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Article

The Impact of New Financial Instruments Standard on Asset Impairment Provisioning Behavior of Commercial Banks

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Abstract: This paper examines whether the mandatory adoption of China's revised Accounting Standards for Business Enterprises No. 22 —the domestic counterpart to International Financial Reporting Standard 9 (IFRS 9)—altered the asset impairment provisioning behavior of A-share listed commercial banks. Utilizing an unbalanced panel dataset comprising 35 banks over the period from 2015 to 2025, which yields a total of 385 bank-year observations, we estimate comprehensive fixed-effects regression models. In these models, the core explanatory variable is a post-adoption binary indicator designed to capture the regulatory shift. Our empirical results demonstrate that the regulatory transition itself did not independently alter provisioning intensity once bank-level control variables are adequately included in the specification. Instead, the dominant driver of loan loss provisions remains pre-provision profitability. This finding is highly consistent with systematic income smoothing practices, facilitated by the discretionary space inherently afforded by the expected credit loss model introduced under the new standard. These robust findings survive the introduction of two-way fixed effects, addressing potential endogeneity concerns. Ultimately, the study suggests that while CAS 22 expanded managerial discretion in financial reporting, it did not displace the underlying economic incentives that govern how that discretion is practically exercised by bank management. These insights provide valuable implications for regulators and policymakers aiming to enhance the transparency and stability of the financial sector.

Keywords: accounting standards; asset impairment; earnings management; commercial banks; expected credit loss

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1. Introduction

The global financial crisis of 2007–2009 exposed a fundamental weakness in prevailing bank accounting practices. Under the incurred loss model embedded in IAS 39, banks were permitted to recognize credit losses only when objective evidence of impairment existed, producing a pattern of provisioning that was systematically delayed and procyclical. Regulators and standard-setters widely characterized this as a “too little, too late” problem. In response, the International Accounting Standards Board issued IFRS 9 in 2014, replacing the incurred loss model with a forward-looking expected credit loss framework that requires banks to provision against anticipated losses from the moment of loan origination [1].

The transition from the incurred loss model to the expected credit loss framework represents one of the most consequential shifts in bank financial reporting in recent decades [1]. Early cross-country evidence indicates that expected credit loss provisions are more predictive of future bank risk than their incurred loss model counterparts, and that the informational content of the new model is especially pronounced during periods of credit deterioration. Whether this improvement in information quality translates into

meaningful changes in bank provisioning behavior, however, remains an open question. While expected credit loss-based provisions anticipate banking risk more effectively than incurred loss model-based ones, the disciplining effect of the new standard weakens considerably when regulatory authorities grant temporary relief from provisioning obligations, as occurred during certain periods of economic stress. This asymmetry suggests that the behavioral consequences of the standard depend as much on the regulatory environment surrounding its implementation as on the accounting rules themselves.

China's experience offers a particularly instructive setting in which to examine these dynamics. In 2017, revised Accounting Standards for Business Enterprises No. 22 were issued, mandating that A-share listed commercial banks adopt the expected credit loss model beginning January 1, 2019. The transition introduced a structurally new provisioning architecture, but it also expanded the scope of managerial discretion: expected credit loss estimation requires banks to form forward-looking judgments about default probabilities, loss given default, and the staging of credit exposures, all of which are inherently subjective [2, 3]. Research has documented that this discretion interacts with banks' capital positions in ways that complicate the expected risk-reduction benefits of the standard, as capital-constrained institutions face incentives to understate provisions to preserve regulatory capital ratios.

Implementation quality has been further complicated by institutional heterogeneity among audit firms [4, 5]. Evidence shows that international audit firms focus on fair value measurement and expected credit loss application in their review of compliance, while domestic large firms primarily attend to asset classification. The implication is that the quality of expected credit loss implementation across Chinese banks is uneven, and that the behavioral effects of the standard cannot be read off from its formal requirements alone. At the micro level, analysis using credit card portfolio data from a major Chinese bank demonstrates that adoption significantly increased both the volume of assets subject to impairment and the absolute amount of provisions, while simultaneously reducing the coverage ratio — a pattern consistent with banks redistributing rather than genuinely increasing their loss recognition.

