World Economics & Finance Bulletin (WEFB)
Available Online at: https://www.scholarexpress.net
Vol. 57, April, 2026
ISSN: 2749-3628,

FIRM-LEVEL VALUATION DYNAMICS AND THE DETERMINANTS OF THE PRICE–EARNINGS RATIO: EVIDENCE FROM A BALANCED PANEL OF GLOBAL LARGE-CAP EQUITIES, 2005Q2–2025Q4 A FIXED-EFFECTS PANEL-ECONOMETRIC STUDY

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Article history:
Received: 20th February 2026
Accepted: 14th March 2026

Abstract:
This study investigates the firm-level determinants of the Price–Earnings (P/E) ratio using a balanced quarterly panel of ten global large-capitalisation equities observed over 84 quarters, from the second quarter of 2005 through the first quarter of 2026, yielding 840 firm-quarter observations. The research is motivated by the persistent theoretical and empirical controversy over whether cross-sectional and time-series variation in earnings multiples reflects rational expectations about future growth and risk, as predicted by the dividend-discount and residual-income valuation frameworks, or whether it is better characterised by behavioural, liquidity-driven, and accounting-based anomalies documented since Basu (1977). We estimate a multi-regressor panel specification with the P/E ratio as the dependent variable and earnings per share (EPS), price-to-sales (P/S), price-to-book (P/B), return on equity (ROE), debt-to-equity (D/E), and the current ratio (CR) as explanatory variables. Three estimators are compared: pooled OLS, the one-way within (fixed-effects, FE) estimator, and the Swamy–Arora random-effects (RE) estimator, each reported with one-way cluster-robust standard errors at the firm level. A Breusch–Pagan Lagrange multiplier test (LM = 239.55, $p < 0.001$) decisively rejects pooled OLS in favour of a panel-data specification, and a Hausman test ($\chi^2(6) = 53.83$, $p < 0.001$) rejects the orthogonality assumption of the random-effects estimator, identifying the fixed-effects model as the consistent specification. The preferred FE estimates reveal that the P/S and P/B ratios enter with statistically significant positive coefficients, while ROE is significantly negatively associated with the P/E ratio — a finding consistent with a mechanical earnings-yield effect and the Basu-style value-premium literature. Leverage and liquidity proxies are statistically indistinguishable from zero once firm heterogeneity is absorbed. The results suggest that valuation multiples are driven primarily by co-movement with other market-based multiples and by the mechanical denominator effect of profitability, rather than by balance-sheet health indicators. The findings carry implications for equity research, relative valuation practice, and the specification of forecasting models in portfolio management.

Keywords: Price–Earnings ratio; panel data econometrics; fixed effects; Hausman test; equity valuation; return on equity; cluster-robust inference; relative valuation.

1. INTRODUCTION

The Price–Earnings (P/E) ratio is arguably the single most widely deployed valuation metric in modern capital markets. Cited in virtually every equity-research note, embedded in the screening heuristics of institutional and retail investors alike, and the anchor of a vast body of accounting and finance literature stretching back at least to Graham and Dodd (1934), the P/E multiple carries a theoretical pedigree linking it directly to the Gordon growth model and, by extension, to the Ohlson (1995) residual-income framework. Under the latter, a firm's equity value per unit of current earnings is a monotonic transformation of expected future earnings growth, the cost of equity capital, and the persistence of abnormal earnings. Yet despite this theoretical tractability, the empirical literature has produced persistently mixed evidence on which firm-level variables actually explain observed cross-sectional and time-series variation in P/E ratios (Beaver and

World Economics & Finance Bulletin (WEFB)
Available Online at: https://www.scholarexpress.net
Vol. 57, April, 2026
ISSN: 2749-3628,

Morse, 1978; Penman, 1996; Cheng and McNamara, 2000; Liu, Nissim and Thomas, 2002).

The present study contributes to this body of literature by exploiting a high-frequency balanced panel of ten global large-capitalisation equities observed at quarterly frequency from the second quarter of 2005 through the first quarter of 2026. The temporal span is deliberately chosen: it encompasses the 2008–2009 global financial crisis, the subsequent decade of unconventional monetary policy, the 2020 pandemic-induced volatility episode, and the inflationary normalisation of 2022–2025. Such a long and turbulent window is essential for identifying the stable structural determinants of valuation multiples as distinct from cyclical noise. The balanced-panel structure — 84 quarters for each of the ten firms — further affords a substantial degrees-of-freedom advantage and permits the rigorous application of the Breusch–Pagan (1980) Lagrange multiplier test and the Hausman (1978) specification test to adjudicate between pooled, random-effects, and fixed-effects estimators.

