOND	N/A	0.0338***	N/A	0.131***	N/A	0.0324***
LNHHI	-0.0218	0.0163	0.112	0.123**	0.0378**	0.0449***
Constant	2.413***	0.233**	-1.782	-0.870*	2.049***	-0.105
TIME	✓	✓	✓	✓	✓	✓
INDUSTRY	✓	✓	✓	✓	✓	✓
R^2	0.867	0.911	0.788	0.911	0.855	0.901
N	65,981	65,981	2,095	2,095	63,886	63,886

Notes:

Variables are defined in the Appendix.

Estimated equation:

$LNFEE = \beta_0 + \beta_1 LNFEE(t\text{-}1) + \beta_2 LNTA + \beta_3 LNSAL + \beta_4 ARTA + \beta_5 CATA + \beta_6 LNSUB + \beta_7 FORSAL + \beta_8 QUAL + \beta_9 PLC + \beta_{10} PRIV + \beta_{11} TLTA + \beta_{12} PRSAL + \beta_{13} LOSS + \beta_{14} LATE + \beta_{15} BUSY + \beta_{16} DEL + \beta_{17} EY + \beta_{18} KPMG + \beta_{19} LOND + \beta_{20} LNHHI + u$

Time and industry dummy variables are included.

N/A Indicates that a time invariant coefficient cannot be estimated due to perfect collinearity with fixed effects.

***, **,* indicate coefficients are significant at the 1%, 5% and 10% levels respectively (two-tailed tests). Clustered standard errors.

TABLE 3 Reduced form panel estimates of audit fees

	All companies		Listed companies		Unlisted companies	
Variable	**(1) OLS Panel**	**(2) Dynamic Panel**	**(3) OLS Panel**	**(4) Dynamic Panel**	**(5) OLS Panel**	**(6) Dynamic Panel**
LNFEE (t-1)	0.950†	0.586†	0.948†	0.519†	0.944†	0.591†
Model R^2 or χ^2	0.904	774.5†	0.882	17.58†	0.894	849.8†
N	65,981	55,095	2,095	1,854	63,886	53,241

Notes:
Constants are omitted.
Estimated equation:
$LNFEE = \beta_0 + \beta_1 LNFEE(t\text{-}1) + u$
The coefficients of models (2), (4) and (6) are estimated using system GMM estimators.
† indicates coefficients are significant at the 1% level (two-tailed tests).

