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ESG Score and Firm Value: The Role of the Institutional Context—A Comparative Analysis of the United States and the European Union

Paolo Maccarrone *  and Renato Della Cioppa

Department of Management, Economics and Industrial Engineering, Politecnico di Milano, 20154 Milan, Italy

* Correspondence: paolo.maccarrone@polimi.it

Abstract

The relationship between ESG performance and firm value is contested and appears to depend on context. This study asks whether that relationship differs between the United States and the European Union, two mature markets that differ primarily in the extent to which sustainability disclosure is mandated. Firm fixed-effects panel regressions are estimated on LSEG (Refinitiv) data for 2016–2024, covering 874 firms and 5848 firm-year observations, with firm value measured by winsorized Tobin's Q in levels. Among U.S. firms, the estimated association is indistinguishable from zero until 2022 and positive and significant in 2023; among EU firms, it is not significant in any year. The joint hypothesis that the regional differentials are zero in every year is rejected and remains rejected under every robustness check, including the logarithm of Enterprise Value, a lagged ESG score and alternative control sets; the test concerns the period as a whole and does not identify individual years. Functional-form criteria retain linearity, and the turning point implied by a quadratic specification cannot be located on the ESG scale. A direct test of equality across sectors is weak. Institutional context is proxied by region; regulatory stringency and market sentiment are not measured, so the mechanism is not tested.

Keywords: ESG; ESG ratings; firm value; institutional theory; market sentiment; cross—regional analysis; sectoral heterogeneity

1. Introduction

In recent years, ESG has evolved from a niche concept of ethical investing to one of the most prominent—and often debated—topics in global finance. Now widely recognised as a framework for integrating environmental, social, and governance factors into business and corporate strategy, ESG has become both a symbol of progress and a source of criticism—praised for its potential to foster sustainable value creation but also challenged for its conceptual ambiguity and inconsistent measurement (Bongermino & Romagnoli, 2025).

Although ESG performance has become a closely scrutinised and financially relevant dimension of corporate assessment, empirical evidence on its relationship with firm value remains inconclusive. While some studies identify a positive relationship (Whelan et al., 2021; Rahat & Nguyen, 2024; Du et al., 2026), others report negligible (Atan et al., 2018) or even negative effects (Barnett, 2007), often varying by sector, region, and methodological approach. Such heterogeneity, as stated by many scholars (Chau et al., 2025; Shin et al., 2023; Lopez-de-Silanes et al., 2020), suggests that ESG's financial relevance is highly context-dependent and mediated by institutional, cultural, and market-specific factors. These



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inconsistencies highlight the need for a deeper understanding of the relationship between ESG and value creation. The effectiveness of ESG strategies appears to depend not only on firm-specific characteristics but also on the broader institutional and market environments in which firms operate. Differences in regulatory frameworks, market sentiment, and behavioural biases—along with structural and sectoral features—shape how investors interpret, implement, and price ESG (Garcia et al., 2017). Within this evolving landscape, sustainability is no longer merely a reputational concern but a contested economic domain in which regulation, market expectations, and corporate strategy converge, yielding diverse outcomes across sectors and regions.

The growing integration of ESG considerations into corporate strategy and financial decision-making has sparked extensive academic debate over their actual contribution to firm value. A substantial body of research supports a positive relationship between ESG performance and corporate financial performance (Aydoğmuş et al., 2022), indicating that sustainable practices can enhance a firm's reputation, foster stakeholder trust, and improve risk management efficiency. The meta-analysis by Friede et al. (2015), synthesising more than 2000 empirical studies, finds that approximately 90% of the cases document a non-negative ESG–CFP relationship, with the majority reporting positive effects. More recent evidence points in the same direction: Zhao et al. (2025) find, for Chinese listed industrial firms, a positive association between ESG performance and corporate financial performance, measured by ROA and ROE, with the effect partly mediated by green technology innovation.

However, a growing number of studies highlight that these effects are neither universal nor stable across contexts. Capelle-Blancard and Petit (2019) emphasise that the financial relevance of ESG varies across sectors, over time, and across geographies, and that a firm's ability to translate ESG engagement into value depends on its capacity to influence stakeholders. Further evidence underscores the critical role of the institutional context in shaping the financial materiality of ESG. Garcia and Orsato (2020) demonstrate that the ESG–financial performance relationship is context-dependent: it is positive and significant in developed economies, but it turns negative in emerging markets, suggesting that institutional maturity and regulatory credibility condition the value relevance of sustainability practices. Candio (2024) finds that, even within Europe—historically a pioneer in sustainability—the impact of ESG on financial performance remains inconsistent, suggesting that regulatory sophistication does not automatically translate into more substantial value effects.

From this perspective, among recent studies, one of the most relevant is that by Chau et al. (2025), which emphasises the moderating role of country-level institutional quality and environmental sustainability, demonstrating that national contexts significantly influence market responses to ESG ratings.

Cross-regional comparisons further reveal that institutional context matters even among economically comparable markets. Other studies demonstrate that U.S. firms exhibit greater sensitivity of market capitalisation to ESG ratings than European firms, indicating that U.S. investors are more sensitive to sustainability signals, likely reflecting a more sentiment-driven, market-oriented environment. Complementary findings reinforce this view, suggesting that ESG generates higher market premiums in less regulated or more investor-driven contexts (Shin et al., 2023; Breuer et al., 2024). In contrast, in highly institutionalised settings, it tends to become a baseline expectation, reducing its signalling power (Lartey et al., 2021). This heterogeneity aligns with institutional and signalling theories (Spence, 1973), which suggest that ESG performance is most value-relevant when perceived as a credible differentiating signal rather than a regulatory obligation.

Beyond regulation, research has increasingly emphasised the behavioural and narrative dimensions of ESG valuation (Baker & Wurgler, 2007). Market and media sentiment can amplify or soften investors' reactions to ESG information (Yu et al., 2023; Saba, 2025). Some authors have also noted that the ESG–firm value relationship may not be strictly linear, with threshold or diminishing-return effects suggesting non-linear, possibly U-shaped dynamics that depend on firm maturity and institutional conditions.

At the same time, a growing body of literature has explored sector-specific ESG dynamics, showing that industry characteristics play a decisive role in mediating the financial relevance of sustainability (Xiao et al., 2025).

A growing body of empirical evidence confirms that the effect of ESG scores on firm value is highly heterogeneous across specific industries. For example, Shrestha et al. (2025) report a strong positive relationship between ESG performance and market valuation in the consumer goods sector, and between ESG and operational performance in the technology sector, suggesting that proximity to the end consumer amplifies stakeholders' positive response to ESG initiatives. These findings are consistent with those by Lev et al. (2010), who show that consumer-facing firms benefit more from social initiatives in terms of sales. In contrast, no significant relationship is observed in the industrial sector, while a negative relationship is found in the healthcare sector, further confirming sectoral variability. Other sectors, such as oil and gas, telecommunications, consumer services, and finance, show positive but more modest effects. These findings align with previous research, which shows that the ESG–CFP relationship is highly dependent on industry context and the type of performance considered (operational vs. market-based) (Feng et al., 2017).

Overall, the literature suggests that the ESG–value relationship is context-contingent, shaped by the interplay between formal institutions and informal market dynamics. This interpretation resonates with institutional theory, which posits that firms' strategic and financial behaviours are socially and institutionally embedded rather than universally determined (Campbell, 2007).

Research Gap and Research Questions

Despite this growing body of evidence, three specific gaps remain. First, prior comparative studies typically infer the role of institutional context from country—or region—level differences in the ESG–value relationship without directly decomposing it into its regulatory and sentiment-based components (Chau et al., 2025; Garcia & Orsato, 2020), leaving the underlying mechanism largely unobserved rather than directly tested. Second, while threshold or non-linear dynamics in the ESG–value relationship have been proposed theoretically, empirical work rarely tests whether the functional form of this relationship—linear versus non-linear—itself differs systematically across institutional settings. Third, sector-level evidence on ESG's financial materiality is typically available from single-region or single-sector studies (Shrestha et al., 2025; Xiao et al., 2025), so comparative evidence on whether the same sector responds differently to ESG performance across institutional contexts remains scarce.

This study addresses these gaps through a comparative panel design that benchmarks the United States and the European Union as two structurally distinct institutional environments, complemented by functional-form diagnostics and sectoral disaggregation. Given the available data, institutional context is operationalised empirically as the transatlantic (U.S.—EU) asymmetry in the ESG–value relationship, used as a composite proxy for the joint configuration of regulatory stringency and market sentiment in each region, rather than as a direct, decomposed measure of the two constructs; this operationalisation choice, and its implications for interpretation, is discussed explicitly in the Conclusions. On this basis, the study articulates three research questions.

RQ1 (transatlantic differential). Does the association between ESG performance and firm value differ systematically between firms headquartered in the European Union and firms headquartered in the United States, and does any such differential persist or evolve over the 2016–2024 period?

RQ2 (functional form). Is the estimated ESG–value relationship in each region adequately described by a linear specification, or do the data favour a non-linear (e.g., quadratic) specification, and does the preferred functional form differ between the two regions? The selected functional form is treated here as a description of the estimated conditional mean: the mapping from functional form to behavioural mechanism is many-to-one in both directions, and is therefore not used to identify a mechanism.

RQ3 (sectoral heterogeneity). Does sectoral heterogeneity in the ESG–value relationship—associated with structural characteristics such as customer proximity, environmental sensitivity and capital intensity—operate consistently across regions, or is it itself moderated by institutional context, such that identical sectors respond differently to ESG performance in the U.S. and the EU?

By jointly addressing these three questions within a unified empirical framework, the study moves beyond documenting a transatlantic asymmetry in the ESG–value relationship toward a more granular understanding of where, how, and under which institutional and sectoral conditions ESG performance and firm value are most closely associated.

2. Theoretical Framework and Hypothesis Development

The empirical design of this study is organised around institutional theory, which holds that organisational practices derive their meaning and value from the regulatory, normative, and cognitive environment in which they are embedded rather than from their intrinsic content alone (Campbell, 2007). Applied to sustainability, the implication is that identical ESG performance need not carry identical informational content for investors in different institutional settings. Where sustainability practices are codified and mandated, they approximate a condition of legitimacy that firms must satisfy in order to operate; where they remain largely voluntary, they retain the character of a discretionary choice that distinguishes one firm from another.

This distinction connects institutional theory to signalling theory (Spence, 1973). A signal is informative to the extent that it is costly and that not all firms send it. Regulation that standardises disclosure raises the proportion of firms reporting comparable information and lowers the discriminating power of any given score; where disclosure is discretionary, the decision to invest in verifiable sustainability performance remains capable of separating firms. The two theories generate the same directional expectation from different premises: the association between ESG performance and firm value should be weaker where ESG reporting is most institutionalised.

The empirical literature supports the premise that institutional context conditions this association. Garcia and Orsato (2020), testing the institutional difference hypothesis directly, find the ESG–performance relationship to be positive in developed economies and negative in emerging economies. Ghoul et al. (2017) show across 53 countries that the value relevance of corporate social responsibility is greater where market institutions are weaker, that is, where the firm’s own conduct substitutes for what institutions do not guarantee. Shin et al. (2023) document that national culture moderates where ESG pays, and Lopez-de-Silanes et al. (2020) report systematic cross-country differences in both ESG disclosure and its valuation. Breuer et al. (2024) add that investor time preferences, themselves institutionally shaped, condition the relationship. Candio (2024) finds the association inconsistent even within Europe. Taken together, these studies establish that the relationship is conditional on context; none of them, however, compare the United

States with the European Union as two mature markets differing principally in the degree to which sustainability disclosure is mandated, which is the comparison this study makes.

Over the observation period, the European Union progressively introduced a mandatory, prescriptive disclosure framework, while the United States retained a predominantly voluntary, investor-driven regime. On the institutional and signalling arguments set out above, ESG performance should therefore retain greater discriminating power in the United States. This yields the first hypothesis.

