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* Views expressed are those of the author and do not necessarily reflect official positions of De Nederlandsche Bank.

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Abstract

How do unexpected changes in macroprudential capital buffer requirements impact bank valuation, measured by price-to-book ratios? This study addresses this question by constructing macroprudential capital buffer "surprises" from market reactions to buffer announcements and estimating their effects, using panel local projections, on the price-to-book ratios of a panel of large European banks. The analysis shows that unexpected buffer surprises are associated with a short-run decline in price-to-book ratios, followed by a sustained increase in the weeks following the announcement. Such an increase is consistent with market recognition of reduced risk, despite higher buffer requirements that could lower distributable resources, suggesting that the risk channel dominates the payout channel in the valuation of large European banks.

Keywords: Capital regulation; Macroprudential policy; Bank valuation
JEL Codes: G21, G28, G32

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1 Introduction

The setting of macroprudential capital buffers after the Global Financial Crisis (GFC) coincided with low bank valuations, raising the question of whether low valuations are linked to higher capital requirements. Since the GFC, the so-called price-to-book (PtB) ratio, which compares market capitalisation to book equity, has been depressed, especially in Europe. Many European banks have been trading below their book value, suggesting that investors view them as less able to deliver value [Martínez et al. 2024]. Compared with banks in other parts of the world, European banks have consistently shown lower PtB ratios, reinforcing concerns about their valuation [Bogdanova et al. 2018].

Industry voices contend that macroprudential capital buffer requirements depress bank valuations, arguing that if a bank must finance a larger share of its assets with equity rather than debt, its return on equity declines mechanically because profits need to be spread over a larger equity base. Assuming investors value banks as a stream of discounted dividends, this lower return should translate into a lower valuation attached to the bank. From this perspective, stricter macroprudential capital buffers should make banks less attractive to shareholders, but does this claim hold empirically? Do higher macroprudential buffers depress bank valuations?

Based on the literature, the expected relationship between regulation and bank valuation is either negative or negligible. For a panel of banks in the United States, Calomiris and Nissim [2014] find that the decline in PtB ratios since the GFC has been due not only to declines in several key determinants of the PtB ratio, but also to changes in the premium investors place on key variables such as leverage. Bogdanova et al. [2018] replicate the model of Calomiris and Nissim [2014] for an international sample of banks, concluding that the decline in PtB ratios has been the result of bank-specific variables, and that their findings ”cast some doubt on explanations that assign a large role to regulation as a source of low current bank valuations”. On the other hand, Chousakos and Gorton [2017], who also employ a panel data model to investigate the determinants of PtBs in a sample of US and European banks, argue that since the macro-economy and the zero lower bound do not explain the low PtB ratios, regulation is a key remaining candidate. Caparusso et al. [2023] put this to the test, estimating a panel data model of PtBs and, among others, a management buffer, and find little correlation between PtBs and management buffers, which casts ”doubt on arguments that more stringent capital requirements have driven the decline in G-SIBs' franchise values”. A limitation of this literature is that buffer requirements are treated as low-frequency policy variables whose variation is endogenous to macro-financial and banking conditions. As a result, existing estimates struggle to fully distinguish the valuation effect of the buffer policy itself from the factors that inform it.

This paper contributes to the literature on the relationship between macroprudential regulation and bank valuation in two key respects. First, the study has a relatively high degree of validity for the current policy environment, given its reliance on a sample of European banks during a period of active macroprudential policy in the current regulatory context. Second, to address concerns about the endogeneity in bank valuation and buffer policy, the paper identifies market-based buffer surprises that are plausibly exogenous. While constructed surprises follow a similar logic to Bluwstein and Patozi [2024], who construct binary sector-level shocks from these market responses, this study differs from this approach by identifying bank-level buffer surprises that vary in intensity, allowing estimation of dynamic valuation responses to continuous policy shocks. More broadly, this paper complements the existing literature on bank valuation and capital regulation by shifting the focus from low-frequency policy variation to high-frequency market-priced surprises. By doing so, it isolates the component of buffer policy that is orthogonal to contemporaneous information about bank fundamentals, addressing a key identification challenge.

The empirical analysis reveals a robust, statistically significant positive effect of an increase in the buffer on bank valuation, measured by the PtB ratio. This positive effect is not an artefact of the empirical strategy, but holds for both standard panel data models with fixed effects, in the vein of Calomiris and Nissim [2014], and for local projections using plausibly exogenous buffer surprises. While buffer increases are associated with a short-term decline in equity prices for a few days, this effect reverses over time, with PtB ratios rising significantly in the weeks following the announcement. The results hold up under several robustness checks, including sample restrictions and alternative specifications. Taken together, the evidence is consistent with a positive effect of unexpected buffer tightening on bank valuation.

The observed positive effect of increases in buffers on bank valuation suggests that the risk channel dominates the payout channel. The payout channel suggests that when banks need to increase equity or have lower distributable resources due to higher buffer requirements, this mechanically reduces the return on equity, and lower returns should, if all else remains equal, drive down valuations. On the other hand, the risk channel suggests that bank risk decreases as banks become less levered, thereby reducing the cost of equity and boosting valuation. The empirical literature has documented a negative relationship between capital strength and the cost of equity [Belkhir et al. 2021, deBandt et al. 2014, Berger et al. 2022, Kovner and Van Tassel 2022], suggesting that more conservative capital structures are associated with lower equity risk premia. The findings in this paper are consistent with the risk channel dominating the payout channel, which helps explain why increases in capital buffers may ultimately support rather than depress bank valuations.

The remainder of the paper is structured as follows. Section 2 documents the positive correlation between buffer requirements and bank valuation and motivates an identification approach. Section 3 describes the dataset. Section 4 presents the baseline results. Section 5 reports the results of the robustness analyses. Section 6 concludes.

2 Identification

Estimating the effect of macroprudential capital buffer requirements on bank valuation is challenging because buffers are policy instruments that respond endogenously to evolving macro-financial conditions and bank characteristics [Kim and Mehrotra 2022]. As a result, raw correlations between buffers and valuation may reflect not only the causal effect of buffer policy, but also the underlying risks and supervisory assessments that motivate buffer setting. To address this challenge, this section first documents the basic empirical regularity in the data and highlights why standard panel approaches are unlikely to yield causal estimates. It then develops a market-based identification strategy that exploits the joint response of equity returns and credit spreads around buffer announcements to distinguish policy-dominant buffer shocks from announcements dominated by other information, such as bank fundamentals.