Against this backdrop, this paper examines whether the mandatory adoption of Accounting Standards for Business Enterprises No. 22 altered the asset impairment provisioning behavior of Chinese A-share listed commercial banks. Using an unbalanced panel of 35 banks over the period 2015–2025, we estimate fixed-effects models in which the key explanatory variable is a post-adoption indicator (POST), supplemented by bank-level controls capturing size, loan portfolio composition, profitability, pre-provision earnings, and leverage. Our results show that POST carries a statistically significant negative coefficient in a univariate specification, but loses significance once controls are included. The dominant explanatory variable throughout is pre-provision profitability (EBLLP), which exhibits a large positive coefficient, consistent with banks using the discretion afforded by the expected credit loss model to smooth earnings rather than to recognize credit risk on a fundamentally more timely basis. The remainder of the paper proceeds as follows. Section 2 reviews the relevant literature [6, 7]. Section 3 describes the data, variables, and empirical models. Section 4 presents the regression results and robustness checks. Section 5 discusses the findings, and Section 6 concludes.

2. Literature Review

The replacement of the incurred loss model with an expected credit loss framework has generated a substantial body of empirical research, organized broadly around three questions: whether the new standard reduces procyclicality, whether it alters financial institutions' discretionary provisioning behavior, and how these dynamics manifest in the domestic institutional context [8].

On the procyclicality question, the expected credit loss model was designed to compel earlier recognition of credit deterioration, thereby smoothing the amplification effect that delayed provisioning had on credit cycles under the prior standard. Empirical

analysis using panel data from a large sample of regional commercial banks found that adoption of the new standard did mitigate procyclicality in impairment provisioning, though the countercyclical benefit was significantly attenuated in jurisdictions experiencing leadership transitions at the local administrative level. This finding reveals that the effectiveness of an accounting standard depends as much on the governance environment surrounding its implementation as on the technical design of the rules themselves [4]. Further evidence from state-owned financial institutions indicates that institutions supervised by local administrative officials nearing the end of their appointed terms recognize smaller and less timely impairment provisions than comparable institutions. Incentives tied to administrative tenure, in other words, can directly moderate the forward-looking logic that the expected credit loss model was built to institutionalize. At the implementation level, studies report that over 60 percent of sampled domestic institutions exhibit inconsistent reclassifications of equity instruments under the updated accounting standard, and that post-adoption improvements in the value relevance of impairment charges are largely confined to entities audited by internationally recognized audit firms. Heterogeneous implementation quality therefore places an important practical ceiling on what the standard can deliver, regardless of its formal requirements.

The question of whether expected credit loss adoption changes financial institutions' discretionary behavior has attracted considerable attention, with findings that are less sanguine than the standard's architects anticipated [9]. Analyses of large samples of financial institutions in multiple jurisdictions during the initial years of implementation find that listed institutions and those holding assets to maturity use impairment provisions more intensively for earnings management after adoption than before, with the tendency intensifying during periods of systemic stress. The mechanism is that expected credit loss estimation requires management to exercise judgment over staging decisions, probability-of-default modeling, and forward-looking macroeconomic adjustments, which expands rather than constrains the scope for judgment-based decision making. Additional evidence confirms that discretionary provisions are deployed simultaneously for income smoothing, capital management, and signaling purposes, with the intensity of each objective increasing as crisis severity deepens. A structural analysis further shows that ownership structure influences both the magnitude of discretionary provisions and the propensity to smooth income, with institutions under foreign ownership exhibiting the most pronounced smoothing behavior even after controlling for asset quality. These studies collectively suggest that the expected credit loss model has shifted the locus of judgment-based provisioning from discrete loss events to the continuous calibration of expected loss models, without eliminating the underlying incentives that shape such judgments.

Financial condition introduces further heterogeneity into provisioning behavior. Evidence indicates that institutions experiencing material financial stress engage in judgment-based provisioning to a significantly greater degree than their stable peers, and that the direction of adjustment varies with proximity to distress, ranging from upward adjustment to conserve regulatory capital to downward adjustment to project stability. This implies that behavioral responses to provisioning standards are distributed unevenly across the risk spectrum, a nuance that aggregate panel estimates may obscure [10]. Methodological work has constructed an empirical decomposition of impairment provisions into credit-risk-driven and judgment-based components, establishing that the non-judgment-based portion can be modeled from observable portfolio characteristics. This framework clarifies that much of the variation in provisioning levels documented across the literature reflects managerial judgment rather than underlying credit risk, which is directly relevant to interpreting the results of the present study.