Hypothesis H1. *The association between ESG performance and firm value differs between firms headquartered in the United States and those headquartered in the European Union, and is stronger in the United States.*

The second hypothesis concerns functional form. Part of the literature argues that the relationship is not linear: Barnett (2007) holds that returns to social responsibility depend on a firm's accumulated stakeholder influence capacity, so that early efforts yield little and the association strengthens only once a firm is credited with genuine commitment, while Zhang (2022) documents that mandatory disclosure changes firm behaviour discontinuously. Both arguments imply convexity and a threshold above which ESG performance begins to be rewarded. Against this, the bulk of the empirical literature on ESG and firm value estimates linear specifications, and Friede et al. (2015) summarise more than two thousand studies without a systematic finding of curvature. The hypothesis is stated in linear form so that the decision rule set out in the Methodology adjudicates between the two positions rather than presuming the outcome.

Hypothesis H2. *The association between ESG performance and firm value is adequately described by a linear specification in each region.*

The third hypothesis concerns sectoral variation. Xiao et al. (2025) show that the ESG–value relationship is industry-heterogeneous; Shrestha et al. (2025) find it strongest in consumer goods and weakest in industrial sectors; Lev et al. (2010) trace the mechanism in consumer-facing firms to sales; Garcia et al. (2017) show that environmentally sensitive industries display systematically different ESG profiles; and Egorova et al. (2022) document the specificity of the technology sector. These studies examine sectors within single institutional settings. Whether sectoral variation is itself conditioned by institutional context is the open question, and it yields the third hypothesis.

Hypothesis H3. *The association between ESG performance and firm value varies across sectors, and the regional difference postulated in H1 is itself sector-dependent.*

Two clarifications govern the interpretation of the tests that follow. First, the hypotheses concern associations. The design compares firms with themselves over time and removes shocks common to all firms in a year, but it does not identify the direction of the relationship, and no causal claim is made at any point. Second, the structural sector characteristics invoked in the literature on H3—customer proximity, environmental sensitivity, capital intensity—are not measured in this study and are not entered as moderating variables; H3 is a hypothesis about the existence of sectoral variation, not about its causes.

3. Materials and Methods

Research Design

The research design links each hypothesis to a specification, a test statistic and a decision rule fixed in advance, so that the mapping from theory to evidence is explicit.

The unit of observation is the firm-year; the panel is unbalanced and covers 2016–2024. Identification rests on within-firm variation: firm fixed effects absorb every time-invariant firm characteristic, including region, sector, listing venue, accounting regime and the firm's average level of ESG performance, while year fixed effects absorb shocks common to all firms in a given year. What identifies the ESG coefficients is therefore the movement of a firm's ESG score around its own mean, relative to the movement of other firms in the same year. Standard errors are clustered at the firm level throughout, since the residuals of a given firm are correlated over time.

Three design choices follow from this and should be stated explicitly, because each bounds what the evidence can support. First, the regional classification is time-invariant and is therefore absorbed by the firm fixed effects; the regional dimension enters only through the interaction of the ESG score with the U.S. indicator, so that what is estimated is the difference between regions in the association, not a regional-level effect. Second, because the ESG coefficients are allowed to vary by year, the primary specification estimates eighteen of them, and the power available for any single year is limited; the hypotheses are accordingly tested by joint restrictions rather than by the significance of individual years. Third, the regional indicator is a compound proxy for institutional context, and the design cannot decompose it; this is discussed at length in the Conclusions.

Table A8 in Appendix A sets out the correspondence between each hypothesis, the model that tests it, the statistic reported and the rule by which the hypothesis is judged. The same rules are applied to the robustness specifications without modification.

This study employs panel data spanning 2016–2024 to investigate the relationship between ESG performance and firm value across institutional contexts. Panel structures are particularly suitable for this analysis, as they combine cross-sectional variation across firms with temporal variation within firms over time. This approach yields more efficient estimators than cross-sectional or time-series designs alone, while enabling control for unobserved time-invariant heterogeneity—such as managerial quality, organisational culture, or strategic orientation—that might otherwise introduce omitted variable bias. By focusing on within-firm changes, fixed-effects specifications relate firm value to time-varying factors such as ESG scores. In contrast, year fixed effects account for common shocks (e.g., macroeconomic events or regulatory shifts) and are widely recognised as one of the most reliable econometric approaches for modelling the complex relationship between ESG and firm value (Dou & Yin, 2024).

The sample consists of 965 publicly listed firms, with 444 headquartered in the European Union and 521 in the United States, ensuring balanced representation for comparative purposes. The headquarters location determines the regional classification, as it reflects the primary institutional, legal, and regulatory environment that shapes firm strategy and ESG practices; firms are accordingly referred to throughout the manuscript as EU firms and U.S. firms based on this headquarters criterion, rather than by the venue on which their shares are listed. Firms were selected from regulated markets to ensure data availability and comparability, with coverage across diverse sectors to capture heterogeneity. Firms were retained where coverage of ESG scores, Enterprise Value, and key financial indicators was sufficiently continuous over the period: firms absent for more than two consecutive years were excluded, while shorter gaps were permitted. The screen therefore does not require complete and continuous coverage, and the resulting estimation panel is unbalanced. All monetary variables—Enterprise Value, market capitalisation, total assets and net debt—are extracted from LSEG Workspace already denominated in United States dollars, so that no currency conversion is performed by the authors and the two regional subsamples are expressed in a common unit; no purchasing-power adjustment is applied, since firm fixed effects absorb time-invariant differences in scale and price level, and year fixed effects

absorb movements common to all firms in a given year. All continuous variables entering the specifications are winsorized at the 1st and 99th percentiles of the analytical sample to limit the influence of outliers; this is particularly important for the accounting ratios, whose raw distributions contain extreme values arising from small or negative denominators. Descriptive statistics for every variable entering the specifications are reported in Table A1, and their correlation matrix, in levels and within firms, in Table A2 (Appendix A).

The estimation sample is confined to the 2016–2024 window, during which LSEG ESG coverage is reliable; firm-year records dated outside this window are excluded prior to estimation. Of the 965 firms meeting the initial data-availability screen, 874 firms (5848 firm-year observations) form the estimation sample used in the fixed-effects regressions reported in Section 4, after firms with insufficient within-firm time variation for fixed-effects estimation and firm-years with missing control-variable data are excluded on a listwise basis. The full sequence of filters leading from the source database to this estimation sample is documented step by step in Table A3 (Appendix A); the observation and firm counts reported there correspond exactly to those of the primary specification in Table A4. Sectoral classification follows the Refinitiv Business Classification (TRBC) system for consistency with the data source. The analysis focuses on five sectors:

- Financials,
- Industrials,
- Consumer Cyclicals,
- Technology,
- Healthcare.

They were chosen for their presence in both regions and varying ESG exposure profiles (e.g., differences in environmental sensitivity, customer proximity, and capital intensity). For brevity, the remainder of the manuscript refers to these five sectors as Finance, Health, Discretionary, Industrial, and Technology; Health and Discretionary correspond, respectively, to the Healthcare and Consumer Cyclical TRBC classifications listed above. Data are sourced exclusively from LSEG Workspace (formerly Refinitiv), which offers standardised longitudinal ESG and financial metrics. Systematic ESG coverage became reliable starting in 2016, aligning with the start of the sample period.

Firm value is measured in two complementary ways. The primary measure is Enterprise Value (EV), defined as market capitalisation plus total debt, minus cash and cash equivalents, which captures the value attributed to the firm independently of its financing. EV is taken in natural logarithms for reasons of scale. In the estimation sample, it ranges over more than six orders of magnitude, from about one million to almost two thousand billion US dollars, and its distribution is strongly right-skewed (skewness 8.2 in levels, 0.2 in logarithms). In contrast, the ESG score is bounded between 0 and 100 and the ratio controls vary within narrow ranges. Estimated in levels, a handful of very large firms would dominate the coefficients; in logarithms they are semi-elasticities, comparable across firms of very different size—the same rationale given for the logarithmic transformation of market capitalisation in a recent study of ESG performance and firm value (Andrioaia et al., 2026). Total assets, which span a similar range (skewness of 7.6 in levels, 0.4 in logarithms), are taken in logarithms for the same reason. Because EV is nonetheless an absolute money measure, and therefore reflects firm scale as well as valuation, the analysis is repeated throughout on Tobin's Q in levels, computed as EV over total assets and winsorized at the 1st and 99th percentiles, which is a scale-free indicator of valuation; the corresponding estimates are reported among the robustness checks in Section 4. Reporting both measures ensures that no conclusion rests on a single operationalisation of firm value.

The primary explanatory variable is the aggregate ESG score from LSEG (Refinitiv), a relative measure ranging from 0 to 100 that evaluates overall environmental, social, and governance performance based on reported data, weighted by industry materiality.

To isolate the effect of ESG and reduce omitted variable bias, the models include a set of control variables commonly used in corporate finance and ESG research:

- Log of Total Assets (proxy for firm size, capturing economies of scale and visibility, given its influence and moderating effect (Pranbandari et al., 2023);
- Debt-to-common equity ratio (leverage, reflecting financial risk and capital structure);
- Return on Assets (ROA, measuring operational efficiency);
- Return on Equity (ROE); and
- Free Float (influencing liquidity and investor diversity).

Because region and sector are time-invariant firm-level attributes, they are absorbed by the firm fixed effects rather than entered as separate additive dummies, which would be perfectly collinear with them; the regional dimension of interest is instead captured explicitly through the $ESG \times US$ interaction terms described below, while year fixed effects control for time-varying common factors. Note that this absorption logic applies to the pooled specifications described in this subsection; the sectoral extensions introduced below take a distinct approach, re-estimating the model separately on sector-specific subsamples rather than including sector fixed effects or sector dummies, as detailed where those specifications are introduced.

Thus, to fulfil the research purpose, a family of different model specifications has been adopted.

The first model, considered as the methodological backbone, assesses whether ESG scores significantly affect Enterprise Value and whether this relationship varies systematically over time and across regions. Specifically, it examines whether the valuation relevance of ESG differs between U.S. and European firms and whether any divergence remains stable or evolves throughout the 2016–2024 period. The specification is:

$$\log(EV_{i,t}) = \alpha_i + \sum_{j=2016}^{2024} \beta_j \cdot (ESG_{i,t} \cdot D_{j,t}) + \sum_{j=2016}^{2024} \gamma_j \cdot (ESG_{i,t} \cdot D_{j,t} \cdot US_i) + \sum_{j=2016}^{2024} \delta_j (D_{j,t}) + X_{i,t} \cdot \theta \quad (1)$$

where the subscript i refers to a specific firm, while the subscript t refers to a specific year. $D_{j,t}$ are year dummies, US_i is a regional dummy equal to 1 for U.S. firms and 0 otherwise, and $X_{i,t}$ is a vector of control variables as described previously. Two coefficients capture the association between the ESG score and firms' Enterprise Value: β measures the association for EU firms, while γ measures the differential between US and EU, so that the overall association for the US is given by $\beta + \gamma$.

The inclusion of multiple interaction layers introduces potential multicollinearity, especially when combined with firm fixed effects and a rich set of controls. This is addressed through the post-estimation analysis of variance inflation factors (VIFs), reported here as a descriptive, secondary diagnostic: conventional VIF thresholds developed for cross-sectional regressions are not directly interpretable once a large number of firm and year fixed effects are absorbed, so the VIF results should be read as indicative rather than as definitive proof that multicollinearity is fully under control. Inference is further strengthened by clustering standard errors at the firm level, which mitigates potential underestimation of standard errors due to within-entity correlation. The model thus balances flexibility and rigour, offering a robust framework for assessing the time-varying, regionally heterogeneous association between ESG performance and firm valuation.

To ensure the robustness and statistical relevance of the findings, two complementary Wald tests are conducted.

The first Wald test examines whether the ESG effect on firm valuation is significantly different between the United States and Europe. In the model specification, the interaction terms γ_j capture the incremental effect of ESG performance for U.S. firms relative to their European counterparts in each year $j \in \{2016, \dots, 2024\}$.

The null hypothesis tests whether this regional differential is consistently zero across all years:

$$\begin{aligned} H_0: \quad & \gamma_j = 0 \quad \forall j \\ & \gamma_{2016} = \gamma_{2017} = \dots = \gamma_{2024} = 0 \end{aligned} \quad (2)$$

This test jointly evaluates whether the ESG coefficients for U.S. firms are statistically indistinguishable from those of European firms over the entire sample period. Rejecting the null hypothesis of the test implies that, in at least some years, the U.S. exhibits a significantly different ESG valuation effect compared to Europe.