A key initial question is whether banks facing different macroprudential buffer requirements also exhibit different market valuations. The sign of this relationship is ambiguous ex ante. Higher buffers reduce distributable resources, potentially depressing shareholder value in the short run. At the same time, higher buffers increase loss-absorbing capacity and may lower the perceived riskiness of bank equity, supporting valuation. Moreover, higher buffer calibration may target banks that are already more profitable, so that any cross-sectional association could reflect reverse causality. The evidence in this subsection is descriptive and intended to motivate the identification strategy that follows.

The analysis draws on a panel of large, listed European banks, using data from the European Banking Authority's (EBA) Transparency Exercise (TE) over 2014-2023.[1] Valuation is measured by the PtB ratio, aggregated to an annual level by averaging. The key policy variable is the combined buffer requirement (CBR), which is the sum of all relevant buffer requirements. The CBR is constructed from publicly available sources and bank disclosures. To account for phased-in components such as the capital conservation buffer, the CBR is demeaned by year. The resulting dataset contains 256 bank-year observations for 36 banks over 9 years.

Table 1 reports a sequence of specifications that progressively absorb heterogeneity: a

[1] The bank regulatory dataset is downloadable from: https://doi.org/10.5281/zenodo.19109425

pooled regression (model 1), a levels panel model with bank and year fixed effects (model 2), a first-difference with bank and year fixed effects (model 3), and a model that includes the CBR and headroom, defined as the management buffer held in excess of capital and buffer requirements (model 4). Across these models, the PtB ratio is positively associated with the demeaned CBR and with changes in the CBR. Similarly, the positive coefficient on headroom reported in model 4 aligns with the results of Caparusso et al. [2023], who find a positive association between headroom and the PtB ratio, even though there are differences in frequency, sample, variable definitions and controls. The results in Table 1 suggest that headroom and buffer requirements have separate effects, as the CBR remains significant and positive when including headroom. The fact that increasing the CBR could have positive, standalone effects beyond headroom contradicts the common industry narrative that buffers depress valuation.

Table 1: Regression results of PtB on the combined buffer requirement

Model:	(1)	(2)	(3)	(4)
Constant	0.775***			
	(0.024)			
CBR	0.105***	0.027**		
	(0.025)	(0.011)		
ΔCBR			0.021**	0.022***
			(0.007)	(0.005)
$\Delta Headroom$				0.009***
				(0.002)
Bank FE	No	Yes	Yes	Yes
Year FE	No	Yes	Yes	Yes
Observations	256	256	221	220
R^2	0.064	0.877	0.565	0.573

Notes: Model (1) is a pooled OLS regression without fixed effects. Model (2) is a linear panel data model that includes both time and bank fixed effects. Model (3) is a first-difference panel data model including both bank and time effects. Model (4) is a first-difference panel model that includes headroom. Standard errors are based on Driscoll and Kraay [1998]. Significance: ***: 0.01, **: 0.05, *: 0.1.

However, these results are, at best, conditional associations, and so do not reflect causality. A causal interpretation of buffer-valuation relationships requires variation in buffers that is orthogonal to developments in bank risk and profitability. Such orthogonality is unlikely to hold mechanically, as macroprudential buffers are designed - and shown [Kim and Mehrotra 2022] - to respond to macro-financial conditions. This endogeneity problem parallels the monetary policy identification problem, in which policy responds to the state of the economy, and announcements may reveal information about fundamentals ("information effects") in addition to delivering a policy impulse. A large literature therefore uses short-window market responses around announcements, together with structural restrictions on the joint movement of multiple asset prices, to separate policy shocks from information shocks [Jarociński and Karadi 2020, Altavilla et al. 2019, Nakamura and Steinsson 2018, Kerssenfischer 2022, Miranda-Agrippino and Ricco 2021]. This study uses a similar approach, employing high-frequency market responses to policy announcements to distinguish policy variation from other variation [Bluwstein and Patozi 2024].

The identification strategy exploits an institutional feature of the European framework: buffers are intended to absorb losses and primarily operate through distribution constraints rather than through immediate solvency triggers. Capital requirements are formal preconditions to conducting banking operations, and sustained breaches can lead to severe supervisory actions, including resolution or insolvency, which ultimately result in losses for creditors. Buffers, by contrast, sit above minimum requirements and are designed to restrict dividends and other forms of payouts when breached, without directly threatening authorisation. This institutional distinction implies that a buffer tightening can redistribute value between shareholders and creditors: tighter buffers reduce expected distributions to equity holders but increase protection for creditors by strengthening resilience.

This logic yields a classification of buffer announcements into policy shock or news shock, similar to that of Bluwstein and Patozi [2024]. A policy-dominant buffer tightening should be associated with a decline in equity prices (lower expected distributions) and a tightening of credit spreads (improved creditor protection). Conversely, news shocks dominate when an announcement window contains information that, for example, reduces expected profitability and is identified by deteriorating values of both equity and credit claims, where equity prices fall while credit spreads widen. Market reactions are classified into two categories: (i) buffer announcement surprises when equity returns and credit spreads move in the same direction (concordant moves), and (ii) news shocks when equity returns and credit spreads move in the opposite direction (discordant moves), which suggests that new information dominates the announcement window. Appendix A formalises this logic in a stylised model to clarify the conditions under which the sign identification should hold.

Two conditions are required for this approach to deliver plausibly exogenous buffer surprises. The first condition is that there should be no anticipation of buffer announcements. In practice, macroprudential authorities typically announce market-sensitive decisions before the market opens, subject to strict confidentiality constraints. The second condition is that the event must generate a repricing relative to prior beliefs. The empirical object is therefore the unexpected component (”surprise”) of the decision. A ”no-change” outcome can constitute an effective tightening or loosening depending on what the markets expect. This strengthens identification by conditioning on surprises rather than nominal policy moves, but also implies that estimated effect sizes should be interpreted as responses to market-priced surprises rather than actual changes in the policy variable.

3 Data

This section implements the identification strategy developed in Section 2 using a novel dataset of European macroprudential capital buffer announcements. The key challenge is to align policy decisions with market prices in a way that accurately captures the timing and informational content of buffer announcements. To this end, the section describes the construction of the announcement dataset, discusses the availability and suitability of market variables for measuring announcement-window responses, and details how bank-level buffer surprises are identified and merged with valuation and balance-sheet data.