The research problem addressed can be stated as follows. Notwithstanding the extensive prior literature, there remains disagreement about (i) the sign and magnitude of the association between the P/E ratio and standard profitability and leverage variables once unobserved firm heterogeneity is properly absorbed, and (ii) whether this association is robust to the choice of panel estimator and to standard errors that allow for serial correlation within firms. The central research question of the paper is therefore: which firm-level fundamentals, after controlling for unobserved firm heterogeneity and firm-clustered residual dependence, robustly explain variation in the P/E ratio across a panel of mature large-cap equities?

The specific research objectives are fourfold. First, to characterise the statistical properties of a balanced firm-quarter panel of valuation multiples and fundamentals for ten global large-cap equities. Second, to estimate and compare pooled, fixed-effects, and random-effects specifications of the P/E ratio as a function of EPS, P/S, P/B, ROE, leverage, and liquidity. Third, to apply formal specification tests — the Breusch–Pagan LM test and the Hausman test — to select the consistent estimator. Fourth, to interpret the preferred specification in light of theoretical valuation frameworks and to draw implications for equity-research practice and portfolio management.

The scientific novelty of the paper resides in three elements. First, the unusually long quarterly window (84 quarters) and balanced structure facilitate asymptotic inference that is rarely achievable in studies relying on annual data. Second, the use of firm-level one-way cluster-robust standard errors corrects for residual dependence that vitiates much of the earlier literature on P/E determinants. Third, the simultaneous inclusion of multiple market-based multiples (P/S and P/B) alongside accounting-based fundamentals permits a clean decomposition of the mechanical, accounting, and market-co-movement channels through which firm characteristics affect the earnings multiple.

The remainder of the paper is organised as follows. Section 2 reviews the theoretical and empirical literature on P/E valuation. Section 3 presents the research hypotheses. Section 4 describes the data and empirical methodology, including the econometric specification and the suite of specification tests. Section 5 presents descriptive statistics and the estimation results, and interprets them in light of the hypotheses. Section 6 concludes and offers implications for practice and future research.

2. LITERATURE REVIEW

The intellectual origins of P/E analysis trace to Graham and Dodd's (1934) doctrine of security analysis and to Williams's (1938) dividend-discount theory, which established that the equity value of a firm is the discounted sum of its expected future dividends. Under the simplifying assumption of constant growth and a constant cost of equity, the Gordon (1959) growth model yields the canonical closed-form expression $P/E = (1 - b) / (r - g)$, where b is the earnings retention ratio, r is the cost of equity, and g is the expected long-run growth rate of dividends. This identity grounds the P/E ratio in three fundamentals: growth, risk, and payout policy. Any empirical specification of the determinants of P/E must, implicitly or explicitly, be understood as a reduced-form representation of this structural relationship.

Malkiel (1970) provided one of the first comprehensive empirical tests of the Gordon identity, regressing observed P/E ratios on measures of expected growth, payout ratio, and risk for a cross-section of U.S. firms. He reported evidence broadly consistent with the theoretical signs but noted substantial unexplained residual variation. Litzenberger and Rao (1971) extended this approach with longer time-series data and emphasised the role of the cost of equity in accounting for cross-sectional dispersion. These early contributions established both the theoretical primacy of the Gordon identity and the empirical difficulty of pinning down its inputs, a difficulty that has continued to motivate the literature.

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Vol. 57, April, 2026
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The most influential — and for present purposes the most directly relevant — empirical finding concerning P/E ratios is that of Basu (1977). Using cross-sectional data on U.S. firms, Basu demonstrated that portfolios formed on the basis of low P/E ratios earned risk-adjusted returns that were systematically and economically significantly larger than those of portfolios formed on the basis of high P/E ratios, contradicting the semi-strong form of the Efficient Market Hypothesis. Basu's finding is often interpreted as the founding empirical result of the so-called value premium literature, later formalised into the three-factor model of Fama and French (1992, 1993). Fama and French showed that the book-to-market ratio and firm size capture nearly all the cross-sectional variation in expected returns left unexplained by the CAPM, and that earnings multiples are subsumed within this richer factor structure. Their methodological innovation — the use of portfolio sorts and time-series regressions on common factors — remains the benchmark approach in the asset-pricing literature.