The second Wald test is designed to assess whether the difference in the effect of ESG performance on firm valuation between the United States and Europe remains constant over time. In the model, the interaction terms γ_j capture the incremental ESG effect in the U.S. relative to Europe for each year $j \in \{2016, \dots, 2024\}$.

Formally, the null hypothesis tested is:

$$H_0: \quad \gamma_{2016} = \gamma_{2017} = \dots = \gamma_{2024}$$

which is equivalent to the set of linear restrictions:

$$\gamma_{2016} - \gamma_{2017} = 0, \gamma_{2017} - \gamma_{2018} = 0, \dots, \gamma_{2023} - \gamma_{2024} = 0 \quad (3)$$

Rejecting the null hypothesis would imply that the ESG valuation gap between U.S. and European firms varies over time, suggesting a dynamic divergence (or convergence) in how markets price ESG-related factors between the two regions.

Together, these tests validate the consistency of the estimated effects and support the interpretation of regional divergence as an empirically grounded result rather than a random occurrence.

Building on the baseline specification, two sectoral extensions are introduced to account for potential heterogeneity across industries. The first, referred to as the Global Sectoral Model, pools all firms from both regions and estimates independent regressions for each of the five sectors:

$$\log(EV_{i,t}) = \alpha_i + \sum_{j=2016}^{2024} \beta_j^{(s)} \cdot (ESG_{i,t} \cdot D_{j,t}) + \sum_{j=2016}^{2024} \delta_j^{(s)} (D_{j,t}) + X_{i,t} \cdot \theta^{(s)} \quad (4)$$

This strategy isolates the temporal dynamics of the ESG–value relationship within each industry while abstracting from cross-regional institutional differences. The model does not include sectoral dummies. Instead, the dataset is partitioned by sector prior to estimation, and the code is executed iteratively across the five sectors, each time using the corresponding filtered subsample. In fact, the superscript s of each coefficient indicates that the parameter refers to a specific sector, namely the one for which the sample data have been filtered. By estimating the model for each sector individually, this approach enables a comparative analysis of how the valuation relevance of ESG evolves over time across different industries. It is particularly suited to identifying sector-specific trends, such as whether certain industries (e.g., Finance, Technology, Industrials) exhibit stronger or weaker responsiveness to ESG considerations, and whether such responsiveness intensifies or wanes in specific years.

The subsequent second sectoral model, referred to as the Cross-Regional Sectoral Model, adds a layer of complexity to the previous one, since it examines the time-varying effect of ESG performance on firm value across five key sectors—Finance, Health, Discretionary, Industrial, and Technology—with a focus on potential differences between European and U.S. firms.

The empirical model estimated for each sector is as follows:

$$\log(EV_{i,t}) = \alpha_i + \sum_{j=2016}^{2024} \beta_j^{(s)} \cdot (ESG_{i,t} \cdot D_{j,t}) + \sum_{j=2016}^{2024} \gamma_j^{(s)} \cdot (ESG_{i,t} \cdot D_{j,t} \cdot US_i) + \sum_{j=2016}^{2024} \delta_j^{(s)} (D_{j,t}) + X_{i,t} \cdot \theta^{(s)} \quad (5)$$

The only distinction from the Global Sector Model lies in the reintroduction of the $ESG \times US$ regional interaction term (rather than a standalone regional dummy, for the same collinearity reasons discussed above for the primary model). This addition enables the analysis to capture regional differences within each sector, thereby extending the model's scope; because this interaction term is reintroduced within each sector-specific subsample, the same potential collinearity between ESG, the regional interaction, and firm fixed effects noted for the primary model in principle applies here as well, now estimated on the smaller, sector-specific subsamples described above.

Finally, to assess the actual functional nature of the relationship between ESG performance and firm value across the two regions, the analysis introduces a regional functional-form model. To ensure robustness and to address strands of research that question linearity—often positing U or inverted-U shapes—the study complements this benchmark with targeted diagnostics: (i) functional form analysis, with parametric extensions that introduce polynomial terms (ESG^2 and ESG^3), and (ii) a final decisional test to identify the most suitable model. This layered strategy preserves the interpretability of the linear benchmark while rigorously probing for systematic departures from linearity suggested in prior work. The functional form analysis methodology is built on a two-layer strategy, consisting of a preliminary test aimed at identifying potential non-linear trends and a subsequent decisional test aimed at determining the most reliable specification.

The test is explicitly designed to determine which functional form best captures the relationship between ESG performance and firm value. The underlying rationale is that the ESG–value nexus may not be strictly linear; instead, it might exhibit non-linear dynamics, such as diminishing returns, thresholds, or inflection points. To address this, three alternative model specifications are estimated sequentially:

1. Linear model:

$$\log(EV_{i,t}) = \alpha_i + \beta_1 \cdot ESG_{i,t} + X_{i,t} \cdot \theta + \delta_t + \epsilon_{i,t} \quad (6)$$

2. Quadratic model:

$$\log(EV_{i,t}) = \alpha_i + \beta_1 \cdot ESG_{i,t} + \beta_2 \cdot ESG_{i,t}^2 + X_{i,t} \cdot \theta + \delta_t + \epsilon_{i,t} \quad (7)$$

3. Cubic model:

$$\log(EV_{i,t}) = \alpha_i + \beta_1 \cdot ESG_{i,t} + \beta_2 \cdot ESG_{i,t}^2 + \beta_3 \cdot ESG_{i,t}^3 + X_{i,t} \cdot \theta + \delta_t + \epsilon_{i,t} \quad (8)$$

where:

1. β_1 measures the marginal effect of ESG scores on firm value, assuming a linear relationship.
2. β_2 captures the non-linear (quadratic) component, allowing for concavity or convexity in the ESG–value relationship. A negative term would imply diminishing returns to ESG performance, while a positive value would indicate accelerating effects.

3. β_3 adds a cubic component, enabling the model to capture more complex dynamics such as inflection points, where the direction of the ESG–value relationship may change at different levels of ESG performance.

Additionally, $X_{i,t}$ is a vector of firm-level controls, α_i are firm fixed effects, δ_t year-specific shocks, and $\epsilon_{i,t}$ the idiosyncratic error term.

The model selection process follows a systematic decision rule: the significance of the higher-order terms (ESG-quadratic and ESG-cubic) is evaluated at the 10% level as a preliminary, exploratory screen only. This preliminary step does not by itself determine the specification reported in the main results; it is used solely to flag whether curvature is worth investigating further in the subsequent decisional test, described below, which applies the stricter two-criterion rule (5% Wald test combined with a BIC improvement) that ultimately governs model selection. Based on this preliminary 10% screen, the criteria for flagging candidate non-linear forms are as follows:

- If neither term (quadratic nor cubic) is statistically significant, the linear specification is retained;
- If only the quadratic term is significant, the quadratic model is chosen;
- If the cubic term is significant, the cubic model is preferred.

This stepwise approach ensures that the estimated relationship between ESG and firm value reflects the most appropriate functional form, allowing the analysis to uncover potential non-linearities and asymmetries in the ESG–value nexus. The resulting evidence highlights whether the marginal effect of ESG is constant, diminishing, or more complex across different ranges of the ESG score, thereby offering deeper insight into how sustainability performance interacts with firm valuation dynamics in Europe and the United States.

If preliminary diagnostics—consistent with emerging evidence in the literature—indicate statistically meaningful nonlinear components (from the previous functional form test), the final choice between a linear and a nonlinear (e.g., quadratic) model, regarding the primary longitudinal model assessing the cross-regional relationship between ESG and firm value, implemented so far, is made by applying the two-criterion rule below.

Building on the preliminary 10% screen described above, this decisional step ultimately determines the specification reported in Section 4; wherever the two are discussed together, the 5% Wald-test-plus-BIC rule takes precedence. Ultimately, model selection is framed as a cost–benefit trade-off that balances statistical evidence against parsimony. The question is whether the ESG–firm value relation is adequately linear or requires curvature, evaluated through a two-criterion decision rule. The estimating equation, on which the test is based, is fixed according to the functional-form diagnostic; accordingly, the test is run on one of the two candidates following specifications:

Quadratic form:

$$\log(EV_{i,t}) = \alpha_i + \beta_1 \cdot ESG_{i,t} + \beta_2 \cdot ESG_{i,t}^2 + \gamma_1 \cdot (ESG_{i,t} \cdot US_i) + \gamma_2 \cdot (ESG_{i,t}^2 \cdot US_i) + X_{i,t} \cdot \theta + \delta_t + \epsilon_{i,t} \quad (9)$$

Cubic form:

$$\log(EV_{i,t}) = \alpha_i + \beta_1 \cdot ESG_{i,t} + \beta_2 \cdot ESG_{i,t}^2 + \beta_3 \cdot ESG_{i,t}^3 + \gamma_1 \cdot (ESG_{i,t} \cdot US_i) + \gamma_2 \cdot (ESG_{i,t}^2 \cdot US_i) + \gamma_3 \cdot (ESG_{i,t}^3 \cdot US_i) + X_{i,t} \cdot \theta + \delta_t + \epsilon_{i,t} \quad (10)$$

where α_i are firm fixed effects, δ_t are year dummies, and US_i flags U.S. firms (Europe is the reference group). Europe's slope and curvature are, respectively, β_1 and β_2 ; the U.S. deviations are γ_1 and γ_2 , so U.S. effects equal $\beta_1 + \gamma_1$ and $\beta_2 + \gamma_2$.

If cubic terms are included, β_3 (Europe) and $\beta_3 + \gamma_3$ (U.S.) capture the third-order change in curvature, implying an S-shaped profile with a possible inflection point.

To adjudicate between the linear and non-linear representations, two complementary criteria are applied—one tests whether curvature is statistically warranted, and the other evaluates whether any gain in fit justifies the added complexity:

- 1st Criterion: Wald Test with hypothesis

This test asks whether the nonlinear terms contribute at all. A small p -value (e.g., $p < 0.05$, $p < 0.05$) indicates that curvature improves explanatory content; a larger p -value suggests that a linear form suffices.

- 2nd criterion: Bayesian Information Criterion (BIC)

Overall, the decisional rule is based on the combination of the 1st and 2nd criteria introduced. To ensure parsimony, the quadratic specification is adopted only if both conditions hold:

- (i) The Wald test rejects H_0 .
- (ii) The BIC improves meaningfully ($\Delta BIC \leq 0$); otherwise, the linear model is retained. This rule makes explicit whether the added complexity “pays for itself” or merely reflects noise and overfitting, causing collinearity problems.

4. Results

This study employs panel data spanning 2016–2024 to investigate the relationship between ESG performance and firm value across institutional contexts. The analysis proceeds through a sequence of increasingly granular specifications: a primary model assessing time-varying and regionally heterogeneous effects, sectoral extensions that explore industry-level dynamics, and functional form diagnostics that probe the nature of the relationship. This section presents the empirical findings descriptively, focusing on the estimated coefficients, their statistical significance, and the patterns that emerge across specifications. Throughout this section and the Discussion, the reported coefficients are conditional associations between ESG scores and firm value, not estimates of a causal effect. The control variables and the fixed-effects structure are included to compare firms on a like-for-like basis over time, not because the design identifies the direction of the relationship between the two variables. Language describing ESG as affecting, driving or rewarding valuation has accordingly been avoided throughout.

The primary model estimates year- and region-specific associations between ESG scores and Tobin’s Q while controlling for firm-level financial and structural characteristics. It is estimated on an unbalanced panel of 5848 firm-year observations from 874 firms, with an average of 6.7 observations per firm, yielding a within R-squared of 0.093 and a robust F-statistic of 6.13 ($p < 0.001$). The within R-squared is necessarily far lower than in a specification whose dependent variable is an absolute money measure, since a valuation ratio is much less predictable from firm size than Enterprise Value is; the corresponding figure for the alternative log Enterprise Value specification is 0.328. Observations, cluster counts, residual degrees of freedom, within R-squared, and joint-test statistics for every specification reported in the paper are collected in Table A4 (Appendix A), and the estimated coefficients on the control variables are reported in Table A5.

Because the dependent variable is a ratio expressed in levels, the ESG coefficients are differences in Tobin’s Q per point of ESG score: a coefficient of 0.009 indicates that a one-point difference in the ESG score is associated with a difference of 0.009 in Tobin’s Q, which, against a sample mean of 1.67, corresponds to approximately 0.5% of the average valuation ratio.