To identify market responses to variations in buffer policy, the publication date of each announcement must be accurate. Measurement error in announcement timing would tend to bias the estimated response toward contemporaneous news shocks rather than policy shocks. To ensure announcement date accuracy, publication dates were extracted from the notification templates, cross-checked against other available data sources, and manually corrected where inconsistencies arose. Observations with inconsistencies were manually reviewed and corrected. For the buffers set for global systemically important institutions (G-SIIs), the publication date of the original G-SII list was used instead of the publication date of the European implementation.

Intraday timestamps are systematically unavailable for macroprudential announcements, and capital buffer decisions are published outside of trading hours. Therefore, this study uses market prices as close as possible to the announcement. For equities, the overnight return is defined as the difference between the previous day's close and the next day's opening price. For Credit Default Swap spreads, only daily prices are available, as intraday opening and closing quotes are not available. This identification strategy assumes that macroprudential authorities follow a pre-market publication convention. This assumption is supported by consultations with experts at the Dutch central bank (De Nederlandsche Bank; DNB) and checks on a subset of announcements with known publication dates and times.

The data collection yields 538 bank-announcement combinations. Applying the joint-response classification described in Section 2 produces 143 buffer surprises. Many announcements do not yield identifiable policy surprises because market reactions are consistent with prior expectations or because other information dominates the announcement window; importantly, some ”no change” decisions generate identifiable surprises, while some nominal buffer changes do not. This underscores that surprise is the unexpected component of announcements, as priced by markets, rather than the change in the policy variable.

Table 2: Descriptive statistics

Variable / Statistic	Mean	St. Dev.	Min	Max
PtB	0.782	0.341	0.120	1.900
Surprise	0.003	0.285	−9.780	45.210
RoE	0.060	0.047	−0.158	0.314
Leverage ratio	0.055	0.015	0.000	0.125
Asset density	0.055	0.015	0.000	0.125
CET1-ratio	0.143	0.027	0.075	0.251

Notes: The table reports descriptive statistics for the main variables used in the empirical analysis. The price-to-book ratio (PtB) is sourced from Refinitiv and measured at the daily frequency. The surprise is defined as the change in a bank's 5-year CDS spread on the day of a macroprudential buffer announcement, with the sign adjusted to reflect the policy surprise's direction. Return on equity (RoE), leverage ratio, asset density, and CET1 Ratio are derived from the EBA Transparency Exercise and matched at the bank-date level, with missing values filled forward. The final dataset includes 20 large European banks and spans 47,556 daily observations.

The key shock variable is the buffer announcement surprise, simply labelled as the "surprise" throughout this study, which is defined as the announcement-day change in a bank's 5-year CDS spread, with the sign aligned with the buffer surprise as implied by the joint equity-CDS response. Positive realisations correspond to tighter-than-expected buffer surprises (market-implied tightening), while negative realisations correspond to looser-than-expected surprises. Using the continuous CDS change rather than a 0/1 event dummy preserves variation in shock intensity, which is valuable given the limited number of identified events.

Buffer surprises are merged with daily valuation and balance-sheet variables to form a long daily panel suitable for local projections. PtB ratios are sourced from Refinitiv at a daily frequency. Balance sheet controls - return on equity, leverage ratio, asset density, and the common equity tier 1 (CET1) ratio - are sourced from the EBA TE and matched at the bank-date level, with missing values carried forward. The final dataset contains 20 large European banks and spans 47,556 daily observations. The sample is restricted to banks with available and varying CDS spreads; this limits external validity to larger institutions but is fundamental to a CDS-based identification strategy.

4 Results

This section presents the empirical results of local projections of PtBs on surprises, using the dataset from the previous section. The first subsection introduces the baseline specification. The second subsection presents the results.

4.1 Baseline specification

To trace the dynamic response of the PtB ratio to a surprise at various horizons, the baseline specification uses panel local projections [Jordà 2005] of the form:

$$PtB_{it+h} = \alpha_i^h + \gamma_t^h + \beta^h Surprise_{it} + \omega^h Q_{it}^h + \epsilon_{it}^h \tag{1}$$

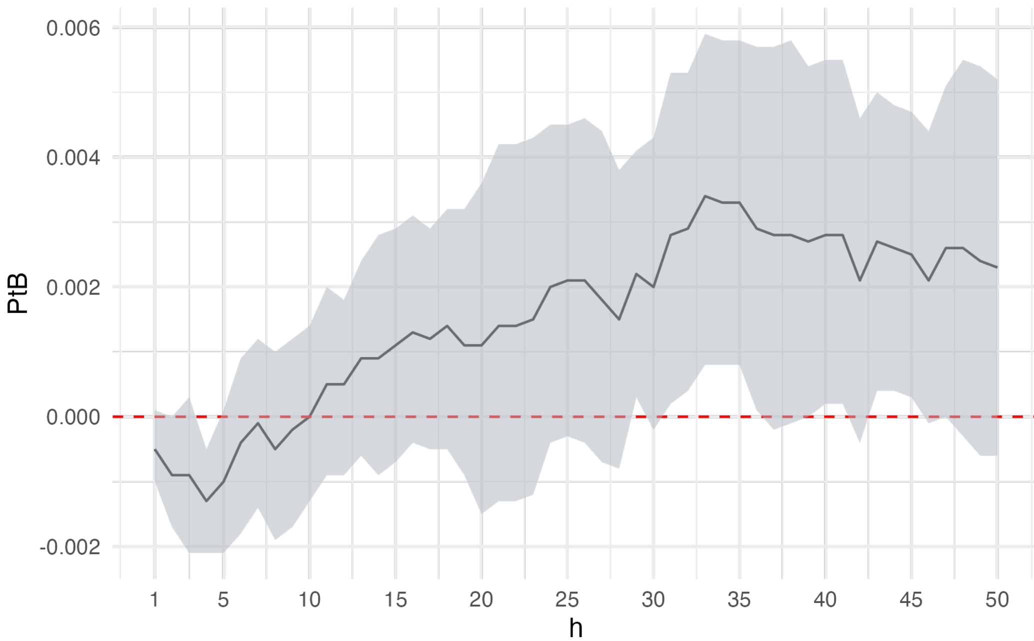
where $i = 1, 2, 3, ..., N$ denotes the bank index, $t = 1, 2, 3, ..., T$ indexes the daily observation period, and $h = 1, 2, 3, ..., H$ denotes the forecast horizon, where $H = 50$ trading days corresponds to approximately ten weeks ahead. The dependent variable is the PtB ratio at horizon h, with the $Surprise_{it}$ constructed as described in Section 3.