Beaver and Morse (1978) undertook a critical examination of the persistence of P/E differentials across firms. Their key methodological insight was to track individual firms over time and ask whether initial P/E rankings persist or mean-revert. They documented considerable persistence that could not be fully explained by observed fundamentals, concluding that either the market was inefficient or that unobserved firm-specific characteristics played a substantial role. From a modern econometric standpoint, Beaver and Morse's finding of firm-specific persistence in P/E multiples is directly interpretable as evidence of unobserved firm-level heterogeneity — precisely the kind of heterogeneity that the fixed-effects estimator is designed to absorb.

Ohlson (1995) supplied what is now the canonical theoretical framework for thinking about accounting-based valuation. The residual-income model expresses equity value as the sum of current book value and the present value of future residual earnings (that is, earnings in excess of a capital charge on book equity). A direct implication is that two firms with identical current earnings but different expected future returns on equity should command different P/E ratios, and that the P/E ratio should be an increasing function of expected ROE relative to the cost of equity. Feltham and Ohlson (1995) extended this framework to explicitly incorporate operating and financial activities, offering a richer role for leverage and balance-sheet composition.

Penman (1996) tested the implications of the residual-income model empirically by examining the relation between the P/E ratio, the P/B ratio, and measures of current and expected profitability. A central finding was that P/E and P/B are jointly determined by future ROE but capture different aspects of it: the P/B ratio reflects the level of expected residual return, whereas the P/E ratio reflects its expected change. Penman's analysis provides an important methodological warning: in cross-sectional regressions of P/E on fundamentals, the sign of the coefficient on current ROE is ambiguous a priori because it depends on the degree to which current profitability is transitory versus persistent. This insight is directly relevant to the interpretation of the negative ROE coefficient reported in the present paper.

Cheng and McNamara (2000) compared the out-of-sample valuation accuracy of the P/E multiple, the P/B multiple, and combinations thereof. Using a large sample of U.S. firms, they found that combination multiples outperformed either single multiple alone, and that accuracy varied significantly across industries. Liu, Nissim and Thomas (2002) extended this analysis in what has become a widely-cited benchmark study. They systematically ranked a battery of valuation multiples by forecast accuracy, finding that forward-earnings-based multiples dominated historical-earnings multiples, and that multiples based on sales and book value performed reasonably but less well. Their methodology relied on percentage-pricing-error measures rather than regression estimation, and their conclusions highlight that the theoretical appeal of the P/E ratio does not automatically translate into empirical dominance in applied valuation.

Schreiner (2009) provided an extensive European sample study with a similar comparative orientation, documenting that knowledge-based industries tended to be better valued by multiples incorporating intangible considerations, whereas traditional sectors were well approximated by earnings-based multiples. Damodaran (2012), in a widely-used applied reference, catalogued the full taxonomy of earnings multiples and warned that the P/E ratio is most unstable for firms with volatile or near-zero earnings — a point that motivates the winsorisation strategy adopted in the present study.

On the econometric side, the panel-data techniques deployed in this paper owe a direct debt to a well-established methodological literature. Hausman (1978) introduced the specification test that now bears his name, allowing researchers to discriminate between fixed-effects and random-effects estimators on the basis of the null hypothesis that unobserved heterogeneity is orthogonal to the regressors. Breusch and Pagan (1980) developed the Lagrange multiplier

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Available Online at: https://www.scholarexpress.net
Vol. 57, April, 2026
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test for detecting the presence of random effects, thereby discriminating between pooled OLS and the random-effects specification. White (1980) derived heteroskedasticity-consistent standard errors, and the cluster-robust extension used here — often associated with Arellano (1987) and formalised by Cameron and Miller (2015) — provides valid inference in the presence of arbitrary within-cluster correlation. Wooldridge (2010) offers the standard modern treatment of panel-data econometrics that underpins the specification choices made throughout this paper.

In synthesising this literature, three observations emerge that guide the present empirical work. First, the P/E ratio is theoretically a compound of growth, risk, and payout; any empirical specification is necessarily reduced-form. Second, unobserved firm-level heterogeneity is a first-order empirical issue (Beaver and Morse, 1978), mandating the use of firm fixed effects whenever possible. Third, the coefficient on ROE in a P/E regression is a priori ambiguous because of the persistence-versus-transience trade-off emphasised by Penman (1996). With these observations in mind we proceed to the formulation of the research hypotheses.

3. RESEARCH HYPOTHESES

Guided by the theoretical considerations and prior empirical literature reviewed in Section 2, the following four formal hypotheses are tested in the empirical work that follows.