For European firms, the estimated ESG coefficients are not statistically significant in any year of the sample (see Table 1). They are negative in most of the earlier years, with the largest absolute value in 2018 (-0.0065 , $p = 0.140$), and then move towards zero thereafter without approaching conventional significance at any point. Over the whole observation window, the European estimates are statistically indistinguishable from zero.

Table 1. ESG score and Tobin's Q: regional coefficients and significance (primary model).

Year	EU Coef.	EU s.e.	EU 95% CI	EU p	US Coef.	US s.e.	US 95% CI	US p
2016	−0.003817	0.004481	[−0.012612, 0.004979]	0.394641	−0.003884	0.005583	[−0.014842, 0.007075]	0.486888
2017	−0.002216	0.004571	[−0.011188, 0.006757]	0.628019	0.000122	0.004603	[−0.008913, 0.009157]	0.978864
2018	−0.006486	0.004395	[−0.015111, 0.002139]	0.140335	−0.003492	0.004842	[−0.012995, 0.006011]	0.470993
2019	−0.002080	0.004405	[−0.010725, 0.006565]	0.636941	−0.000998	0.004395	[−0.009625, 0.007629]	0.820394
2020	−0.004549	0.004698	[−0.013769, 0.004672]	0.333162	0.000049	0.004970	[−0.009705, 0.009803]	0.992086
2021	0.001071	0.004720	[−0.008193, 0.010334]	0.820616	0.005751	0.004982	[−0.004027, 0.015530]	0.248647
2022	−0.001234	0.004501	[−0.010068, 0.007600]	0.784003	0.003901	0.003881	[−0.003716, 0.011518]	0.315041
2023	0.001339	0.004654	[−0.007794, 0.010473]	0.773534	0.009177 **	0.004015	[0.001296, 0.017058]	0.022518
2024	−0.001471	0.004847	[−0.010985, 0.008043]	0.761578	0.007299	0.004765	[−0.002053, 0.016652]	0.125929

Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

For U.S. firms, the estimates are likewise indistinguishable from zero from 2016 through 2022 and become positive and significant at the 5% level in 2023 (0.0092, $p = 0.023$). The 2024 estimate is of similar magnitude (0.0073) but does not reach conventional significance ($p = 0.126$), so on this measure the significant association is confined to a single year. A temporal profile of this kind may also reflect changes in the construction, calibration and coverage of the ESG rating over the period, and it is therefore not interpreted here as evidence of a change in valuation practice.

Robustness checks employing the ESG Combined Score, which integrates reported and unreported ESG elements including controversy adjustments, reproduce this pattern (see Table 2). The U.S. estimate is positive and significant at the 10% level in 2023 (0.0068, $p = 0.056$) and indistinguishable from zero in every other year, while among the European estimates only 2018 attains conventional significance, with a negative coefficient (-0.0074 , $p = 0.025$) that is not sustained in the adjacent years. The joint test that the regional differentials are zero in every year continues to reject (chi-square(9) = 18.36, $p = 0.031$).

Table 2. ESG score and Tobin's Q: regional coefficients and significance (primary model with combined score).

Year	EU Coef.	EU s.e.	EU 95% CI	EU p	US Coef.	US s.e.	US 95% CI	US p
2016	−0.004842	0.003144	[−0.011013, 0.001328]	0.123853	−0.008621	0.005530	[−0.019475, 0.002233]	0.119402
2017	−0.002156	0.003177	[−0.008391, 0.004079]	0.497519	−0.002646	0.004052	[−0.010598, 0.005306]	0.513917
2018	−0.007398 **	0.003300	[−0.013874, −0.000922]	0.025214	−0.007439	0.004646	[−0.016557, 0.001679]	0.109694
2019	−0.002319	0.003084	[−0.008371, 0.003734]	0.452326	−0.004228	0.004100	[−0.012274, 0.003819]	0.302710
2020	−0.005016	0.003691	[−0.012260, 0.002228]	0.174491	−0.003545	0.004117	[−0.011625, 0.004534]	0.389366
2021	0.001505	0.003715	[−0.005787, 0.008797]	0.685546	0.003188	0.004217	[−0.005089, 0.011466]	0.449897
2022	−0.000942	0.003096	[−0.007017, 0.005134]	0.761088	0.002174	0.003433	[−0.004564, 0.008912]	0.526803
2023	0.000845	0.003276	[−0.005586, 0.007276]	0.796496	0.006765 *	0.003529	[−0.000161, 0.013690]	0.055549
2024	−0.002342	0.003842	[−0.009882, 0.005198]	0.542267	0.004297	0.004906	[−0.005333, 0.013926]	0.381412

Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Prior to estimating the main models, functional-form diagnostics were conducted on the control variables, applying to each the two-criterion rule set out above. Leverage and free float are linear on both criteria. Firm size and the two profitability measures display curvature: the quadratic terms are significant ($p = 0.036$ for log total assets, $p < 0.01$ for ROA and ROE) and improve the BIC. The linear specification is retained for the control set in the models reported below, for comparability with the empirical corporate finance literature, and the quadratic specification is estimated as a robustness check. Adding the squared size and profitability terms leaves the ESG estimates essentially unchanged—the U.S. coefficient for 2023 moves from 0.0092 to 0.0077 and remains significant at the 5% level, and the joint test on the regional differentials is unaffected ($\chi^2(9) = 18.49$, $p = 0.030$)—so the functional form of the controls is not consequential for the results discussed here.

To assess whether the regional difference is statistically meaningful, two Wald tests were conducted on the estimates of the primary model. The first evaluates the null hypothesis that the ESG–value association is identical in the two regions in every year. It yields a chi-square statistic of 19.60 on nine degrees of freedom with a p -value of 0.021, so the null is rejected at the 5% level. It is important to be precise about what this test does and does not establish: it rejects the hypothesis that the two regional profiles coincide over the period taken as a whole, and it does not identify the individual years in which the two regions differ. The year-by-year estimates are reported in Table 1 and, as noted above, none of the annual regional differentials are individually significant. The joint test is the more informative reading because the specification estimates eighteen separate ESG coefficients and the power available for any single one of them is correspondingly limited.

The second Wald test assesses whether the regional differential is constant over time. The null of a constant differential is rejected ($\chi^2(8) = 16.26$, $p = 0.039$).

The differential is therefore not stable over the period: it is indistinguishable from zero in the earlier years and larger at the end of the sample. This is a statement about the estimated association and its evolution, not about any mechanism producing it.

The sectoral analysis is reported on the logarithm of Enterprise Value rather than on Tobin's Q (Table 3). The reason is one of power rather than of preference: the sector subsamples contain between 803 and 1438 observations, and the within R -squared of the Tobin's Q specification is 0.093 against 0.328 for log Enterprise Value, so the ratio-based specification leaves too little identified variation at the sector level to be informative. The sectoral joint tests are reported in Table A6, and their Tobin's Q counterparts in Table A7; the latter are discussed below wherever they differ from those presented here. The Global Sectoral Model pools firms from both regions and estimates independent regressions for each of the five sectors. In the Health sector, the coefficients are indistinguishable from zero until 2021 and then become positive and significant in 2022 (0.0104, $p = 0.044$) and 2023 (0.0128, $p = 0.023$), and weakly so in 2024 (0.0098, $p = 0.085$). The Technology sector displays the opposite sign in the earlier part of the period, with negative and significant estimates in 2018 (−0.0067, $p = 0.047$) and 2020 (−0.0087, $p = 0.007$). In the Industrial sector, a single year attains weak significance (2020: 0.0045, $p = 0.077$), and in the Finance and Discretionary sectors no year reaches conventional significance.

The Cross-Regional Sectoral Model reintroduces regional interactions within each sector, allowing the estimated association to vary with both industry and institutional context. Each sectoral specification estimates eighteen ESG coefficients on between 803 and 1438 observations, so the power available for any single year is limited, and the year-by-year stars are not the appropriate summary. Three joint tests are therefore reported for each sector in Table A6: whether the European coefficients are jointly zero, whether the U.S. coefficients are jointly zero, and whether the regional differentials are jointly zero. Every

sector rejects at least one of the three at the 10% level, but which one it rejects differs across sectors, and that is the substantive result of this section.

Table 3. Summary statistics and output of the global sectoral analysis.

Year	Coefficient	Std. Error	95% CI	p-Value
Panel: Discretionary				
2016	−0.000154	0.003215	[−0.006495, 0.006187]	0.961833
2017	0.000132	0.003169	[−0.006118, 0.006383]	0.966786
2018	−0.000882	0.003060	[−0.006917, 0.005153]	0.773498
2019	0.001611	0.002995	[−0.004297, 0.007518]	0.591362
2020	0.000136	0.003069	[−0.005917, 0.006189]	0.964725
2021	0.002836	0.003188	[−0.003452, 0.009124]	0.374795
2022	0.001615	0.003131	[−0.004559, 0.007789]	0.606537
2023	0.002216	0.003061	[−0.003821, 0.008254]	0.469994
2024	0.002157	0.003176	[−0.004106, 0.008420]	0.497744
Clusters = 197; residual degrees of freedom = 1127.				
Panel: Finance				
2016	−0.000239	0.004404	[−0.008924, 0.008446]	0.956764
2017	0.001089	0.002547	[−0.003933, 0.006111]	0.669452
2018	−0.000227	0.002293	[−0.004749, 0.004295]	0.921270
2019	0.001595	0.002235	[−0.002812, 0.006003]	0.476216
2020	0.000187	0.002491	[−0.004725, 0.005098]	0.940373
2021	−0.003388	0.003381	[−0.010055, 0.003280]	0.317569
2022	−0.002235	0.002234	[−0.006641, 0.002171]	0.318297
2023	0.000306	0.002095	[−0.003825, 0.004438]	0.883903
2024	0.000639	0.002878	[−0.005036, 0.006314]	0.824535
Clusters = 200; residual degrees of freedom = 1216.				
Panel: Health				
2016	−0.001256	0.004238	[−0.009636, 0.007125]	0.767430
2017	0.001481	0.004484	[−0.007387, 0.010350]	0.741689
2018	0.001862	0.005783	[−0.009574, 0.013298]	0.747895
2019	0.000683	0.004386	[−0.007992, 0.009358]	0.876509
2020	0.000962	0.004471	[−0.007881, 0.009805]	0.829973
2021	0.004012	0.004874	[−0.005627, 0.013651]	0.411916
2022	0.010384 **	0.005104	[0.000290, 0.020477]	0.043860
2023	0.012799 **	0.005563	[0.001798, 0.023800]	0.022925
2024	0.009830 *	0.005658	[−0.001360, 0.021020]	0.084613
Clusters = 136; residual degrees of freedom = 645.				

Table 3. Cont.

Year	Coefficient	Std. Error	95% CI	p-Value
Panel: Industrial				
2016	0.002947	0.002672	[−0.002328, 0.008222]	0.271629
2017	0.000459	0.002261	[−0.004005, 0.004923]	0.839435
2018	0.000365	0.002400	[−0.004373, 0.005103]	0.879355
2019	0.001931	0.002384	[−0.002774, 0.006637]	0.418950
2020	0.004473 *	0.002517	[−0.000496, 0.009443]	0.077351
2021	0.001523	0.002644	[−0.003696, 0.006742]	0.565280
2022	0.000605	0.002717	[−0.004758, 0.005968]	0.823975
2023	0.001277	0.002951	[−0.004548, 0.007102]	0.665657
2024	−0.001154	0.003338	[−0.007744, 0.005435]	0.729945
Clusters = 170; residual degrees of freedom = 1030.				
Panel: Technology				
2016	−0.003099	0.003964	[−0.010923, 0.004726]	0.435441
2017	−0.004078	0.003665	[−0.011314, 0.003157]	0.267437
2018	−0.006649 **	0.003328	[−0.013219, −0.000079]	0.047327
2019	−0.003232	0.003153	[−0.009456, 0.002991]	0.306712
2020	−0.008654 ***	0.003160	[−0.014891, −0.002417]	0.006822
2021	−0.005043	0.003497	[−0.011946, 0.001860]	0.151127
2022	0.000779	0.003453	[−0.006036, 0.007595]	0.821666
2023	−0.001130	0.003371	[−0.007784, 0.005525]	0.737949
2024	−0.001874	0.004164	[−0.010094, 0.006346]	0.653228
Clusters = 171; residual degrees of freedom = 846.				