To control for a large number of confounding factors, the specification includes time fixed effects, γ_t^h, and unit fixed effects, α_i^h, and a control matrix Q_{it}, which contains the CET1 ratio, the leverage ratio, the return on equity, and the asset density, leads of the surprise for all periods h, and five lags for the surprise and dependent variable - aligning with the recommendations of Alloza et al. [2025]. The object of interest is the impulse response function (IRF) of the surprises on the PtB ratio, given by a vector of β^h for every h, surrounded by a 95% confidence interval (CI) based on Driscoll and Kraay [1998] standard errors that are robust to temporal and cross-sectional dependence.

4.2 Baseline result

Figure 1 displays the IRF of the PtB ratio to a 1 p.p. surprise, surrounded by a 95% CI, based on the model in equation 1. The PtB declines over the first few days following the

Figure 1: The effects of a 1 p.p. surprise on the PtB ratio



Notes: Impulse response of the price-to-book ratio to a 1% surprise estimated using panel local projections. Shaded areas indicate 95% CIs based on Driscoll-Kraay standard errors. The specification is estimated with a full set of controls, as described in Section 4.1.

announcement, then reverses and rises steadily to a net positive effect. In the very short term (up to 5 days), an increase in capital buffers reduces bank valuation, consistent with the payout channel: higher buffers lower distributable capital, thereby reducing return on equity.

Beyond the five-day horizon, the PtB ratio increases significantly following increases in buffers, consistent with a reassessment of the riskiness of bank equity. This pattern is consistent with the existence of the risk channel: higher buffer requirements reduce perceived bank risk, lowering the cost of equity and boosting valuation. Existing evidence shows that stronger capital positions are associated with lower equity risk premia [Belkhir et al. 2021, deBandt et al. 2014, Berger et al. 2022, Kovner and Van Tassel 2022]. In this setting, the risk channel appears to dominate the mechanical dilution of return on equity implied by the payout channel.

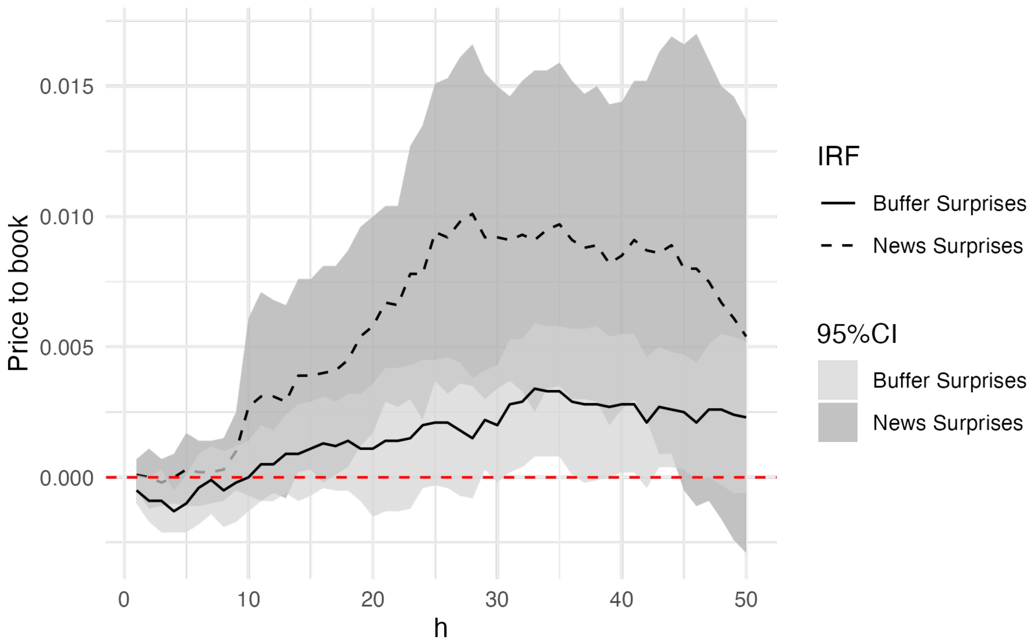
This dynamic pattern contrasts with the predominantly negative or insignificant effects reported in panel-based studies of bank valuation and capital regulation. Existing work finds negative [Chousakos and Gorton 2017], zero [Bogdanova et al. 2018], or negligible [Caparusso et al. 2023] valuation effects of capital regulation. A key reason for this difference is methodological. These studies rely on low-frequency panel regressions in the vein of Calomiris and Nissim [2014], whereas this paper exploits high-frequency, market-based buffer-surprise data within a local projections framework. In addition, this study focuses on large European banks subject to active macroprudential policy and examines market responses at daily rather than quarterly frequencies. These differences allow the analysis to uncover dynamic valuation effects that are difficult to detect in standard panel settings.

These results provide evidence on the direction rather than the effect size of buffer policy, as the identification scheme attempts to capture the surprise element of such a change. The estimated response of the PtB ratios to a surprise is therefore typically small: the effect of a 1 p.p. surprise on the PtB ratio reaches 0.0034 at its peak. Even when accounting for the size of the surprise, with a one standard deviation of the surprise equal to ca. 5 p.p., the IRF peaks at 0.0147, which is far from the fairly sizeable effects found using dynamic panel data models, as shown in Section 2.

5 Robustness analyses

The additional robustness analyses that were performed fall into one of three categories, which are presented in the following subsections. Subsection 5.1 addresses several concerns regarding identification. Subsection 5.2 presents the results of several alternative specifications. Subsection 5.3 shows the results of several alternative sample definitions. Across all robustness checks, the qualitative pattern of an initial decline followed by a persistent increase in the PtB ratio remains intact.