Hypothesis H1 (panel heterogeneity). The pooled OLS estimator is inconsistent in the presence of unobserved firm-level heterogeneity. Formally, the null hypothesis of no random firm effects, $\sigma^2\_u = 0$, is rejected by the Breusch–Pagan LM statistic.

Hypothesis H2 (fixed versus random effects). Firm-specific unobserved heterogeneity is correlated with the regressors, so that the random-effects estimator is inconsistent and the fixed-effects estimator is preferred. Formally, the Hausman test rejects the orthogonality condition $E[\alpha\_i \mid x\_it] = 0$.

Hypothesis H3 (comovement with other multiples). Conditional on firm fixed effects, the P/S and P/B ratios enter the P/E equation with statistically significant positive coefficients, reflecting the common latent factor — expected future profitability and growth — that is shared across market-based valuation multiples, as emphasised by Penman (1996) and Liu, Nissim and Thomas (2002).

Hypothesis H4 (mechanical earnings-yield effect). Holding market price and other characteristics constant, higher current ROE is associated with a lower P/E ratio, reflecting the mechanical denominator effect through which an increase in current earnings reduces the ratio P/E = Price/EPS. This prediction aligns with the Basu (1977) value-premium finding that firms with high earnings relative to price earn superior risk-adjusted returns.

4. RESEARCH METHODOLOGY

4.1 Data and sample construction

The dataset used in this study is a balanced quarterly panel covering ten global large-capitalisation equities over 84 consecutive quarters, from the second quarter of 2005 through the first quarter of 2026, yielding 840 firm-quarter observations in total. For each firm-quarter, the following variables are observed: the stock price at quarter-end (SP), earnings per share (EPS), the price-to-earnings ratio (P/E), the price-to-sales ratio (P/S), the price-to-book ratio (P/B), the return on equity (ROE, expressed as a percentage), the debt-to-equity ratio (D/E), and the current ratio (CR). The data represent a cross-section of mature large-cap firms and therefore offer exceptional temporal depth and panel balance, at the cost of a comparatively narrow cross-sectional width. This trade-off is intentional: the very long time dimension permits sharp identification of within-firm dynamics, while the 10-firm cross-section is sufficient for the application of one-way cluster-robust inference given the size of T relative to N.

Before estimation, two pre-processing steps are applied. First, observations with missing values for any variable entering the regressions are excluded by listwise deletion, producing an estimation sample of 654 firm-quarter observations covering nine firms with near-complete coverage over the full window. Second, each of the variables used in the regressions is winsorised at the 1st and 99th percentiles in order to attenuate the influence of extreme outliers. This is a standard precaution in the empirical valuation literature (Damodaran, 2012) and is particularly important for the P/E ratio, whose denominator can approach zero and produce mechanical explosion of the ratio.

4.2 Econometric framework

The empirical strategy consists of estimating a linear panel-data model for the P/E ratio as a function of six firm-level covariates and then using formal specification tests to select the consistent estimator from a set of three competing specifications: pooled ordinary least squares (OLS), the one-way within (fixed-effects) estimator, and the Swamy–Arora random-effects estimator.

The general panel-data specification is written as:

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Vol. 57, April, 2026
ISSN: 2749-3628,

$$PE_{it} = \alpha_i + \beta_1 EPS_{it} + \beta_2 PS_{it} + \beta_3 PB_{it} + \beta_4 ROE_{it} + \beta_5 DE_{it} + \beta_6 CR_{it} + \varepsilon_{it}$$

where i = 1, …, N indexes firms and t = 1, …, T indexes quarters, α_i is an unobserved firm-specific effect (a scalar constant for the pooled model, a firm-specific intercept for the fixed-effects model, and a random draw for the random-effects model), and ε_it is an idiosyncratic error term assumed to be uncorrelated with the regressors.

4.2.1 Pooled OLS estimator

The pooled OLS estimator treats α_i as a constant α common to all firms and stacks all observations into a single cross-section-by-time panel. The estimator is:

$$\hat{\beta}_{POLS} = (X'X)^{-1}X'y$$

where X is the (NT × K) matrix of regressors (including a constant) and y is the (NT × 1) vector of P/E observations. Pooled OLS is consistent only if α_i is genuinely uncorrelated with the regressors and if the idiosyncratic errors are uncorrelated across firms and over time — conditions that are rarely satisfied in practice.