Standard errors, confidence intervals and *p*-values are taken directly from the estimated firm-clustered covariance matrix. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

In the Finance sector (Table 4), only the U.S. coefficients do so (chi-square(9) = 15.38, $p = 0.081$), and the regional differential is the weakest of the five (chi-square(9) = 5.55, $p = 0.78$). Health (Table 5) is the only sector in which all three tests reject: the European coefficients (chi-square(9) = 21.63, $p = 0.010$), the U.S. coefficients (chi-square(9) = 17.47, $p = 0.042$) and the regional differential (chi-square(9) = 16.27, $p = 0.061$). Discretionary (Table 6) presents the opposite configuration: neither regional level is jointly distinguishable from zero, yet the differential between them is the strongest in the sample (chi-square(9) = 37.47, $p < 0.001$), indicating that the two regions move differently even though neither displays a level association on its own. In Industrial, only the European coefficients approach joint significance (chi-square(9) = 16.41, $p = 0.059$) (Table 7). Lastly, in Technology (Table 8), the European coefficients are jointly significant (chi-square(9) = 17.57, $p = 0.041$) and uniformly negative, while the differential is not. Examining each sector in turn reveals the patterns underlying these joint results. In the Finance sector the panel regression uses the same 200-firm, 1438-observation sample reported in Table 3, augmented with the ESG $\times$ US interaction terms specified in Equation (5). The log of total assets is positive and highly significant, ROE is positive and significant ($p < 0.01$), and leverage is positive and significant ($p < 0.01$), while ROA and free float are statistically indistinguishable from

zero. Turning to the ESG coefficients, no year attains conventional significance in either region, in either direction, and the estimates change sign repeatedly across adjacent years. The joint tests confirm this reading: the regional differential is the weakest observed in any sector ($\chi^2(9) = 5.55$, $p = 0.78$) and is not restored by either robustness check. Finance is therefore the sector in which the data are least informative about the ESG–value association, and the earlier reading of this sector as exhibiting a clear transatlantic contrast is not supported once the estimation sample is restricted to the declared period and the declared winsorization is applied.

In the Health sector, the model is estimated on the same 136-firm, 803-observation sample reported in Table 3, augmented with the $\text{ESG} \times \text{US}$ interaction terms. The inclusion of firm fixed effects ensures that the estimates capture within-firm variation over time.

Table 4. Temporal cross-regional association between ESG score and EV in the Finance sector.

Year	EU Coef.	EU s.e.	EU 95% CI	EU p	US Coef.	US s.e.	US 95% CI	US p
2016	−0.000085	0.006286	[−0.012481, 0.012312]	0.989271	−0.000241	0.003429	[−0.007002, 0.006521]	0.944126
2017	0.001418	0.004011	[−0.006491, 0.009326]	0.724096	0.000107	0.002981	[−0.005771, 0.005985]	0.971443
2018	−0.000005	0.003764	[−0.007426, 0.007417]	0.998978	−0.001087	0.002722	[−0.006456, 0.004281]	0.690021
2019	0.001773	0.003705	[−0.005532, 0.009078]	0.632740	0.001021	0.002487	[−0.003883, 0.005926]	0.681767
2020	0.000563	0.003889	[−0.007105, 0.008232]	0.884955	−0.001162	0.002607	[−0.006303, 0.003978]	0.656156
2021	−0.003262	0.004945	[−0.013012, 0.006489]	0.510243	−0.003922	0.002823	[−0.009490, 0.001645]	0.166318
2022	−0.002201	0.003413	[−0.008930, 0.004529]	0.519776	−0.002200	0.002550	[−0.007228, 0.002828]	0.389296
2023	0.000274	0.003379	[−0.006388, 0.006936]	0.935417	0.000503	0.002345	[−0.004121, 0.005127]	0.830289
2024	0.000816	0.003783	[−0.006644, 0.008275]	0.829506	0.003774	0.003052	[−0.002245, 0.009792]	0.217726

Standard errors are clustered at the firm level (201 clusters); standard errors, confidence intervals and p -values are taken directly from the estimated clustered covariance matrix. The year 2016 is the omitted category for the year dummies, while the ESG coefficients for 2016 are estimated and are therefore reported. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 5. Temporal cross-regional association between ESG score and EV in the Health sector.

Year	EU Coef.	EU s.e.	EU 95% CI	EU p	US Coef.	US s.e.	US 95% CI	US p
2016	−0.001838	0.005803	[−0.013315, 0.009638]	0.751892	−0.005742	0.006980	[−0.019548, 0.008063]	0.412159
2017	0.000582	0.006170	[−0.011620, 0.012785]	0.924969	−0.001313	0.006431	[−0.014030, 0.011405]	0.838565
2018	0.000890	0.007440	[−0.013824, 0.015603]	0.904990	0.002034	0.007326	[−0.012455, 0.016522]	0.781758
2019	−0.000331	0.006127	[−0.012448, 0.011786]	0.956985	−0.001867	0.005664	[−0.013068, 0.009335]	0.742233
2020	0.000075	0.006109	[−0.012006, 0.012157]	0.990173	0.003285	0.005237	[−0.007071, 0.013641]	0.531503
2021	0.003089	0.006426	[−0.009619, 0.015797]	0.631506	0.006418	0.005695	[−0.004845, 0.017681]	0.261752
2022	0.009654	0.006491	[−0.003184, 0.022492]	0.139282	0.012683 **	0.006035	[0.000749, 0.024618]	0.037430
2023	0.012036 *	0.007175	[−0.002154, 0.026226]	0.095767	0.015133 **	0.006058	[0.003151, 0.027114]	0.013699
2024	0.009441	0.007301	[−0.004999, 0.023881]	0.198198	0.012114 *	0.006736	[−0.001208, 0.025436]	0.074348

Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 6. Temporal cross-regional association between ESG score and EV in the Discretionary goods sector.

Year	EU Coef.	EU s.e.	EU 95% CI	EU <i>p</i>	US Coef.	US s.e.	US 95% CI	US <i>p</i>
2016	0.000611	0.004725	[−0.008706, 0.009929]	0.897216	−0.001898	0.004265	[−0.010309, 0.006513]	0.656770
2017	0.000804	0.004815	[−0.008692, 0.010299]	0.867587	−0.000023	0.004046	[−0.008001, 0.007955]	0.995460
2018	−0.000364	0.004678	[−0.009591, 0.008862]	0.938005	−0.000301	0.003783	[−0.007761, 0.007159]	0.936613
2019	0.002450	0.004433	[−0.006291, 0.011192]	0.581021	−0.001736	0.003830	[−0.009289, 0.005818]	0.650946
2020	0.000921	0.004251	[−0.007463, 0.009305]	0.828717	−0.001184	0.004085	[−0.009241, 0.006873]	0.772250
2021	0.003953	0.004423	[−0.004769, 0.012675]	0.372495	0.000599	0.004087	[−0.007461, 0.008660]	0.883553
2022	0.001983	0.004444	[−0.006781, 0.010746]	0.655957	0.002469	0.003652	[−0.004733, 0.009670]	0.499786
2023	0.002674	0.004255	[−0.005718, 0.011066]	0.530460	0.002851	0.003753	[−0.004550, 0.010252]	0.448388
2024	0.002748	0.004295	[−0.005722, 0.011218]	0.523018	0.003337	0.004055	[−0.004660, 0.011334]	0.411531

Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.**Table 7.** Temporal cross-regional association between ESG score and EV in the Industrial sector.

Year	EU Coef.	EU s.e.	EU 95% CI	EU <i>p</i>	US Coef.	US s.e.	US 95% CI	US <i>p</i>
2016	0.004037	0.003794	[−0.003452, 0.011526]	0.288785	0.000462	0.003912	[−0.007260, 0.008185]	0.906094
2017	0.001543	0.003518	[−0.005402, 0.008488]	0.661542	−0.002958	0.003086	[−0.009050, 0.003134]	0.339155
2018	0.001294	0.003517	[−0.005650, 0.008238]	0.713370	−0.000888	0.003190	[−0.007184, 0.005409]	0.781165
2019	0.002845	0.003426	[−0.003918, 0.009607]	0.407447	−0.000415	0.002924	[−0.006187, 0.005356]	0.887256
2020	0.005427	0.003455	[−0.001394, 0.012248]	0.118139	0.002381	0.003259	[−0.004052, 0.008815]	0.465999
2021	0.002525	0.003517	[−0.004418, 0.009468]	0.473758	−0.000118	0.003186	[−0.006407, 0.006170]	0.970456
2022	0.001659	0.003659	[−0.005564, 0.008883]	0.650780	−0.000846	0.002914	[−0.006599, 0.004908]	0.772080
2023	0.002041	0.003841	[−0.005542, 0.009624]	0.595878	0.001166	0.003216	[−0.005183, 0.007515]	0.717455
2024	−0.000112	0.004257	[−0.008515, 0.008291]	0.979096	0.000965	0.003606	[−0.006155, 0.008084]	0.789429

Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.**Table 8.** Temporal cross-regional association between ESG score and EV in the Technology sector.

Year	EU Coef.	EU s.e.	EU 95% CI	EU <i>p</i>	US Coef.	US s.e.	US 95% CI	US <i>p</i>
2016	−0.006625	0.005114	[−0.016720, 0.003471]	0.196948	−0.001985	0.005158	[−0.012166, 0.008197]	0.700900
2017	−0.007550	0.005147	[−0.017710, 0.002610]	0.144242	−0.003620	0.004517	[−0.012537, 0.005296]	0.423932
2018	−0.010063 **	0.004728	[−0.019395, −0.000730]	0.034738	−0.005424	0.004586	[−0.014477, 0.003629]	0.238574
2019	−0.006616	0.004759	[−0.016009, 0.002778]	0.166261	−0.002492	0.004345	[−0.011068, 0.006085]	0.567046
2020	−0.011874 **	0.004855	[−0.021457, −0.002290]	0.015473	−0.007193 *	0.003998	[−0.015086, 0.000699]	0.073765
2021	−0.007986	0.005064	[−0.017982, 0.002010]	0.116633	−0.004425	0.004237	[−0.012789, 0.003938]	0.297751
2022	−0.002719	0.005170	[−0.012925, 0.007487]	0.599690	0.002088	0.003959	[−0.005727, 0.009904]	0.598548
2023	−0.005316	0.005266	[−0.015710, 0.005079]	0.314162	0.002001	0.003862	[−0.005623, 0.009626]	0.605021
2024	−0.006624	0.005917	[−0.018305, 0.005056]	0.264505	0.002442	0.004432	[−0.006308, 0.011191]	0.582455

Standard errors are clustered at the firm level (171 clusters); standard errors, confidence intervals and *p*-values are taken directly from the estimated clustered covariance matrix. The year 2016 is omitted from the year dummies, while the ESG coefficients for 2016 are estimated and therefore reported. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

The log of total assets is positive and highly significant; none of the remaining controls reach conventional significance in this sector. Turning to the ESG coefficients, both regions display the same temporal shape: estimates indistinguishable from zero through 2021, followed by positive estimates in the last three years. In the United States, these become significant in 2022 (0.0127, $p = 0.037$) and 2023 (0.0151, $p = 0.014$) and remain weakly significant in 2024 (0.0121, $p = 0.074$). In Europe, the same upward movement is visible but reaches conventional significance only in 2023 (0.0120, $p = 0.096$). The joint tests reject in both regions, and the differential between them is significant at the 10% level (chi-square(9) = 16.27, $p = 0.061$). Health is thus the one sector in which a positive association is visible on both sides of the Atlantic and in which the two regions nonetheless differ, the difference being one of degree and of timing rather than of direction.

For the Discretionary sector, the model is based on the same 197-firm, 1346-observation sample reported in Table 3. Among the controls, firm size is positive and highly significant, and ROA is positive and highly significant, while ROE is negative and significant ($p < 0.01$). Turning to the ESG coefficients, no year attains conventional significance in either region: the European estimates are positive in eight of the nine years, and the U.S. estimates are negative in the first half of the period and positive in the second, but none are individually distinguishable from zero. The joint tests on the two regional levels likewise fail to reject. The joint test on the differential, however, is the strongest in the sample (chi-square(9) = 37.47, $p < 0.001$) and remains significant under both robustness checks. The Discretionary sector therefore shows a statistically robust difference between the two regions in the absence of a demonstrable association in either one—a configuration that is informative about the comparison, but that does not license any statement about the level of the association in either region.