Figure 2: The effect of news shocks versus surprises on the PtB ratio



Notes: Impulse responses of the PtB ratio to surprises (solid line with light-grey confidence band) and news shocks (dashed line with dark-grey confidence band). Shaded areas denote 95% CIs based on Driscoll-Kraay standard errors. All specifications include the controls as described in the baseline specification, per Section 4.1. In addition, up to $h-1$ leads and five lags of the news surprises shocks were included.

5.1 Identification

There may be some bias in the result, because equity returns are used to identify surprises while also affecting the PtB ratio. At the same time, the equity return influences the PtB because the share price is the only factor in the PtB ratio that varies frequently. Indeed, a regression of first-differenced PtB on overnight returns yields a coefficient of 0.48, significant at the 1% level. However, if the effects of a surprise on the PtB ratio were biased through the negative equity return condition, one would expect a sizeable response. Instead, we observe a small response at $h = 0$, (point estimate: -0.0001, p-value: 0.053). This suggests that any bias, if existent, due to the sign restriction would likely be small. Additionally, note that the directional association is much less straightforward: Kendall's tau, which is a correlation measure based on same- and opposite-directional comovement of two variables, measured between overnight return and first-differenced PtB is 0.35, indicating that in the large majority of cases, the PtB and overnight return move in opposite directions.

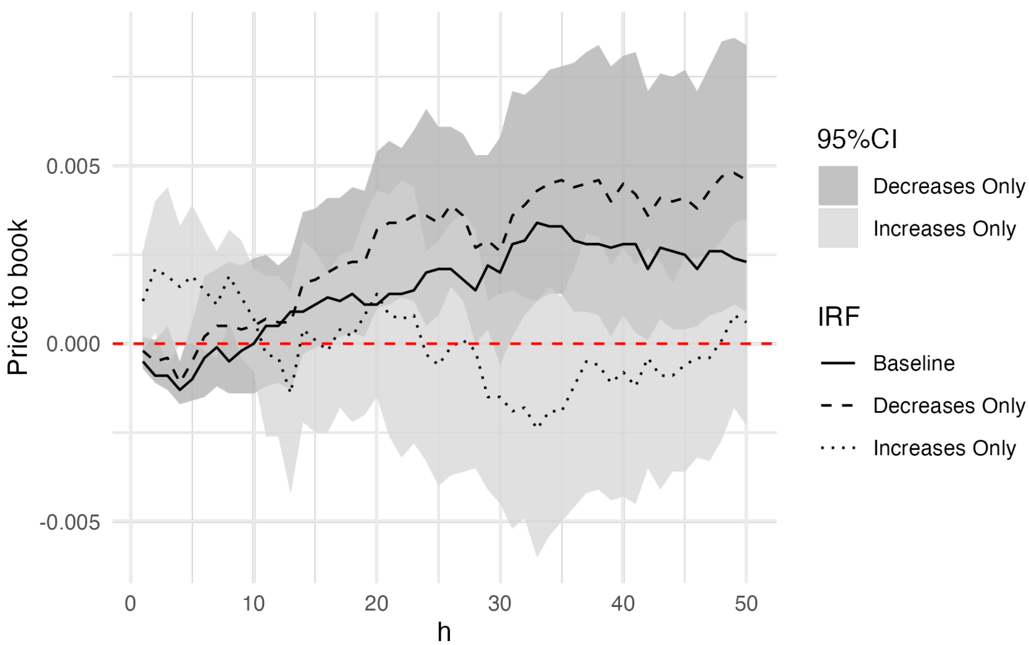
A second concern is whether the identification scheme truly elicits distinct information from the buffer announcements. If the responses of the PtB to news shocks and surprises are highly similar, then using signs on market variables does not yield new information, and one could instead rely on the market response to the buffer announcement rather than on announcements that meet certain conditions. To investigate this, the baseline specification is estimated for news shocks and plotted next to the baseline results in Figure 2, which compares the IRF of the PtB to a news shock and surprises, showing that the PtB responds more strongly to news shocks than surprises, which suggests that using the sign of the equity return to distinguish between news and policy shocks produces distinct responses.

5.2 Alternative specifications

To address the concern about asymmetry between market-implied increases and decreases, the IRF is allowed to vary across positive and negative surprises by interacting the surprise dummies with the surprises in the baseline specification. The results, displayed in Figure 3, suggest that while the IRFs between increases and decreases vary, the confidence intervals of each IRF overlap with the baseline IRF, implying no significant differences between the positive and negative IRFs. The wide confidence intervals are related to the relatively small sample size and the prevalence of negative (81) over positive (61) surprises. Overall, these results are not suggestive of strong asymmetric effects.

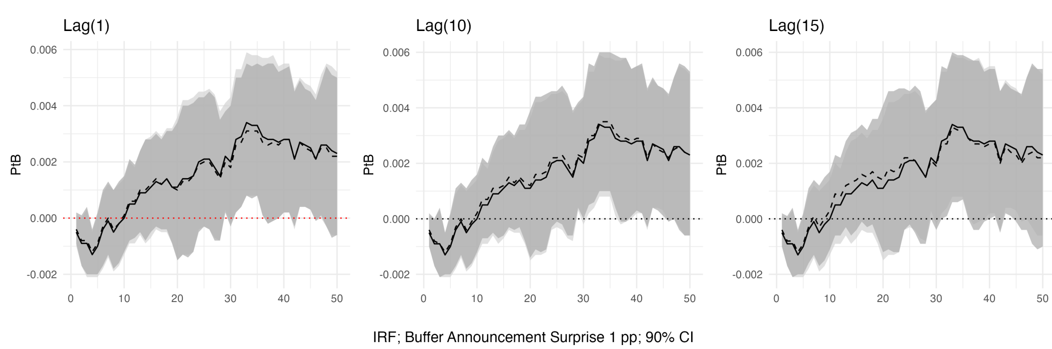
The robustness of the result to alternative lag orders is established by estimating the baseline specification with several alternative lag orders $[1, 10, 15]$. The results displayed in Figure 4 show that the main result of an initially negative response and a consequent increase of the PtB after a positive surprise is confirmed.

Figure 3: The effects of surprises implying increases and decreases compared to the baseline effects



Notes: Impulse response functions of the price-to-book ratio to surprises in the baseline specification (solid line), to market-implied buffer decreases only (dashed line with dark-grey confidence band), and market-implied buffer increases only (dotted line with light-grey confidence band). Responses for decreases are inverted in sign to ensure comparability with the baseline results. Shaded areas denote 95% CIs. All specifications include the controls as described in the baseline specification, per Section 4.1, except that surprises are replaced by increases and decreases are lagged up to five periods, and up to $h-1$ leads are included.