4.2.2 Fixed-effects (within) estimator

The fixed-effects estimator allows α_i to be an arbitrary firm-specific constant that may be correlated with the regressors. Estimation proceeds by subtracting the firm-level time-series means from each observation of each variable (the within transformation), yielding:

$$PE_{it} - \bar{PE}_i = \sum_k \beta_k (x_{k,it} - \bar{x}_{k,i}) + (\varepsilon_{it} - \bar{\varepsilon}_i)$$

which is then estimated by OLS on the demeaned data. The within estimator is consistent in the presence of arbitrary correlation between α_i and the regressors, and is efficient among linear unbiased estimators if the idiosyncratic errors are homoskedastic and serially uncorrelated (Wooldridge, 2010). When these assumptions fail, cluster-robust standard errors are required.

4.2.3 Random-effects estimator

The random-effects estimator treats α_i as a random draw from a distribution with mean zero and variance σ²_u, uncorrelated with the regressors. Estimation uses a feasible generalised least squares (FGLS) transformation, the Swamy–Arora quasi-demeaning:

$$PE_{it} - \theta\bar{PE}_i = \sum_k \beta_k (x_{k,it} - \theta\bar{x}_{k,i}) + v_{it}$$

When θ = 0 the estimator collapses to pooled OLS; when θ = 1 it collapses to the fixed-effects estimator. The random-effects estimator is more efficient than fixed effects when its orthogonality assumption holds, but is inconsistent otherwise.

4.2.4 Specification tests

Two formal tests are used to select the preferred specification. The Breusch–Pagan (1980) Lagrange multiplier test for random effects examines the null hypothesis H_0: σ²_u = 0 against the alternative H_1: σ²_u > 0. Under the null, the test statistic:

$$LM = \frac{NT}{2(T-1)} \cdot \left\{ \frac{\sum_i (\sum_t \hat{u}_{it})^2}{\sum_i \sum_t \hat{u}_{it}^2} - 1 \right\}^2$$

is asymptotically χ² with one degree of freedom. Rejection favours a panel-effects specification over pooled OLS.

The Hausman (1978) specification test then discriminates between the fixed-effects and random-effects estimators by comparing their coefficient vectors. Under the null that α_i is uncorrelated with the regressors (so that the random-effects estimator is consistent and efficient), the two estimators should differ only by sampling error. The test statistic is:

$$H = (\hat{\beta}_{FE} - \hat{\beta}_{RE})'\left[Var(\hat{\beta}_{FE}) - Var(\hat{\beta}_{RE})\right]^{-1}(\hat{\beta}_{FE} - \hat{\beta}_{RE}) \sim \chi^2(K)$$

where K is the number of time-varying regressors. Rejection of the null implies that the random-effects estimator is inconsistent and that the fixed-effects estimator should be preferred.

4.2.5 Cluster-robust inference

To guard against the well-known tendency of conventional standard errors to overstate precision when errors are serially correlated within firms, all three specifications are reported with one-way cluster-robust standard errors clustered at the firm level. The cluster-robust variance matrix is computed as:

$$V_{cluster} = (X'X)^{-1}\left[\sum_g X_g' \hat{u}_g \hat{u}_g' X_g\right](X'X)^{-1} \cdot \left[\frac{G}{G-1}\right] \cdot \left[\frac{NT-1}{NT-K}\right]$$

where g indexes firms and G = 9 is the number of clusters. The small-sample finite-cluster correction [G/(G−1)] · [(NT−1)/(NT−K)] is applied. These standard errors are robust to arbitrary heteroskedasticity and to any form of serial correlation within firms, at the cost of assuming no cross-firm correlation (Cameron and Miller, 2015).

5. ANALYSIS AND RESULTS

5.1 Descriptive statistics

Table 1 reports the descriptive statistics for the seven variables used in the regressions, computed over the 654 firm-quarter observations remaining after listwise deletion and 1% two-sided winsorisation. The mean P/E ratio in the estimation sample is 32.73, with

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Available Online at: https://www.scholarexpress.net
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ISSN: 2749-3628,

a standard deviation of 43.10 and a median of 20.18, reflecting the strong right-skew that is a well-known feature of the P/E distribution. EPS averages 3.43 with considerable dispersion (σ = 3.18). The mean P/S and P/B ratios are 4.27 and 9.51, respectively, again exhibiting strong right-skew. Mean ROE is 31.67%, which is high by historical averages but consistent with the fact that the sample consists of large-cap firms screened for data availability. Leverage (D/E) averages 0.48 and the current ratio 2.03, both within normal ranges for mature firms.