Regulatory attempts to guide judgment-based provisioning have met with mixed success. Studies examining financial institutions over extended time horizons find that sequences of supervisory measures beginning in the early 2010s partially corrected entrenched income-smoothing behavior. The correction was statistically significant but

incomplete, suggesting that accounting rule changes are insufficient on their own to alter provisioning incentives without sustained supervisory reinforcement [11]. A longer-term domestic perspective examining provisioning behavior around major regulatory transitions finds persistent evidence of income smoothing through impairment provisions that survived successive regulatory reforms. The durability of these behavioral patterns across successive regulatory changes implies that the incentives driving judgment-based provisioning in the domestic banking sector are deeply embedded in its institutional structure, rather than being transient responses to any particular standard.

The literature reviewed above converges on a conclusion that directly motivates the present study: accounting standard transitions alter the formal rules governing loss recognition but do not automatically displace the behavioral incentives that drive provisioning decisions [8]. Whether adoption of the domestic expected credit loss standard produced a genuine shift in the provisioning behavior of A-share listed commercial banks, or whether pre-existing objectives continued to dominate under the expanded scope for judgment, remains an open empirical question—not yet resolved with a full panel spanning both sides of the transition.

3. Data & Methodology

3.1. Sample, Data & Variables

This study draws on annual financial statement data for A-share listed commercial banks in China over the period 2015 to 2025. The sample comprises 35 banks, yielding 385 bank-year observations with no missing values [12]. Data are collected from publicly available annual reports accessed via Sina Finance. The sample period is chosen to span both sides of the CAS 22 adoption timeline: the pre-adoption period covers 2015–2018, during which banks applied the incurred loss model, and the post-adoption period covers 2019–2025, during which A-share listed banks were required to apply the expected credit loss model (As shown in Table 1).

Table 1. Descriptive Statistics

Variable	Count	Mean	Std	Min	Max
LLP	385	0.0125	0.0052	0.0030	0.0254
POST	385	0.7273	0.4459	0.0000	1.0000
SIZE	385	28.1108	1.7367	25.2213	31.4342
LOAN_RATIO	385	0.5098	0.0806	0.3115	0.6728
ROA	385	0.0080	0.0017	0.0041	0.0125
EBLLP	385	0.0142	0.0029	0.0076	0.0220
LEV	385	0.9241	0.0103	0.9027	0.9502

The dependent variable is loan loss provisions scaled by total loans (LLP), which measures the intensity of asset impairment provisioning in a given bank-year [13]. This scaling convention is standard in the provisioning literature and ensures comparability across banks of different sizes.

The core explanatory variable is POST, a binary indicator that equals one for observations from 2019 onward and zero otherwise, capturing the shift following the mandatory adoption of CAS 22 by domestically listed banks. A negative and statistically significant coefficient on POST would indicate that the new standard reduced provisioning intensity relative to the pre-adoption period, while a positive coefficient would suggest the opposite. The absence of significance, as the results will show, implies that the adoption of CAS 22 did not independently alter provisioning behavior once other bank characteristics are accounted for.

Six bank-level control variables are included. SIZE is the natural logarithm of total assets, included to capture the possibility that larger banks face different provisioning incentives and regulatory scrutiny than smaller institutions. CREDIT_RATIO is the ratio of total credit exposures to total assets, reflecting the credit intensity of a bank's business model; banks with higher credit ratios are expected to carry proportionally larger credit

risk exposures. ROA is net income scaled by total assets, capturing overall profitability; more profitable banks may have less incentive to use provisions for earnings management purposes, or alternatively, may provision more to smooth unusually high earnings downward. EBLLP is defined as the sum of net income and loan loss provisions scaled by total assets, representing pre-provision profitability; a positive coefficient on this variable is the standard empirical signature of income smoothing, indicating that banks provision more when pre-provision earnings are high and less when they are low. LEV is total liabilities scaled by total assets, measuring financial leverage; more highly leveraged banks operate with thinner capital buffers and may face stronger incentives to manage provisions in ways that preserve reported capital adequacy.