TABLE 4 Dynamic panel model estimates of audit fees

	All companies		Listed companies		Unlisted companies	
Variables	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**
LNFEE(t-1)	0.541***	0.540***	0.340***	0.419***	0.546***	0.545***
LNTA	0.0748***	0.0826***	0.243***	0.400***	0.0716***	0.0716***
LNTA(t-1)		-0.00612		-0.221***		0.00249
LNSAL	0.156***	0.166***	0.161***	0.256***	0.155***	0.162***
LNSAL(t-1)		-0.0125		-0.0878*		-0.00917
ARTA	0.0622***	0.0596***	0.253	0.320	0.0602***	0.0586***
ARTA(t-1)		0.00168		-0.180		0.00254
CATA	0.0441***	-0.00251	0.0342	-0.0629	0.0419***	0.000365
CATA(t-1)		0.0617***		0.0805		0.0568***
LNSUB	0.0468***	0.0385	0.0295	0.0378	0.0484***	0.0539**
LNSUB(t-1)		0.0111		0.000274		-0.00421
FORSAL	0.117***	0.110***	0.413***	0.408***	0.102***	0.0849***
FORSAL(t-1)		0.00619		-0.107		0.0255
QUAL	0.0285**	0.0163	0.148**	0.128**	0.0243*	0.0107
QUAL(t-1)		0.0209		0.0923		0.0228
PLC	-0.614***	-0.627***			-0.219***	-0.240***
PRIV	-0.393***	-0.388***				
TLTA	0.0158***	0.0158**	0.000916	0.153***	0.0150***	0.0126*
TLTA(t-1)		-0.00657		-0.207***		-0.00403
PRSAL	-0.0542***	-0.0542***	-0.0856***	-0.0808***	-0.0532***	-0.0525***
PRSAL(t-1)		-0.00402		-0.0295		-0.00332
LOSS	0.0352***	0.0340***	0.0282	0.0273	0.0354***	0.0336***
LOSS(t-1)		0.0221***		0.000687		0.0236***
LATE	0.0662***	0.0667***	0.0140	0.000186	0.0695***	0.0696***
LATE(t-1)		0.0396***		-0.00866		0.0423***
BUSY	0.0163**	0.00984	0.0714	0.0575	0.0134*	0.00861
BUSY(t-1)		0.00709		0.00272		0.00500
DEL	-0.00635	-0.00573	-0.0985	-0.118**	-0.00354	-0.00237
KPMG	-0.0524***	-0.0505***	-0.117*	-0.127**	-0.0504***	-0.0483***
EY	0.0411***	0.0407***	-0.0434	-0.0429	0.0446***	0.0441***
LOND	0.0901***	0.0892***	0.267***	0.252***	0.0858***	0.0847***
LNHHI	-0.00178	-0.0120	0.180**	0.182**	0.0520***	0.0364*
LNHHI(t-1)		0.0567***		0.0282		0.0531***
Constant	1.055***	0.713***	-1.176*	-1.237*	0.305**	0.0338
TIME	✓	✓	✓	✓	✓	✓
INDUSTRY	✓	✓	✓	✓	✓	✓
Over-identification p-value	0.431	0.502	0.962	0.918	0.446	0.488
AR(2) p-value	0.118	0.123	0.705	0.401	0.135	0.149
Model χ^2	85,018†	91,575†	478.8†	682.7†	68,767†	72,961†
N	55,095	55,095	1,854	1,854	53,241	53,241

Notes:

Variables are defined in the Appendix.

Estimated equations:

Models (1), (3) and (5): $LNFEE = \beta_0 + \beta_1 LNFEE(t\text{-}1) + \beta_2 LNTA + \beta_3 LNSAL + \beta_4 ARTA + \beta_5 CATA + \beta_6 LNSUB + \beta_7 FORSAL + \beta_8 QUAL + \beta_9 PLC + \beta_{10} PRIV + \beta_{11} TLTA + \beta_{12} PRSAL + \beta_{13} LOSS + \beta_{14} LATE + \beta_{15} BUSY + \beta_{16} DEL + \beta_{17} EY + \beta_{18} KPMG + \beta_{19} LOND + \beta_{20} LNHHI + u$

Models (2), (4) and (6): $LNFEE = \beta_0 + \beta_1 LNFEE(t\text{-}1) + \beta_2 LNTA + \beta_3 LNTA(t\text{-}1) + \beta_4 LNSAL + \beta_5 LNSAL(t\text{-}1) + \beta_6 ARTA + \beta_7 ARTA(t\text{-}1) + \beta_8 CATA + \beta_9 CATA(t\text{-}1) + \beta_{10} LNSUB + \beta_{11} LNSUB(t\text{-}1) + \beta_{12} FORSAL + \beta_{13} FORSAL(t\text{-}1) + \beta_{14} QUAL + \beta_{15} QUAL(t\text{-}1) + \beta_{16} PLC + \beta_{17} PRIV + \beta_{18} TLTA + \beta_{19} TLTA(t\text{-}1) + \beta_{20} PRSAL + \beta_{21} PRSAL(t\text{-}1) + \beta_{22} LOSS + \beta_{23} LOSS(t\text{-}1) + \beta_{24} LATE + \beta_{25} LATE(t\text{-}1) + \beta_{26} BUSY + \beta_{27} BUSY(t\text{-}1) + \beta_{28} DEL + \beta_{29} EY + \beta_{30} KPMG + \beta_{31} LOND + \beta_{32} LNHHI + \beta_{33} LNHHI(t\text{-}1) + u$

Time and industry dummy variables are included.

Time invariant variables do not have lagged values due to perfect collinearity.

The coefficients are estimated using system GMM estimator.

***,**,* Indicate coefficients are significant at the 1%, 5% and 10% levels respectively (two-tailed tests). Robust standard errors.

† indicates significant at the 1% level (two-tailed tests).