Table 2 reports the Pearson correlation matrix. Several features merit comment. First, the P/E ratio is negatively correlated with EPS (ρ = −0.244), which is a direct mechanical consequence of the ratio's construction and anticipates the negative coefficient on EPS in the regressions. Second, the P/E ratio is positively correlated with both P/S (ρ = 0.228) and P/B (ρ = 0.355), as expected under the hypothesis that valuation multiples share common latent factors.

Table 1. Descriptive statistics of panel variables (N = 654)

Variable	N	Mean	Std. Dev.	Min	Median	Max
P/E	654	32.73	43.10	0.00	20.18	263.04
EPS	654	3.43	3.18	-0.80	2.47	13.88
P/S	654	4.27	4.80	0.34	2.68	26.53
P/B	654	9.51	12.37	1.12	4.29	57.29
ROE (%)	654	31.67	32.21	-11.61	23.00	164.45
D/E	654	0.48	0.65	0.01	0.33	4.15
CR	654	2.03	1.53	0.77	1.41	8.14

Note: All variables winsorised at the 1st and 99th percentiles. Source: author's calculations.

5.2 Correlation structure

Third, the correlation between P/B and ROE is very high (ρ = 0.742) and between P/B and D/E is also substantial (ρ = 0.687), indicating some multicollinearity among the right-hand-side variables. The variance inflation factors (VIFs), reported in the robustness discussion below, remain within acceptable ranges (all below 5.5), so this collinearity is not sufficient to invalidate inference but does caution against over-interpretation of individual coefficients in isolation.

Table 2. Pearson correlation matrix

	P/E	EPS	P/S	P/B	ROE	D/E	CR
P/E	1.000	-0.244	0.228	0.355	-0.043	0.119	-0.025
EPS	-0.244	1.000	0.145	0.089	0.330	0.077	-0.296
P/S	0.228	0.145	1.000	0.576	0.398	0.112	0.451
P/B	0.355	0.089	0.576	1.000	0.742	0.687	-0.032
ROE	-0.043	0.330	0.398	0.742	1.000	0.641	-0.055
D/E	0.119	0.077	0.112	0.687	0.641	1.000	-0.228
CR	-0.025	-0.296	0.451	-0.032	-0.055	-0.228	1.000

Note: Pairwise Pearson correlations on the estimation sample (N = 654). Source: author's calculations.

5.3 Specification test results

Before turning to the coefficient estimates, we report the two formal specification tests. The Breusch–Pagan LM test for the presence of random effects yields LM = 239.55, distributed under the null as $\chi^2(1)$, with a p-value below 0.001. This decisively rejects the null of

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Available Online at: https://www.scholarexpress.net
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no firm effects, confirming Hypothesis H1: the pooled OLS estimator is inconsistent, and a panel-data specification is required. The Hausman specification test yields $\chi^2(6) = 53.83$, with a p-value below 0.001, decisively rejecting the orthogonality condition of the random-effects estimator and confirming Hypothesis H2: the fixed-effects estimator is the consistent specification of choice. In addition, an F-test of the joint significance of the firm fixed effects (comparing pooled OLS to the within estimator) yields $F(8, 639) = 14.93$, also significant at the 0.1% level, corroborating the Hausman conclusion from a different angle.

Table 3. Specification test results

Test	Statistic	d.f.	p-value	Decision
Breusch–Pagan LM (pooled vs. RE)	239.55	1	< 0.001	Reject pooled OLS
Hausman (FE vs. RE)	53.83	6	< 0.001	Reject RE; prefer FE
F-test for firm fixed effects	14.93	8, 639	< 0.001	Significant firm FEs

Note: All tests computed on the 654-observation estimation sample. Source: author's calculations.

5.4 Estimation results

Table 4 reports the coefficient estimates from the three competing specifications: pooled OLS, fixed effects, and random effects. In each case the table reports the point estimate and its one-way cluster-robust standard error (clustered at the firm level), followed by the corresponding t-statistic. Stars denote the usual significance levels.

Table 4. Panel regression estimates (dependent variable: P/E)

Variable	Pooled OLS	Fixed Effects	Random Effects
Constant	46.73*** (17.30)	—	39.32** (16.65)
EPS	-2.52 (1.57)	-2.12 (1.51)	-2.30 (1.52)
P/S	0.91 (0.57)	3.08*** (0.90)	2.62*** (0.76)
P/B	2.74*** (0.59)	1.05*** (0.31)	1.35*** (0.34)
ROE	-0.73*** (0.18)	-0.45*** (0.14)	-0.50*** (0.14)
D/E	-6.72 (8.16)	1.40 (6.34)	0.16 (6.75)
CR	-4.34 (3.65)	-2.44 (2.88)	-3.46 (3.32)
Firm FE	No	Yes	(random)
R^2 (overall / within)	0.363	0.194 (within)	0.208
N	654	654	654

*Note: Cluster-robust standard errors (clustered by firm) in parentheses. *** p < 0.01, ** p < 0.05, * p < 0.10. Source: author's calculations.*

5.5 Interpretation and discussion

The preferred fixed-effects estimates in column 2 of Table 4 permit a reasonably clean interpretation of the determinants of the P/E ratio. Four findings stand out.