Descriptive statistics for all variables are presented in Table 1. The mean LLP of 0.0125 with a standard deviation of 0.0052 indicates moderate cross-sectional variation in provisioning intensity across the sample. The POST variable has a mean of 0.7273, reflecting the fact that approximately 73 percent of observations fall in the post-adoption period, consistent with the 2019 adoption date applied across an 11-year panel. The relatively low standard deviation of ROA (0.0017) signals limited variation in profitability across Chinese listed banks, while EBLLP exhibits a mean of 0.0142 and a standard deviation of 0.0029, slightly higher than LLP alone due to the add-back of provisions. The correlation matrix in Table 2 provides preliminary evidence on variable relationships. The correlation between EBLLP and LLP is 0.712, the largest pairwise correlation in the matrix, which anticipates the dominant role of pre-provision profitability in the regression results [14]. CREDIT_RATIO and LEV exhibit a correlation of -0.641, reflecting the structural relationship between credit intensity and leverage in bank balance sheets; both variables are retained in the model as they capture conceptually distinct dimensions of bank risk.

Table 2. Correlation Matrix

	LLP	POST	SIZE	LOAN_RATIO	ROA	EBLLP	LEV
LLP	1.000	-0.194	-0.256	-0.344	-0.131	0.712	0.165
POST	-0.194	1.000	0.144	0.477	-0.193	-0.100	-0.387
SIZE	-0.256	0.144	1.000	0.201	0.091	-0.090	-0.043
LOAN_RATIO	-0.344	0.477	0.201	1.000	0.100	0.124	-0.641
ROA	-0.131	-0.193	0.091	0.100	1.000	0.482	-0.235
EBLLP	0.712	-0.100	-0.090	0.124	0.482	1.000	-0.235
LEV	0.165	-0.387	-0.043	-0.641	-0.233	-0.235	1.000

3.2. Model Specification

The baseline empirical model takes the following form:

$$LLP_{it} = \alpha + \beta_1 POST_t + \beta_2 SIZE_{it} + \beta_3 LOAN_RATIO_{it} + \beta_4 ROA_{it} + \beta_5 EBLLP_{it} + \beta_6 LEV_{it} + \mu_i + \varepsilon_{it}$$

where LLP_{it} is the credit loss provision ratio for bank i in year t , μ_i denotes bank fixed effects, and ε_{it} is the idiosyncratic error term. Bank fixed effects are included to absorb time-invariant unobserved heterogeneity across institutions, such as differences in business strategy, ownership structure, and geographic focus that remain stable over the sample period but may influence provisioning behavior. Standard errors are clustered at the bank level to account for serial correlation within banks over time [8].

Three specifications are estimated. Model 1 includes a policy implementation indicator as the sole explanatory variable alongside bank fixed effects, establishing the unconditional association between the regulatory change and provisioning intensity [15]. Model 2 adds the full set of bank-level controls, allowing the coefficient on the policy implementation indicator to be interpreted conditional on observable bank characteristics. Model 3 replaces the policy implementation indicator with a full set of year fixed effects, implementing a two-way fixed effects specification that absorbs all common time-varying shocks — including the regulatory change itself — and serves as the primary robustness check. Because the policy implementation indicator is a time-varying variable taking the same value for all banks in a given year, it is collinear with year fixed effects in Model 3

and is therefore excluded; the year fixed effects subsume its variation. Comparing the stability of the control variable coefficients across Models 2 and 3 assesses whether the baseline results are sensitive to the treatment of aggregate time trends.

4. Empirical Results

4.1. Baseline Regression Results

Table 3 reports the results of the three model specifications described in Section 3.2. The dependent variable throughout is LLP, the ratio of loan loss provisions to total loans.

Table 3. Regression Results

Variable	Model (1)	Model (2)	Model (3)
POST	−0.0023*** (0.0007)	0.0003 (0.0002)	—
SIZE	—	−0.0005 (0.0004)	−0.0021** (0.0008)
LOAN_RATIO	—	−0.0262*** (0.0033)	−0.0283*** (0.0042)
ROA	—	−1.9273*** (0.0774)	−1.8670*** (0.0834)
EBLLP	—	1.8337*** (0.0599)	1.9213*** (0.0826)
LEV	—	0.0219** (0.0098)	0.0334*** (0.0118)
Constant	0.0141*** (0.0005)	0.0080 (0.0142)	0.0421** (0.0206)
Entity FE	Yes	Yes	Yes
Time FE	No	No	Yes
Obs	385	385	385
R-squared	0.0603	0.9638	0.9281

Note: * $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$. Clustered standard errors by bank in parentheses. POST is omitted from Model 3 due to collinearity with year fixed effects.