Figure 4: The effects of a 1 p.p. surprise on the PtB ratio for various lag orders

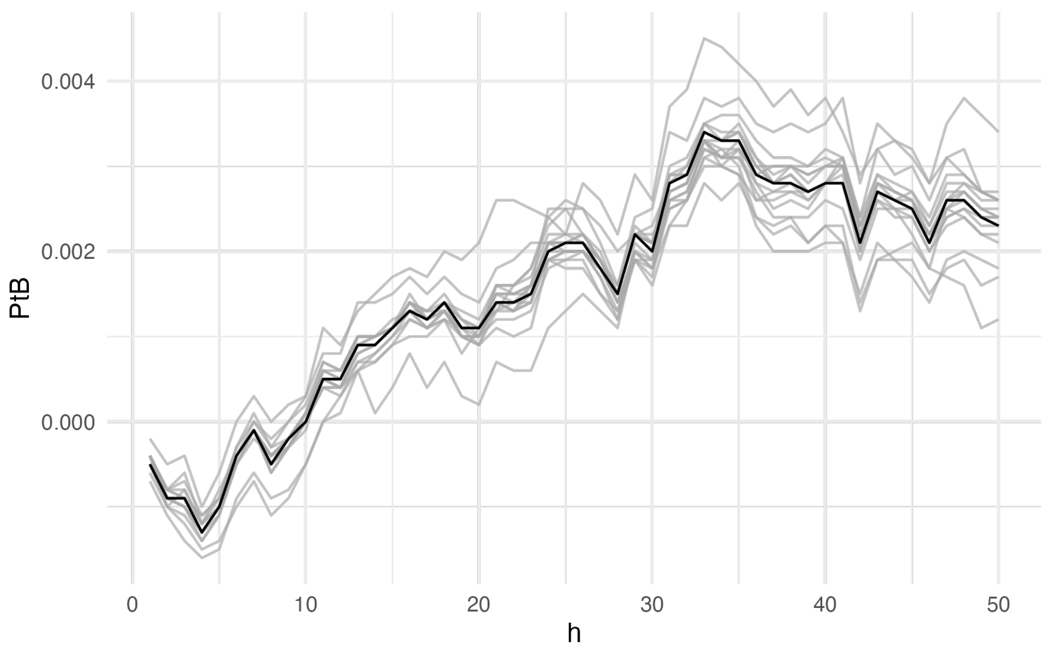


Notes: Impulse responses of the price-to-book ratio to a 1% surprise estimated using panel local projections under alternative lag specifications. The solid line with a light-grey confidence band shows the baseline specification, while the dashed lines with dark-grey confidence bands correspond to specifications with alternative lag lengths $[1, 10, 15]$. Shaded areas denote 95% CIs. Apart from alternative lag orders, the specifications are in line with the baseline specification, per Section 4.1.

5.3 Alternative samples

Given the relatively low number of shocks (143) and banks (20), there is concern that a single bank with idiosyncratic shocks may drive the results. To address this concern, a leave-one-bank-out procedure is applied, in which one bank is iteratively excluded from the sample, and the baseline is re-estimated on the new sample. These 20 alternative IRFs are displayed in Figure 5 and closely track the baseline results, implying no single bank does drives the results.

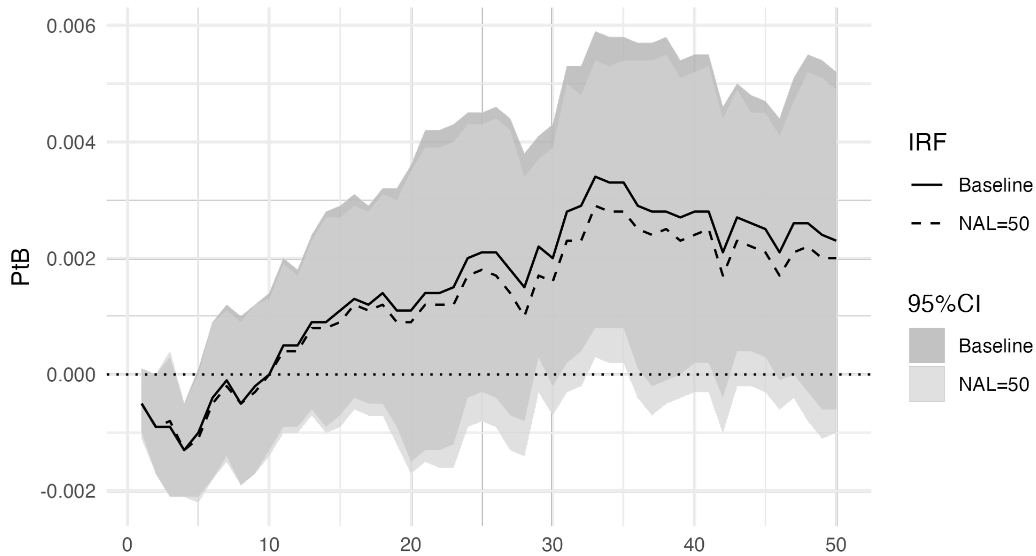
Figure 5: The effects of a 1 p.p. surprise on the PtB ratio when iteratively dropping one bank



Notes: Impulse responses of the price-to-book ratio to a 1% surprise, estimated using a leave-one-bank-out procedure. Each thin line represents an impulse response obtained by excluding one bank from the sample, while the solid black line shows the baseline specification. All specifications include the controls as described in the baseline specification, per Section 4.1.

The sample contains several invalid comparisons: banks that have recently been "treated" with a surprise are used as a "control" for banks that are treated with a surprise in the current period, which can result in biased estimates [Dube et al. 2025, Sant'Anna and Zhao

Figure 6: The effects of a 1 p.p. surprise on the PtB ratio with 50 days of data after an announcement excluded



Notes: Impulse responses of the price-to-book ratio to a 1% surprise estimated using panel local projections. The solid line with dark grey confidence bands shows the baseline specification. In contrast, the dashed line with a light-grey confidence band excludes observations for banks that experienced an unexpected buffer change in the preceding 50 trading days. Shaded areas denote 95% CIs. All specifications include the controls as described in the baseline specification, per Section 4.1.