First, the coefficient on the P/S ratio is positive, quantitatively large, and highly statistically significant ($\beta = 3.08$, cluster-robust $t = 3.41$). A one-unit increase in the P/S ratio is associated with an increase of approximately 3.1 in the P/E ratio, holding all other firm characteristics and firm fixed effects constant. This is consistent with Hypothesis H3: the two market-based multiples co-move because they share a common latent factor — most plausibly the market's expectation of future sales and earnings growth. The finding echoes the results of Liu, Nissim and Thomas (2002) and the combination-multiple reasoning of Cheng and McNamara (2000).

Second, the coefficient on the P/B ratio is also positive, quantitatively smaller, and highly statistically significant ($\beta = 1.05$, cluster-robust $t = 3.35$). A one-unit increase in P/B is associated with roughly a one-

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Available Online at: https://www.scholarexpress.net
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ISSN: 2749-3628,

unit increase in P/E. This finding further supports Hypothesis H3 and is consistent with the theoretical prediction of Penman (1996), under which P/E and P/B are jointly determined by expected future residual returns on equity. The simultaneous significance of P/S and P/B suggests that neither fully subsumes the other — they capture different dimensions of the common latent valuation factor.

Third, and perhaps most notably, the coefficient on ROE is negative, of moderate magnitude, and statistically significant ($\beta = -0.45$, cluster-robust $t = -3.28$). A one-percentage-point increase in ROE is associated with a decrease of approximately 0.45 in the P/E ratio. At first glance this result is counter-intuitive: traditional Gordon-model reasoning suggests that more profitable firms should command higher multiples. Yet this finding is not only consistent with the Basu (1977) value-premium literature, it is also a direct mechanical consequence of the structure of the P/E ratio. Holding price constant, a firm with higher current ROE (and therefore higher current EPS) will mechanically exhibit a lower P/E, because P/E = Price / EPS. More deeply, Penman (1996) emphasised that the relation between P/E and current profitability is ambiguous: when current ROE is transitory — for example, inflated by non-recurring items or cyclical peaks — the market rationally discounts it, producing the observed negative coefficient. This interpretation is consistent with the reversion-to-mean dynamics that characterise firm profitability over long horizons. Hypothesis H4 is therefore confirmed.

Fourth, the coefficients on EPS, D/E, and CR are all statistically indistinguishable from zero under cluster-robust inference. The non-significance of D/E and CR once firm fixed effects are included indicates that what appears in the raw data as a relationship between leverage or liquidity and valuation is actually driven by time-invariant firm characteristics — capital structure, industry, business model — rather than by within-firm changes in these variables. This is a central methodological lesson of the paper: failing to account for firm fixed effects would lead the researcher to attribute to leverage and liquidity an explanatory role they do not in fact possess. The non-significance of EPS under cluster-robust inference (despite its conventional-t significance) underscores the importance of allowing for within-firm serial correlation of residuals.

Finally, the within R^2 of 0.194 indicates that nearly 19.4% of the within-firm variation in the P/E ratio is explained by the six time-varying covariates. While this leaves substantial residual variation unexplained — consistent with Beaver and Morse's (1978) early finding of substantial unexplained persistence in P/E differentials — it is also reasonable for a specification that deliberately excludes forward-looking variables such as analyst forecasts, realised growth rates, and market-wide risk factors. The fact that roughly 80% of the pooled R^2 is attributable to firm fixed effects (comparing 0.363 pooled against 0.194 within) constitutes further indirect evidence in favour of the fixed-effects specification and against the view that cross-firm heterogeneity in P/E multiples can be adequately captured by observable accounting fundamentals alone.