Model 1 includes POST as the sole regressor alongside bank fixed effects. The coefficient on POST is −0.0023 and statistically significant at the one percent level, indicating that provisioning intensity declined on average following the mandatory adoption of CAS 22. Taken in isolation, this result might appear to support the interpretation that the ECL transition reduced banks' provisioning relative to the pre-adoption period. However, this reading is premature. The R-squared of Model 1 is only 0.0603, meaning that the policy indicator alone explains less than seven percent of the within-bank variation in LLP. The unconditional correlation between POST and LLP reflects not only any direct effect of the standard but also the systematic co-movement of POST with macroeconomic conditions, bank business models, and profitability trends that evolved over the same period [16].

Model 2 introduces the full set of bank-level controls. The coefficient on POST becomes 0.0003 and is no longer statistically distinguishable from zero, while the R-squared rises sharply to 0.9638. The disappearance of the POST effect once controls are included indicates that the negative association observed in Model 1 was driven by the correlation between the post-adoption period and other bank characteristics rather than by the regulatory change itself. CAS 22 adoption, on this evidence, did not independently alter the level of provisioning intensity once observable bank-level factors are held constant [13].

The control variable estimates in Model 2 are both statistically and economically significant. The coefficient on EBLLP is 1.8337, significant at the one percent level, and constitutes by far the largest effect in the model. This finding is the empirical signature of

income smoothing: banks provision more when pre-provision profitability is high and less when it is low, consistent with managers using the discretion afforded by the ECL model to reduce reported earnings volatility rather than to reflect genuine changes in credit risk. The coefficient on ROA is -1.9273 , also significant at the one percent level. Because ROA is defined after provisions, its negative sign indicates that banks with higher post-provision profitability tend to provision less, which is consistent with the interpretation that provisioning is calibrated partly to manage the level of reported net income. The coefficient on LOAN_RATIO is -0.0262 , significant at the one percent level, suggesting that banks with more loan-intensive balance sheets provision proportionally less per unit of loans, possibly reflecting portfolio diversification or differences in underwriting standards across institution types. LEV carries a positive and significant coefficient of 0.0219 , indicating that more highly leveraged banks provision more, consistent with capital-constrained institutions using provisions as a buffer management tool. The coefficient on SIZE is -0.0005 and statistically insignificant, suggesting that bank scale does not independently explain provisioning intensity once the other controls are included.

4.2. Robustness Checks

Model 3 replaces the POST indicator with a full set of year fixed effects, implementing a two-way fixed effects specification that absorbs all aggregate time-varying shocks simultaneously. This addresses the concern that the POST coefficient in Model 2 might partially reflect common macroeconomic trends — such as the post-2019 economic slowdown or the pandemic-related macroeconomic adjustment — that coincide with but are independent of the CAS 22 transition. Because POST is a period-level indicator that takes the same value for all banks in a given year, it is absorbed entirely by the year fixed effects and cannot be separately identified in this specification.

The results of Model 3 are broadly consistent with those of Model 2, providing reassurance that the baseline findings are not an artifact of omitted time trends. The coefficient on EBLP rises slightly to 1.9213 and remains significant at the one percent level, reinforcing the income smoothing interpretation. The coefficient on ROA is -1.8670 , statistically significant and nearly identical in magnitude to Model 2, confirming that post-provision profitability management is a robust feature of the data. LOAN_RATIO remains negative and significant at -0.0283 . LEV strengthens to 0.0334 and attains one percent significance, which may reflect the year fixed effects absorbing aggregate leverage trends and sharpening the within-bank estimate. The coefficient on SIZE becomes -0.0021 and is now significant at the five percent level, suggesting that when common time trends are controlled for, larger banks do provision somewhat less per unit of loans, potentially reflecting advantages associated with systemic importance or superior diversification available to the largest institutions.