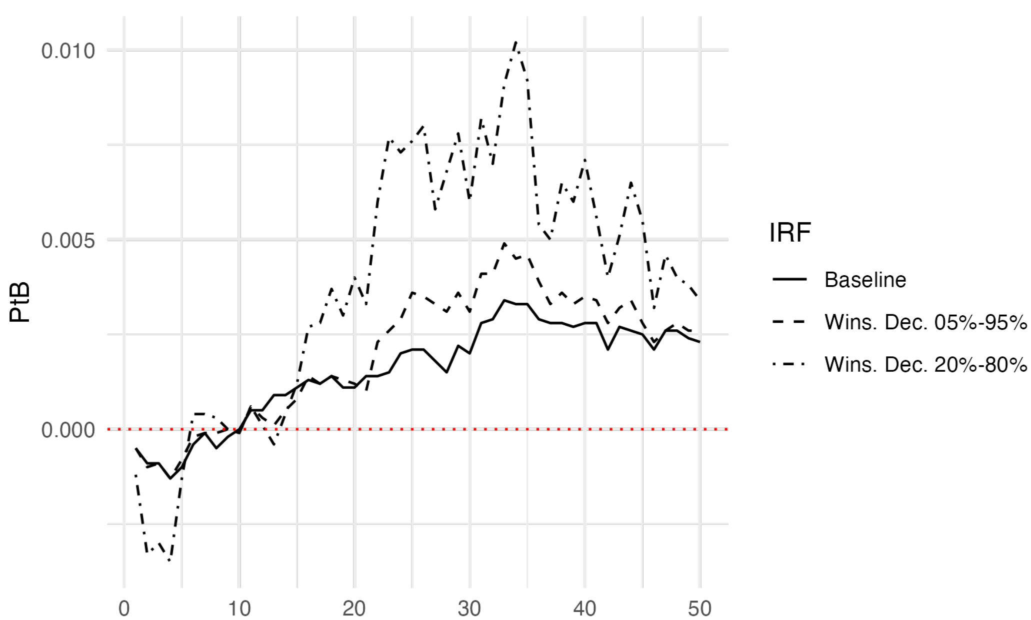
2020]. Unfortunately, no readily available implementable estimator handles both invalid comparisons and a staggered continuous treatment. As a solution, a non-absorbing lag, in the vein of Dube et al. [2025], was applied up to the 50-day horizon, which drops all banks from the sample for the 50 days after a buffer surprise. The results are displayed in Figure 6, showing high similarity between the baseline results and the reduced invalid comparison, with a slight decrease in significance and a slightly less pronounced IRF. Qualitatively, however, the results remain the same, suggesting they are not due to invalid comparisons.

The sample contains several highly volatile periods resulting from large common shocks, most notably Covid-19 and the invasion of Ukraine, which may affect the baseline IRFs. To ensure the robustness of the baseline IRFs to periods with large common shocks, the onset of Covid-19 (from March 11, 2020, to June 11, 2021) and the Ukraine invasion (from February

24, 2022, to May 24, 2022) are excluded from the sample, and the baseline specification is re-estimated. The resulting IRF is displayed in Figure 8, which displays results that are highly similar to the baseline. However, there is a slight decrease in significance, as expected when a substantial share of the data is excluded, leading to the conclusion that the results are not attributable to large common shocks.

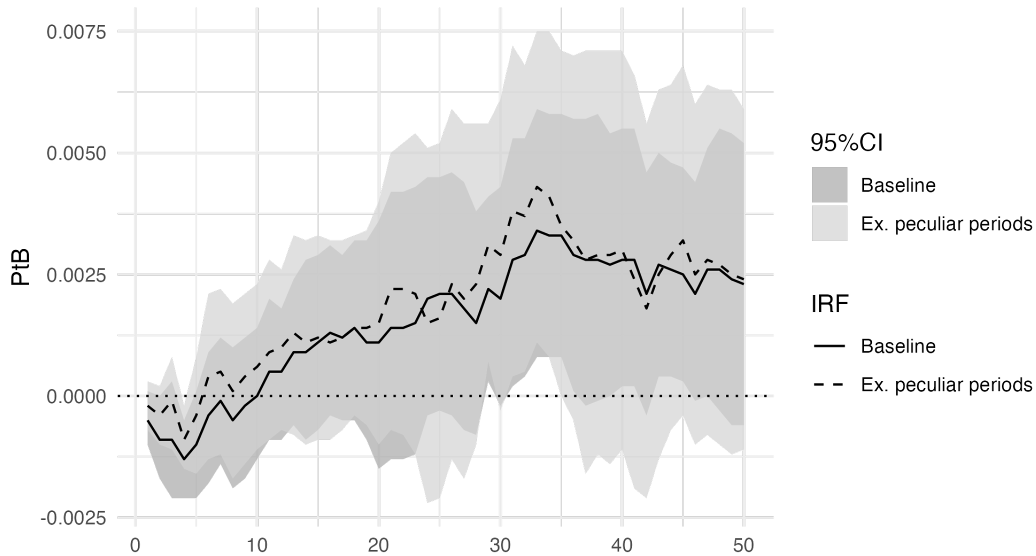
Due to the relatively low number of surprises (143) and the wide dispersion in surprise sizes, the baseline results may be affected by outliers. To reduce reliance on outliers without causing a major reduction in sample size, the data were winsorised at the 5%-90% and 20%-80% percentiles, and the baseline was re-estimated. The resulting IRFs are displayed in Figure 7, which shows that the baseline results are conservative, whereas winsorisation produces progressively more pronounced IRFs.

Figure 7: The effects of a 1 p.p. surprise on the PtB ratio with surprises winsorised at various levels



Notes: Impulse responses of the price-to-book ratio to a surprise, estimated using panel local projections. The solid line shows the baseline specification, while dashed lines correspond to specifications in which the surprises are winsorized at the 5-95% and 20-80% percentiles. All specifications include the controls as described in the baseline specification, per Section 4.1.

Figure 8: The effects of a 1 p.p. surprise on the PtB ratio with peculiar periods excluded



Notes: Impulse responses of the price-to-book ratio to a surprise, estimated using panel local projections. The solid line with dark grey confidence band shows the baseline specification. In contrast, the dashed line with the light-grey confidence band corresponds to a specification that excludes periods of heightened market volatility associated with the Covid-19 outbreak (March 11, 2020, to June 11, 2021) and the Ukraine invasion (February 24, 2022, to May 24, 2022). Shaded areas denote 95% CIs. All specifications include the controls as described in the baseline specification, per Section 4.1.