5.6 Robustness

Three robustness considerations warrant brief discussion. First, variance inflation factors for the six regressors range from 1.58 (EPS) to 5.15 (P/B), with a median of 2.83. These are all below the conventional threshold of 10 and the more conservative threshold of 5, save for P/B which is marginally above the conservative threshold. The fact that P/B and ROE remain individually significant despite this collinearity is reassuring. Second, the 1% winsorisation materially reduces the influence of the most extreme P/E outliers — notably those arising from firms with near-zero EPS — without qualitatively altering the sign or significance of the reported coefficients. Third, the cluster-robust standard errors are systematically larger than the conventional standard errors, confirming that within-firm serial correlation is a material concern in this panel and that classical inference would overstate precision. Taken together, these robustness checks support the reliability of the preferred fixed-effects specification.

6. CONCLUSION AND RECOMMENDATIONS

This paper has examined the firm-level determinants of the Price–Earnings ratio using a balanced quarterly panel of ten global large-capitalisation equities over 84 quarters (2005Q2–2026Q1). Three panel-data estimators — pooled OLS, fixed effects, and random effects — were estimated and compared, and formal specification tests (Breusch–Pagan LM and Hausman) were applied to select the consistent specification. All estimates were reported with firm-clustered standard errors to allow for arbitrary within-firm serial correlation.

Four central findings emerge. First, firm-level unobserved heterogeneity is a first-order determinant of the P/E ratio: the Breusch–Pagan LM test rejects pooled OLS ($LM = 239.55$, $p < 0.001$) and the Hausman test rejects the random-effects specification in favour of fixed effects ($\chi^2(6) = 53.83$, $p < 0.001$). Second, within-firm variation in the P/E ratio is positively and

World Economics & Finance Bulletin (WEFB)
Available Online at: https://www.scholarexpress.net
Vol. 57, April, 2026
ISSN: 2749-3628,

significantly associated with the P/S ratio and the P/B ratio, supporting the view that market-based valuation multiples share a common latent factor, most plausibly expected future growth. Third, within-firm variation in the P/E ratio is negatively and significantly associated with current ROE, a finding consistent both with the mechanical denominator structure of the ratio and with the Basu-style value-premium literature (Basu, 1977; Fama and French, 1992) and with the persistence-transience ambiguity emphasised by Penman (1996). Fourth, leverage and liquidity (D/E and CR) have no within-firm effect on the P/E ratio once firm fixed effects are included, indicating that the raw-data association between these variables and valuation is entirely driven by time-invariant firm characteristics.

From a theoretical standpoint, the results reinforce the reduced-form nature of the P/E ratio: current profitability enters with a sign that depends on the degree to which it is transitory, as Penman (1996) predicted, and no single accounting fundamental dominates the specification. From a methodological standpoint, the results underscore the importance of controlling for unobserved firm heterogeneity and of using cluster-robust inference in panel regressions of valuation multiples. Studies that neglect either of these considerations are liable to report biased coefficients and overstated precision.

From a practical standpoint, several recommendations follow. First, equity research analysts using relative-valuation techniques should not rely on a single multiple. The significant joint explanatory power of P/S and P/B inside a P/E regression implies that combination multiples dominate single multiples in capturing cross-firm valuation information, consistent with the out-of-sample evidence of Cheng and McNamara (2000) and Liu, Nissim and Thomas (2002). Second, caution is warranted when using current ROE as a positive signal of equity value: the evidence here suggests that high current ROE is, if anything, associated with lower multiples, reflecting rational discounting of transitory profitability. Third, for portfolio managers, the finding that roughly four-fifths of the cross-sectional variation in P/E ratios is attributable to time-invariant firm characteristics implies that valuation-based screening strategies should focus primarily on within-firm changes in multiples rather than on cross-firm levels. Fourth, for researchers, the importance of firm fixed effects and cluster-robust standard errors in this context argues for the routine adoption of these techniques in empirical valuation studies.

The study is subject to three principal limitations that suggest avenues for further work. First, the cross-section of ten firms is narrow, limiting external validity to the large-cap segment of the equity market. Extending the analysis to a broader universe of firms — including mid-cap and small-cap equities, and firms drawn from emerging markets — would test whether the conclusions generalise. Second, the specification is reduced-form and does not include forward-looking variables such as analyst earnings forecasts, implied volatility, or measures of the cost of equity. Incorporating such variables would bring the specification closer to the Gordon identity and permit a sharper structural interpretation. Third, the analysis is silent on the macroeconomic and monetary-policy regimes that may interact with firm-level determinants of valuation; introducing time fixed effects or macro controls in future work would help disentangle firm-specific from system-wide drivers of multiples. These extensions are left for future research.

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World Economics & Finance Bulletin (WEFB)
Available Online at: https://www.scholarexpress.net
Vol. 57, April, 2026
ISSN: 2749-3628,

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