For the Industrial sector, the regression is estimated on the same 170-firm, 1222-observation sample reported in Table 3, providing an unbalanced panel over the nine-year horizon. Control variables behave consistently with expectations: firm size and ROA are positive and highly significant, ROE is negative and significant ($p < 0.05$), and leverage is positive and significant ($p < 0.05$). Turning to the ESG coefficients, the European estimates are positive in eight of the nine years, with the largest value in 2020 (0.0054, $p = 0.118$), and jointly approach significance (chi-square(9) = 16.41, $p = 0.059$); no single year, however, reaches conventional significance. The U.S. estimates fluctuate around zero and are jointly indistinguishable from it, as is the regional differential (chi-square(9) = 13.41, $p = 0.145$). The Industrial sector thus displays a consistently signed but individually weak European pattern and no transatlantic contrast.

The Technology sector includes the same 171-firm, 1039-observation sample reported in Table 3. The overall explanatory power of the model is relatively high, with a within R-squared of 0.451. Among the control variables, the log of assets and ROA are positive and highly significant, while ROE, leverage, and free float do not reach conventional significance levels. Regarding the ESG coefficients, Technology is the only sector in which the European estimates are negative in every year of the sample, ranging from -0.0027 to -0.0119 , with statistical significance in 2018 (-0.0101 , $p = 0.035$) and 2020 (-0.0119 , $p = 0.015$) and joint significance over the period as a whole (chi-square(9) = 17.57, $p = 0.041$). The U.S. estimates are negative in the first half of the sample—weakly significant in 2020 (-0.0072 , $p = 0.074$)—and turn positive but insignificant from 2022. The regional differential is not jointly significant (chi-square(9) = 10.00, $p = 0.351$). The substantive finding in this sector is therefore not a transatlantic gap but a persistent negative association between ESG scores and enterprise value among European technology firms, which runs opposite in sign to the association documented for U.S. firms in the primary model and which the present design cannot attribute to any specific mechanism.

Finally, the regional functional-form model investigates whether the ESG–value relationship is adequately described by a linear specification in each region, applying the two-criterion decision rule set out in the Methodology. The two criteria disagree, and the disagreement is reported rather than resolved by selection. The quadratic terms are jointly significant ($\chi^2(2) = 8.84$, $p = 0.012$), and, taken separately, the curvature is positive and significant in both regions (Europe: 1.83×10^{-4} , $p = 0.032$; United States: 2.16×10^{-4} , $p = 0.043$). The Bayesian Information Criterion, however, deteriorates when the quadratic terms are added ($\Delta\text{BIC} = +4.4$): with 5848 observations, the criterion imposes a penalty of about 8.7 per parameter, which the improvement in fit does not cover. The cubic terms are neither significant ($\chi^2(2) = 1.92$, $p = 0.382$) nor BIC-improving ($\Delta\text{BIC} = +13.6$). Since the decision rule stated in the Methodology requires both criteria to be satisfied, the linear specification is retained for the models reported above, and the quadratic specification is reported here as a secondary result. The imprecision of the implied turning point limits its substantive content. For Europe, the point estimate is an ESG score of 67.5, but its 95% Fieller confidence interval ranges from 18.7 to 521.7; that is, it extends well beyond the upper limit of the 0–100 scale: the data bound the turning point from below rather than from above. For the United States, the implied turning point is 32.9, and its interval covers almost the entire scale. The appropriate conclusion is that some convexity is detectable while its location is not, and the paper therefore does not advance a threshold interpretation. The limited power of within-firm identification is relevant here: the within-firm standard deviation of the ESG score is 6.83 points, compared with 22.10 points in levels, and the typical within-firm range spans approximately ± 8 points.

Robustness to Dynamic Specification and to the Measurement of Firm Value

The specification reported in Table 1 measures ESG performance contemporaneously with firm value and scales value by the asset base. Three further specifications examine whether the comparative pattern documented above depends on the timing convention, on the valuation measure, or on the control set. None are intended to identify a causal effect.

The first specification replaces the contemporaneous ESG score with its one-year lag, so that firm value in year t is related to ESG performance recorded in year $t - 1$. The estimation sample falls to 5149 firm-year observations from 850 firms, and the change is not simply the loss of each firm's first observation. Relative to the sample in Table 1, 779 observations are lost: 734 are first observations of a firm, and 45 are observations for which the preceding year is missing from the panel, so no lagged score exists even though the firm is observed earlier. A further 80 observations enter that are absent from Table 1, namely firm-years in which the contemporaneous ESG score is missing while the lagged score is available. The two samples are therefore overlapping rather than nested, which is why the observation counts in Table A4 differ across specifications. The results are reported in Table 9. The U.S. estimates for 2023 and 2024 are positive and significant at the 10% level (0.0094 , $p = 0.050$ and 0.0103 , $p = 0.063$); no European coefficient is positive and significant in any year, and the joint test on the regional differentials continues to reject ($\chi^2(8) = 20.56$, $p = 0.008$). The temporal location of the association is therefore unchanged when ESG performance is measured before firm value, although the individual estimates are less precisely determined.

The second specification retains the contemporaneous ESG score and replaces the dependent variable with the natural logarithm of Enterprise Value, the measure used in the earlier version of this analysis. It is an absolute money measure rather than a valuation ratio, and is reported here for comparability with the part of the literature that adopts it. The results are in Table 10. The within R-squared is much higher (0.328 versus 0.093), which reflects the mechanical association between an absolute measure of value and firm size

rather than the additional explanatory power of ESG. On this measure, the U.S. estimates are positive and significant at the 5% level in both 2023 (0.0044, $p = 0.013$) and 2024 (0.0044, $p = 0.021$); no European coefficient is significant in any year, and the joint test rejects more strongly than on Tobin's Q (chi-square(9) = 29.25, $p < 0.001$). The comparative result is thus confirmed on both measures of value, and is somewhat sharper on the absolute measure.

Table 9. Primary model estimated with the one-year lagged ESG score.

Year	EU Coef.	EU s.e.	EU 95% CI	EU p	US Coef.	US s.e.	US 95% CI	US p
2017	−0.005352	0.004273	[−0.013738, 0.003034]	0.210708	0.000274	0.006208	[−0.011911, 0.012458]	0.964850
2018	−0.007566 *	0.004172	[−0.015754, 0.000622]	0.070068	−0.000796	0.006075	[−0.012720, 0.011127]	0.895754
2019	−0.004348	0.004056	[−0.012310, 0.003613]	0.283987	0.000177	0.005528	[−0.010674, 0.011028]	0.974489
2020	−0.008760 **	0.004381	[−0.017359, −0.000162]	0.045858	−0.002534	0.006287	[−0.014873, 0.009805]	0.687014
2021	−0.000798	0.004189	[−0.009020, 0.007423]	0.848873	0.006832	0.005868	[−0.004685, 0.018348]	0.244639
2022	−0.004718	0.004385	[−0.013324, 0.003889]	0.282278	0.003945	0.004858	[−0.005590, 0.013481]	0.416950
2023	−0.001460	0.004239	[−0.009780, 0.006861]	0.730693	0.009420 *	0.004804	[−0.000009, 0.018849]	0.050211
2024	−0.002986	0.004647	[−0.012107, 0.006135]	0.520702	0.010303 *	0.005533	[−0.000558, 0.021163]	0.062950

Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 10. Primary model estimated with an alternative measure of firm value (log Enterprise Value).

Year	EU Coef.	EU s.e.	EU 95% CI	EU p	US Coef.	US s.e.	US 95% CI	US p
2016	−0.001026	0.002333	[−0.005605, 0.003553]	0.660137	−0.001103	0.002099	[−0.005222, 0.003016]	0.599447
2017	−0.000225	0.002108	[−0.004362, 0.003913]	0.915196	−0.000272	0.001916	[−0.004032, 0.003489]	0.887337
2018	−0.001271	0.002079	[−0.005352, 0.002810]	0.541288	0.000065	0.001841	[−0.003548, 0.003678]	0.971716
2019	0.000858	0.001965	[−0.002998, 0.004715]	0.662298	0.000186	0.001710	[−0.003170, 0.003543]	0.913254
2020	0.000187	0.001996	[−0.003731, 0.004104]	0.925535	0.000482	0.001801	[−0.003052, 0.004016]	0.789095
2021	0.000484	0.002175	[−0.003784, 0.004752]	0.823974	0.000525	0.001807	[−0.003021, 0.004071]	0.771576
2022	0.001283	0.002057	[−0.002754, 0.005319]	0.532905	0.002116	0.001714	[−0.001247, 0.005479]	0.217155
2023	0.002250	0.002116	[−0.001904, 0.006403]	0.288105	0.004364 **	0.001748	[0.000934, 0.007795]	0.012711
2024	0.001205	0.002192	[−0.003097, 0.005507]	0.582603	0.004419 **	0.001918	[0.000655, 0.008183]	0.021443

Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

A third specification removes the two profitability controls, and a fourth replaces the debt-to-equity ratio with net debt over total assets; both are reported in Tables A4 and A5 and leave the comparative result in place (chi-square(9) = 19.38, $p = 0.022$ and chi-square(9) = 24.25, $p = 0.004$ respectively). The same checks were then applied to the five cross-regional sectoral models. Reporting them in full would require ten additional tables; Table 11 therefore summarises them through the joint test on the regional differential, which is the quantity of interest in those specifications.

On the Tobin's Q measure, the regional differential is significant in only one sector, Discretionary (chi-square(9) = 19.04, $p = 0.025$), and it also survives the lagged specification ($p = 0.014$). On log Enterprise Value, the same sector rejects far more strongly ($p < 0.001$), and Health rejects at the 10% level ($p = 0.061$), while Finance, Industrial and Technology do not reject on either measure. The sectoral evidence for a transatlantic differential therefore rests on the Discretionary sector, is measure-dependent elsewhere, and is reported with that qualification in the Discussion.

Table 11. Joint test of the regional differential across specifications, by sector.

Sector	Firms	Baseline (Tobin's Q)	ESG (t−1), Tobin's Q	log(Enterprise Value)
Finance	200	9.01 (0.436)	10.84 (0.211)	5.55 (0.784)
Health	136	11.39 (0.250)	9.28 (0.320)	16.27 (0.061)
Discretionary	197	19.04 (0.025)	19.18 (0.014)	37.47 (<0.001)
Industrial	170	9.91 (0.358)	9.40 (0.310)	13.41 (0.145)
Technology	171	6.96 (0.641)	8.27 (0.408)	10.00 (0.351)

Each cell reports the chi-square statistic of the joint test that the regional differentials are zero in every year, with the *p*-value in parentheses, estimated on the cross-regional sectoral model. Standard errors are clustered at the firm level. The first two columns use Tobin's Q as the dependent variable, and the third uses log Enterprise Value. The lagged specification tests eight restrictions rather than nine, since the first year of each firm is lost. Sample sizes differ across columns for the reasons set out above.

Taken together, these checks indicate that the comparative result of the paper—a difference between the United States and the European Union in the association between ESG scores and firm value, concentrated in the closing years of the sample—is not an artefact of contemporaneous timing, of the valuation measure, or of the control set (see Table 12 for a summary of all the checks). A direct test of sectoral heterogeneity was also conducted, since the separate sector regressions compare each sector with zero rather than with one another. Estimating a single model in which the ESG term and the ESG × US interaction are allowed to differ by sector, the null that the European association is the same in all five sectors is rejected only at the 10% level (chi-square(4) = 7.84, *p* = 0.098), the null that the regional differential is the same in all five sectors is not rejected (chi-square(4) = 2.26, *p* = 0.688), and the joint null of no sectoral heterogeneity is rejected only at the 10% level (chi-square(8) = 14.00, *p* = 0.082). The sectoral patterns described above are therefore descriptive of each sector taken separately; the evidence that they differ from one another is weak, and the evidence that the transatlantic differential itself varies by sector is absent. These checks do not establish the direction of any of the reported relationships and are not offered for that purpose.

Table 12. Comparative summary of estimated ESG associations and regional differences by sector (log Enterprise Value specification, Tables 4–8).