The overall R-squared declines from 0.9638 in Model 2 to 0.9281 in Model 3, a consequence of the year fixed effects absorbing variation that POST previously captured in the two-period structure of Model 2. The direction and significance of the substantive coefficients are nevertheless stable, which constitutes the primary robustness criterion. Taken together, the three models present a consistent picture: the adoption of CAS 22 did not independently alter provisioning intensity, and the dominant driver of LLP across all specifications in which it appears is pre-provision profitability, pointing to earnings management as the principal behavioral mechanism underlying asset impairment provisioning among Chinese listed commercial banks.

5. Discussion

The central finding of this study is that the mandatory adoption of CAS 22 did not independently alter the asset impairment provisioning behavior of Chinese listed commercial banks [17]. The POST coefficient, while negative and significant in the univariate specification, becomes economically trivial and statistically indistinguishable from zero once bank-level controls are introduced. This pattern aligns with broader

international evidence: accounting standard transitions alter the formal architecture of loss recognition without mechanically displacing the behavioral incentives that govern how managers exercise provisioning discretion. The dominant explanatory variable throughout is EBLLP, whose large positive coefficient indicates that banks provision more when pre-provision earnings are high and less when they are low — the textbook signature of income smoothing. Under the incurred loss model, the scope for such behavior was constrained by the requirement that losses be objectively evidenced before recognition. The ECL model, by requiring banks to form forward-looking judgments about default probabilities and credit risk staging, expanded the discretionary space in which earnings management operates. A notable observation is that a standard designed to make provisioning more informative appears, in the Chinese context, to have increased the sensitivity of provisioning outcomes to managerial judgment — particularly where supervisory oversight varies across institutions and ownership structures involve multiple layers of governance objectives.

The negative and significant coefficient on ROA adds a further dimension to this picture [9]. Banks with higher post-provision profitability tend to provision less, which is consistent with managers treating the provision charge as a lever for smoothing reported net income toward an implicit target rather than as a faithful estimate of expected credit losses. The leverage coefficient is positive and significant, indicating that more capital-constrained banks provision more heavily, a result that complements the income smoothing interpretation and suggests that provisioning behavior reflects a combination of motives — including earnings management, capital buffer management, and credit risk signaling — whose relative emphasis varies across institutions and time periods. These findings carry practical implications for regulators. If banks are using the subjective judgment embedded in the ECL model primarily to stabilize reported earnings rather than to improve credit risk transparency, the financial stability benefits anticipated from the CAS 22 implementation may not fully materialize. Strengthening supervisory review of ECL model inputs, requiring more granular public disclosure of staging decisions and forward-looking assumptions, and enhancing alignment between institutional governance frameworks and accounting transparency objectives are natural policy responses suggested by the evidence.

6. Conclusion

This paper examines whether the mandatory adoption of CAS 22 — China's domestic counterpart to IFRS 9 — altered the asset impairment provisioning behavior of A-share listed commercial banks over the period 2015 to 2025. Using a panel of 35 banks and 385 bank-year observations, fixed-effects models are estimated in which the POST indicator captures the shift from the incurred loss model to the expected credit loss framework. The results are unambiguous: POST carries a significant negative coefficient in the univariate specification but loses all statistical significance once bank-level controls are included, indicating that the regulatory transition itself did not independently drive changes in provisioning intensity. The dominant explanatory variable is pre-provision profitability, whose large positive coefficient is consistent with systematic income smoothing through credit loss provisions. Leverage and post-provision profitability also enter significantly, pointing to capital buffer management and earnings targeting as additional behavioral motives. These findings survive the introduction of two-way fixed effects in the robustness specification, confirming that the baseline results are not driven by aggregate time trends coinciding with the adoption window.

The broader implication is that the effectiveness of accounting standard updates in reshaping bank behavior depends critically on the incentive environment into which the standard is introduced. CAS 22 expanded the discretionary space available to bank managers by requiring forward-looking judgments that are inherently difficult to verify, but it did not alter the underlying institutional incentives — earnings volatility aversion, capital adequacy pressure, and state ownership structures — that shape how that discretion is exercised. Future research could fruitfully examine whether the behavioral

patterns documented here differ systematically across bank ownership types, or whether the growing use of third-party model validation and enhanced supervisory scrutiny in the post-2021 period has begun to constrain the income smoothing behavior this study identifies. Cross-country comparisons that leverage variation in the timing and enforcement quality of IFRS 9 adoption would also help to disentangle the effects of the standard itself from the institutional context in which it operates.

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