Av. Association with EV of a +1 Point Difference in ESG Score	Financial	Healthcare	Discretionary	Industrials	Technology
EU association: mean annual coefficient (joint <i>p</i> -value)	−0.01% (0.254)	+0.37% (0.010)	+0.18% (0.249)	+0.24% (0.059)	−0.73% (0.041)
US association: mean annual coefficient (joint <i>p</i> -value)	−0.04% (0.081)	+0.47% (0.042)	+0.05% (0.496)	−0.00% (0.306)	−0.21% (0.072)
Regional differential (joint <i>p</i> -value)	0.784	0.061	<0.001	0.145	0.351
Differential significant under both robustness checks	No	Lagged only	Yes	No	No
Exposure to structural drivers (interpretive, not tested) *					
Consumer Proximity	Mid	High	High	Low	Mid
Environmental Sensitivity	Low	Mid	Mid	High	Mid
Capital Intensity	Low	Mid	Mid	High	Low

*. The three rows below the joint tests record the qualitative exposure of each sector to structural drivers, based on SASB Materiality Mapping and the LSEG materiality index. They are reported to organise the discussion and are not entered as moderating variables in any specification; no test of their association with the estimated coefficients is performed in this paper.

Collectively, these results establish four regularities. First, in the primary specification, the ESG–value association is statistically indistinguishable from zero among European firms in every year of the sample, and among U.S. firms it is positive and significant only in 2023. Second, the joint test that the regional differentials are zero in every year is rejected in the primary model and in every robustness specification examined, so the two regional profiles differ over the period as a whole, even though no single annual differential is individually significant. Third, the functional-form criteria disagree: curvature is statistically detectable but is not favoured by the information criterion, and the implied turning point cannot be located on the ESG scale. Fourth, the sectoral regressions display different patterns across industries, but a direct test of equality across sectors rejects only weakly for the European coefficients and not at all for the regional differential.

The discussion that follows considers each hypothesis in turn against the prior literature, and is concerned with candidate accounts of these patterns rather than with mechanisms established by the analysis.

5. Discussion

5.1. Hypothesis 1: The Transatlantic Difference

The evidence is consistent with H1 in direction and weak in magnitude. The two regional profiles differ over the period as a whole; the only significant annual estimate belongs to the United States, and the European estimates are indistinguishable from zero throughout. This places the paper alongside the studies that find the ESG–value relationship conditional on institutional setting rather than universal. [Garcia and Orsato \(2020\)](#) report a sign reversal between developed and emerging economies; [Ghoul et al. \(2017\)](#) find the value relevance of corporate social responsibility greater where market institutions are weaker; [Shin et al. \(2023\)](#) find national culture moderating where ESG pays. The present result extends this line of evidence by comparing two mature markets, where the institutional difference is one of regulatory approach rather than development, and finds a difference there as well. It, on the other hand, diverges from the part of the literature reporting a positive and reasonably general association, such as the aggregate reading of [Friede et al. \(2015\)](#) or the firm-level evidence of [Aydoğmuş et al. \(2022\)](#). Two features of the present design account for much of that divergence: the estimates are within-firm, so they exclude the cross-sectional component on which much of that literature rests, and they are annual, so they do not average over a period. [Candio \(2024\)](#), who finds the relationship inconsistent within Europe, and [Capelle-Blancard and Petit \(2019\)](#), who find its financial relevance varying across sectors, time and geography, are closer to the evidence reported here. The divergence is therefore less a contradiction than a difference in what is being estimated.

5.2. Interpreting the Regional Difference

([Dorothy & Endri, 2024](#)) One reading of the regional contrast is that, where sustainability disclosure is extensively mandated, ESG performance operates as a condition of legitimacy rather than as a distinguishing signal, so that variation in scores carries little incremental information for valuation; where disclosure remains discretionary, more of its signalling capacity survives. This is the mechanism anticipated in the theoretical framework and is consistent with [Dorothy and Endri \(2024\)](#) and the broader institutional literature. The present design cannot establish it. The regional indicator bundles regulatory regime with accounting conventions, index composition, ownership structure and rating coverage, and any of these would generate the same pattern; moreover, the association appears in the last years of a nine-year window, which is also when ESG rating methodologies changed most. The reading is offered as an interpretation, and the Conclusions record it as such.

5.3. Hypothesis 2: Functional Form

H2 is retained, but the evidence is more interesting than a simple retention suggests. Curvature is statistically detectable in both regions and is positive, consistent with the threshold argument: Barnett (2007) holds that returns to social responsibility accrue only once a firm has accumulated sufficient stakeholder influence capacity, and Zhang (2022) documents discontinuous responses to mandatory disclosure. What the data do not support is the quantitative claim that follows from that argument, namely that a threshold can be located. The information criterion does not favour the quadratic specification, and the implied turning point has a confidence interval extending beyond the upper limit of the ESG scale. An earlier version of this analysis reported a threshold at an ESG score of approximately 68; the corrected estimates do not sustain that claim, and it is withdrawn. The appropriate statement is that convexity may be present, but these data do not identify its location.

5.4. Hypothesis 3: Sectoral Variation

H3 is not supported in the form in which it was stated, and the reason is instructive. The sector-level regressions display different patterns: Health shows a positive association in both regions on the logarithmic scale; Technology shows a persistently negative European association; Discretionary shows the only robust regional differential; and Finance shows nothing that survives a robustness check. These patterns are broadly consistent with the sectoral literature, which associates ESG sensitivity with customer proximity (Shrestha et al., 2025; Lev et al., 2010) and with industry-level heterogeneity more generally (Xiao et al., 2025; Garcia et al., 2017). However, comparing each sector with zero is not the same as comparing sectors with one another: when the comparison is made directly, the evidence is weak for the European coefficients and absent for the regional differential. The sectoral literature cited above generally reports within-sector estimates without testing cross-sector equality, and the present result suggests that this distinction matters: patterns that look different may not be distinguishable. The structural drivers invoked in that literature are in any case not measured here and are not entered as moderators, so their role remains an interpretive device rather than a finding.

Taken as a whole, the analysis supports a narrow and specific claim: over 2016–2024 the association between ESG performance and firm value differed between the United States and the European Union, with the only statistically significant estimate appearing in the United States in the closing years of the period. It does not support the stronger claims that are sometimes made in this literature—that the difference can be attributed to regulation or to market sentiment, that a threshold of ESG performance can be identified above which value accrues, or that sectors differ from one another in a way the data can distinguish. Setting out what the evidence does not support is as much a contribution as setting out what it does, in a field where measurement choices and specification choices are known to drive results.

6. Conclusions

This study has examined whether the association between ESG performance and firm value differs across institutional contexts, comparing firms headquartered in the United States and the European Union over 2016–2024, using data on 874 firms and 5848 firm-year observations, with firm value measured by Tobin's Q in levels.

H1 is supported. The null that the regional differentials are zero in every year is rejected ($\chi^2(9) = 19.60, p = 0.021$) and remains rejected with the ESG score lagged, with firm value measured by the logarithm of Enterprise Value, with the profitability controls removed and with leverage redefined as net debt over assets. The direction is as

predicted: the only significant annual estimate is positive and belongs to the United States (2023, $p = 0.023$), while no European estimate is significant in any year. The support is for a difference over the period as a whole; the test does not identify which years differ.

H2 is retained. The quadratic terms are jointly significant ($p = 0.012$), but the information criterion deteriorates, and the rule fixed in advance requires both, so linearity stands. The convexity that the quadratic specification detects cannot be located: the Fieller interval around the European turning point extends beyond the ESG scale. The threshold reading proposed in part of the literature therefore receives no usable support.

H3 is not supported as stated. Sector regressions display different patterns, but a direct test of equality across sectors rejects only at the 10% level for the European coefficients ($p = 0.098$) and not at all for the regional differential ($p = 0.688$)—and the latter was the part of the hypothesis concerning the interaction between sector and institutional context.

Three limitations bound these findings. The regional indicator is a compound treatment, bundling regulatory regime and investor sentiment with accounting conventions, index composition, ownership concentration, liquidity and ESG rating coverage, which the design cannot separate; the institutional reading offered in the Discussion is therefore an interpretation. The two regional samples occupy different parts of the ESG distribution (means of 55.5 and 41.8), so where the relation departs from linearity, part of the estimated difference may reflect their positions within that distribution. Moreover, identification is narrow: firm and year fixed effects absorb 96.4 percent of the variance of log Enterprise Value, so a modest time-varying factor affecting both disclosure and valuation would suffice to generate associations of the size reported. The estimates are associations throughout: no causal claim is made, and because ESG scores show no association with realised profitability one, two or three years ahead, the process linking sustainability performance to valuation is not documented here.

Four extensions follow directly. First, replacing the regional indicator with direct measures of regulatory stringency and of market sentiment would allow the components of the institutional proxy to be separated and tested individually. Second, disaggregating the European Union by member state, or by an index of institutional quality, would establish whether the European result conceals variation across national regimes. Third, repeating the analysis using ratings from a second provider would show how far the findings depend on a single rating methodology, given the documented divergence between providers. Fourth, a longer horizon than the nine years available here would permit the expectations channel to be examined, which is what the present data cannot observe. For practice, the results caution against making general claims that markets reward sustainability performance: over the period examined, the association is evident in one region, only in the final years, and unevenly across sectors.

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code, the documentation of the sample-construction filters and the full output of every specification reported in the paper are available from the corresponding author on reasonable request.

Conflicts of Interest: The authors declare no conflicts of interest.

Appendix A. Descriptive Statistics and Sample Construction

Table A1. Descriptive statistics, estimation sample 2016–2024.

Panel A. Full Sample								
Variable	N	Mean	Std. Dev.	Min	P25	Median	P75	Max
Tobin's Q	5848	1.670	1.984	0.054	0.593	1.029	1.987	14.181
log(Enterprise Value)	5848	1.374	2.139	−3.828	−0.151	1.161	2.899	6.133
ESG score	5848	48.595	22.104	8.780	29.995	46.489	67.142	91.310
log(Total Assets)	5848	1.371	2.181	−3.935	−0.122	1.182	2.664	7.283
Debt-to-equity ratio	5848	0.949	1.479	0.000	0.166	0.513	1.113	11.663
Return on assets (%)	5848	2.618	13.048	−72.785	0.771	3.658	8.285	30.848
Return on equity (%)	5848	7.452	31.086	−207.753	3.700	11.018	19.509	71.149
Free float (%)	5848	82.376	21.217	14.911	71.854	92.103	98.548	100.000
Panel B. By Region								
Variable	N (EU)	Mean (EU)	SD (EU)	N (US)	Mean (US)	SD (US)	Diff. (US-EU)	
Tobin's Q	2903	1.569	1.883	2945	1.769	2.074	+0.200	
log(Enterprise Value)	2903	1.536	2.106	2945	1.214	2.160	−0.322	
ESG score	2903	55.467	22.049	2945	41.820	19.967	−13.647	
log(Total Assets)	2903	1.555	2.321	2945	1.189	2.018	−0.366	
Debt-to-equity ratio	2903	0.945	1.386	2945	0.953	1.566	+0.008	
Return on assets (%)	2903	4.375	9.984	2945	0.885	15.291	−3.490	
Return on equity (%)	2903	10.874	21.886	2945	4.079	37.738	−6.795	
Free float (%)	2903	76.489	23.671	2945	88.179	16.548	+11.690	

Notes. Panel A and Panel B are computed on the estimation sample of the primary model reported in Table 1: 5848 firm-year observations on 874 firms between 2016 and 2024, obtained through the filters documented step by step in Table A3. All continuous variables are winsorized at the 1st and 99th percentiles before estimation. Monetary quantities are denominated in United States dollars as extracted from LSEG Workspace. ROA, ROE and free float are reported here in percentage points; they enter the regressions divided by 100, that is as ratios, and the coefficients in Table A5 are expressed on that scale.

Table A2. Correlation matrix.

Panel A. Levels								
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	
(1) log(Enterprise Value)	1.000							
(2) ESG score	0.671	1.000						
(3) log(Total Assets)	0.874	0.660	1.000					
(4) Debt-to-equity ratio	0.200	0.128	0.242	1.000				

Table A2. *Cont.*

(5) Return on assets (%)	0.312	0.205	0.231	−0.093	1.000		
(6) Return on equity (%)	0.323	0.241	0.282	−0.113	0.861	1.000	
(7) Free float (%)	0.167	0.117	0.180	−0.052	0.015	0.061	1.000
Panel B. Within Firms, After Firm and Year Fixed Effects							
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
(1) log(Enterprise Value)	1.000						
(2) ESG score	0.169	1.000					
(3) log(Total Assets)	0.520	0.351	1.000				
(4) Debt-to-equity ratio	0.053	0.025	0.063	1.000			
(5) Return on assets (%)	0.209	−0.007	0.071	−0.186	1.000		
(6) Return on equity (%)	0.167	−0.003	0.077	−0.317	0.775	1.000	
(7) Free float (%)	0.124	0.138	0.166	−0.009	0.046	0.056	1.000

Notes. Pearson correlations on the 5848 firm-year observations of the estimation sample. Panel A reports correlations in levels; Panel B reports them after the two-way within transformation, that is on the variation that identifies the coefficients once firm and year fixed effects are absorbed. The distinction is material: firm and year fixed effects absorb 96.4 percent of the variance of log Enterprise Value, leaving the 3.6 percent on which the within-firm correlations are computed, and the standard deviation of the ESG score falls from 22.10 points in levels to 6.83 points within firms.

Table A3. Sample construction.

Step	Filter Applied	Firm-Year Obs.	Firms
1	Firm-year records in the source database (LSEG Workspace)	9672	965
2	Less records with a missing observation year	9670	965
3	Restricted to the 2016–2024 analysis window	8703	965
4	Less 18 duplicated firm-year records (two firms)	8685	965
5	Enterprise Value strictly positive	7755	962
6	Total Assets strictly positive	7746	962
7	ESG score non-missing	6792	961
8	Control variables non-missing (listwise deletion)	5881	907
9	Firms with at least two observations (estimation sample)	5848	874

Notes. Every step in this table is reproducible from the source database. Step 3 restricts the records to the 2016–2024 window over which ESG coverage is reliable; the earlier version of this analysis did not apply this restriction. Step 4 removes two firms recorded twice in the source extract, one of which carried two different sector labels and has been assigned to Finance. Step 9 is applied within each specification, so the estimation sample differs slightly across models; the figures shown are those of the primary model reported in Table 1 and coincide with the first row of Table A4.

Table A4. Model statistics for all reported specifications.

Specification	Obs.	Clusters	Param.	Residual d.f.	Within R ²	Robust F	Wald $\chi^2(\gamma)$	p
Table 1—Primary model, Tobin's Q (ESG score)	5848	874	31	4943	0.0928	6.13	19.60	0.021
Table 2—Primary model, Tobin's Q (ESG combined score)	5847	874	31	4942	0.0929	6.02	18.36	0.031

Table A4. Cont.

Specification	Obs.	Clusters	Param.	Residual d.f.	Within R ²	Robust F	Wald $\chi^2(\gamma)$	p
Table 3—Global sectoral model (Discretionary)	1346	197	22	1127	0.4554	14.47	—	—
Table 3—Global sectoral model (Finance)	1438	200	22	1216	0.3892	23.60	—	—
Table 3—Global sectoral model (Health)	803	136	22	645	0.3834	13.69	—	—
Table 3—Global sectoral model (Industrial)	1222	170	22	1030	0.3670	20.46	—	—
Table 3—Global sectoral model (Technology)	1039	171	22	846	0.4385	17.58	—	—
Table 4—Cross-regional sectoral model (Finance)	1438	200	31	1207	0.3937	18.96	5.55	0.784
Table 5—Cross-regional sectoral model (Health)	803	136	31	636	0.3956	11.40	16.27	0.061
Table 6—Cross-regional sectoral model (Discretionary)	1346	197	31	1118	0.4687	12.70	37.47	<0.001
Table 7—Cross-regional sectoral model (Industrial)	1222	170	31	1021	0.3750	16.53	13.41	0.145
Table 8—Cross-regional sectoral model (Technology)	1039	171	31	837	0.4509	14.28	10.00	0.351
Table 9—Lagged ESG score	5149	850	28	4271	0.1134	5.30	20.56	0.008
Table 10—log(Enterprise Value)	5848	874	31	4943	0.3281	34.97	29.25	<0.001
Robustness—controls excluding ROA and ROE	6120	893	29	5198	0.0729	6.61	19.38	0.022
Robustness—leverage as net debt over assets	6289	925	31	5333	0.0891	6.26	24.25	0.004
Functional form—quadratic specification	5848	874	17	4957	0.0874	7.32	—	—
Functional form—cubic specification	5848	874	19	4955	0.0880	6.67	—	—

Notes. Within R-squared is computed on the firm-demeaned data. The robust F-statistic tests the joint significance of all estimated slope coefficients using the clustered covariance matrix. The Wald statistic tests the joint hypothesis that the regional differentials are zero in every year, and is not defined for the global sectoral models of Table 3 or for the functional-form specifications, which contain no year-specific regional interaction. The lagged specification tests eight restrictions rather than nine. Parameters counts the estimated slope coefficients, excluding the absorbed firm fixed effects; residual degrees of freedom are observations less parameters less firm fixed effects. Standard errors are clustered at the firm level throughout.

Table A5. Estimated coefficients on the control variables.

Specification	log(Assets)	ROA	ROE	Leverage	Free Float
Table 1 (primary, Q)	−0.5867 *** (0.1106)	+2.9573 *** (0.6961)	−0.3567 (0.2575)	−0.0205 (0.0274)	+0.4423 (0.4662)
Table 2 (combined score, Q)	−0.5910 *** (0.1103)	+2.9648 *** (0.6973)	−0.3558 (0.2582)	−0.0205 (0.0275)	+0.4599 (0.4696)

Table A5. Cont.

Specification	log(Assets)	ROA	ROE	Leverage	Free Float
Table 3 (Discretionary)	+0.7869 *** (0.0830)	+3.7762 *** (0.5175)	−0.4262 *** (0.1393)	−0.0061 (0.0253)	+0.4769 (0.4527)
Table 3 (Finance)	+0.6831 *** (0.0632)	−0.4179 (0.3700)	+0.8041 *** (0.2885)	+0.1635 *** (0.0432)	+0.4223 (0.2693)
Table 3 (Health)	+0.9550 *** (0.0980)	−0.1416 (0.5052)	−0.0427 (0.2378)	−0.0431 (0.0307)	+0.6250 (0.4763)
Table 3 (Industrial)	+0.7635 *** (0.0709)	+3.0362 *** (0.6229)	−0.3390 ** (0.1497)	+0.0267 * (0.0144)	−0.0054 (0.2764)
Table 3 (Technology)	+0.7594 *** (0.0925)	+1.9618 *** (0.5338)	−0.0253 (0.2025)	+0.0101 (0.0202)	+0.0410 (0.4965)
Table 4 (Finance)	+0.6612 *** (0.0707)	−0.4678 (0.3793)	+0.8493 *** (0.2992)	+0.1685 *** (0.0431)	+0.3900 (0.2672)
Table 5 (Health)	+0.9696 *** (0.0982)	−0.1412 (0.5066)	−0.0618 (0.2399)	−0.0480 (0.0297)	+0.5844 (0.4596)
Table 6 (Discretionary)	+0.7916 *** (0.0832)	+3.7813 *** (0.5050)	−0.4277 *** (0.1373)	−0.0087 (0.0251)	+0.4928 (0.4403)
Table 7 (Industrial)	+0.7560 *** (0.0695)	+2.9775 *** (0.6268)	−0.3144 ** (0.1521)	+0.0294 ** (0.0142)	−0.0246 (0.2749)
Table 8 (Technology)	+0.7421 *** (0.0936)	+1.9717 *** (0.5586)	−0.0566 (0.2074)	+0.0072 (0.0207)	−0.0114 (0.5068)
Table 9 (lagged ESG, Q)	−0.7273 *** (0.1455)	+3.0253 *** (0.7857)	−0.0951 (0.2921)	−0.0125 (0.0299)	+0.8721 (0.6826)
Table 10 (log Enterprise Value)	+0.7552 *** (0.0360)	+1.0165 *** (0.2520)	+0.0319 (0.0984)	+0.0307 * (0.0182)	+0.2726 (0.1893)
Robustness (no ROA/ROE)	−0.5411 *** (0.1102)	not included	not included	−0.0312 * (0.0188)	+0.7702 (0.4883)
Robustness (leverage = net debt/assets)	−0.5456 *** (0.1040)	+2.5472 *** (0.6553)	−0.2448 (0.1798)	−0.3381 (0.3097) [net debt/assets]	+0.4154 (0.5101)

Notes. Each cell reports the estimated coefficient with its firm-clustered standard error in parentheses. Assets are in natural logarithms; ROA and ROE enter as ratios, that is percentage points divided by 100, and are reported in percentage points in Table A1; free float is expressed as a share. In the specifications whose dependent variable is Tobin's Q in levels (Tables 1, 2 and 9), the coefficient on log(Assets) is negative, which is the ordinary relationship between size and a valuation ratio; in the log Enterprise Value specification of Table 10 it is positive. Significance: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A6. Joint tests by sector; cross-regional sectoral model.

Sector	Firms	Europe: H0 All Beta = 0	United States: H0 All Beta + Gamma = 0	Differential: H0 All Gamma = 0
Finance	200	11.33 (0.254)	15.38 (0.081)	5.55 (0.784)
Health	136	21.63 (0.010)	17.47 (0.042)	16.27 (0.061)
Discretionary	197	11.40 (0.249)	8.39 (0.496)	37.47 (0.000)

Table A6. *Cont.*

Sector	Firms	Europe: H0 All Beta = 0	United States: H0 All Beta + Gamma = 0	Differential: H0 All Gamma = 0
Industrial	170	16.41 (0.059)	10.58 (0.306)	13.41 (0.145)
Technology	171	17.57 (0.041)	15.75 (0.072)	10.00 (0.351)

Notes. Each cell reports the chi-square statistic of the joint test on the nine annual coefficients, with the *p*-value in parentheses; all tests have nine degrees of freedom and use the firm-clustered covariance matrix. The first column tests whether the European coefficients are jointly zero, the second whether the U.S. coefficients are jointly zero, and the third whether the regional differentials are jointly zero.

Table A7. Joint tests by sector; cross-regional sectoral model; Tobin's Q.

Sector	Firms	Europe: H0 All Beta = 0	United States: H0 All Beta + Gamma = 0	Differential: H0 All Gamma = 0
Finance	200	7.35 (0.601)	8.01 (0.533)	9.01 (0.436)
Health	136	13.31 (0.149)	15.01 (0.091)	11.39 (0.250)
Discretionary	197	8.88 (0.448)	11.08 (0.270)	19.04 (0.025)
Industrial	170	8.49 (0.485)	6.75 (0.663)	9.91 (0.358)
Technology	171	13.66 (0.135)	13.28 (0.150)	6.96 (0.641)

Notes. As in Table A6, with Tobin's Q in levels as the dependent variable. The lower power of the ratio specification at sector level is the reason the sectoral analysis is reported on log Enterprise Value in the main text.

Table A8. From hypotheses to tests.

Hypothesis	Statement	Specification	Test Statistic	Decision Rule
H1	The ESG–value association differs between U.S. and EU firms and is stronger in the United States	Equation (1): $ESG \times year$ and $ESG \times year \times US$ interactions, firm and year fixed effects	Wald test of the joint null that all nine regional differentials are zero	Rejection at the 5% level supports H1; the test does not identify which individual years differ
H2	The association is adequately described by a linear specification in each region	Equations (9) and (10): quadratic and cubic terms in ESG, by region	Wald test on the higher-order terms and change in the Bayesian Information Criterion	Linear form retained unless the Wald test rejects at 5% and the BIC improves; both criteria required
H3	The association varies across sectors and the regional difference is itself sector-dependent	Equations (4) and (5) by sector, and a pooled model with sector-specific ESG and $ESG \times US$ terms	Joint tests by sector and a direct Wald test of equality of the ESG terms across sectors	Rejection of equality across sectors at the 5% level supports H3; sector-by-sector significance alone does not

Notes. The decision rules were fixed before estimation and are applied unchanged to the robustness specifications. Equation numbers refer to Section 3.

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