6 Conclusion

This paper studies how macroprudential capital buffer requirements affect bank valuation in the European post-crisis regime. The central challenge is identification: buffer policy responds to macro-financial conditions and bank characteristics, and announcement windows can bundle policy actions with information about fundamentals. To address this, the paper constructs surprises to buffer requirement announcements using the joint response of equity returns and CDS spreads around buffer announcements, and estimates dynamic valuation effects using panel local projections.

The results point to a clear dynamic pattern. Following an unexpected buffer tightening, PtB ratios decline initially, consistent with a short-run payout channel in which higher buffers reduce expected distributions. This effect reverses over subsequent weeks and turns persistently positive at longer horizons, consistent with a risk channel in which stronger buffers reduce the perceived riskiness of bank equity. The qualitative pattern is robust across alternative specifications, lag choices, and sample restrictions. This pattern is difficult to reconcile with a dominating role for the payout channel, which predicts unambiguously depressed bank valuations due to higher buffers.

From a policy perspective, these findings do not support concerns about the adverse valuation effects of macroprudential buffers for large European banks operating under the current framework over the recent period. Instead, this study provides evidence that macroprudential buffers are associated with higher bank valuations, consistent with the risk channel dominating the payout channel. Future work could sharpen the mechanism by linking buffer policy directly to the payout and risk channels, for instance, by estimating the impact of buffer policy on distributable resources and bank risk from financial statement data, and relating these to bank valuation.

References

Alloza, M., Gonzalo, J., and Sanz, C. (2025). Dynamic effects of persistent shocks. *Journal of Applied Econometrics*, 40(4):380–394.

Altavilla, C., Brugnolini, L., Gürkaynak, R. S., Motto, R., and Ragusa, G. (2019). Measuring Euro area monetary policy. *Journal of Monetary Economics*, 108:162–179.

Belkhir, M., Ben Naceur, S., Chami, R., and Samet, A. (2021). Bank capital and the cost of equity. *Journal of Financial Stability*, 53.

Berger, A. N., El Ghoul, S., Guedhami, O., and Roman, R. A. (2022). Geographic deregulation and banks' cost of equity capital. *Journal of International Money and Finance*, 120:102498.

Bluwstein, K. and Patozi, A. (2024). The effects of macroprudential policy announcements on systemic risk. *Bank of England Staff Working Paper 1080*.

Bogdanova, B., Fender, I., and Takats, E. (2018). The ABCs of bank PBRs. *BIS Quarterly Review, March*.

Calomiris, C. W. and Nissim, D. (2014). Crisis-related shifts in the market valuation of banking activities. *Journal of Financial Intermediation*, 23(3):400–435.

Caparusso, J., Lewrick, U., and Tarashev, N. A. (2023). Profitability, valuation and resilience of global banks: A tight link. *BIS Working Paper 1114*.

Chousakos, K. T. and Gorton, G. B. (2017). Bank health post-crisis. *NBER Working Paper 23167*.

deBandt, O., Camara, B., Pessarossi, P., and Rose, M. (2014). Regulatory changes and the cost of equity: Evidence from French banks. *Débats économiques et financiers 11*.

Driscoll, J. C. and Kraay, A. C. (1998). Consistent covariance matrix estimation with spatially dependent panel data. *The Review of Economics and Statistics*, 80(4):549–560.

Dube, A., Girardi, D., Jordà, O., and Taylor, A. M. (2025). A local projections approach to difference-in-differences. *Journal of Applied Econometrics*, 40(7):741–758.

Jarociński, M. and Karadi, P. (2020). Deconstructing monetary policy surprises: The role of information shocks. *American Economic Journal: Macroeconomics*, 12(2):1–43.

Jordà, O. (2005). Estimation and inference of impulse responses by local projections. *American Economic Review*, 95(1):161–182.

Kerssenfischer, M. (2022). Information effects of euro area monetary policy. *Economics Letters*, 216:110557.

Kim, S. and Mehrotra, A. (2022). Examining macroprudential policy and its macroeconomic effects: Some new evidence. *Journal of International Money and Finance*, 128:102697.

Kovner, A. and Van Tassel, P. (2022). Evaluating regulatory reform: Banks' cost of capital and lending. *Journal of Money, Credit and Banking*, 54(5):1313–1367.

Martínez, F. G., Jiménez, J. D., and Matas, R. Q. S. (2024). Reevaluating bank price-to-book ratios: An in-depth analysis of equity components across economic cycles. *Journal of Risk and Financial Management*, 17(8):363.

Miranda-Agrippino, S. and Ricco, G. (2021). The transmission of monetary policy shocks. *American Economic Journal: Macroeconomics*, 13(3):74–107.

Nakamura, E. and Steinsson, J. (2018). High-frequency identification of monetary non-neutrality: The information effect. *The Quarterly Journal of Economics*, 133(3):1283–1330.

Sant'Anna, P. H. C. and Zhao, J. (2020). Doubly robust difference-in-differences estimators. *Journal of Econometrics*, 219(1):101–122.

A Appendix

Below follows a short, stylised model of the market response to a buffer announcement. The main goal of the model is to illustrate that using common assumptions, the identification assumption as stated in Section 2 emerges: a change in the buffer requirement results in a same-directional move of equity returns and credit spreads, while a change in risk results in an opposite-directional move of equity returns and credit spreads.

A.1 Set-up

The setup includes three agents: a credit investor, an equity investor, and a bank manager. The bank has credit and equity instruments in issue, which are priced by credit and equity investors. The model has three periods, $t = 0, 1, 2$. All agents are risk-neutral.

A.2 The bank and its regulation

The banker can invest in a single project. This project has two states in each period: a good state and a bad state. In the good state it generates return R_t^H with probability $1-p$, while it generates return R_t^L with probability p, in periods $t = 1, 2$. The return in the bad state is negative: $R_t^L < 0$. The bank is subject to a regulation that allows it to distribute dividends, D_t, only if it meets its combined buffer requirement, CBR, and its capital requirement, CR. The CR is constant for all periods, but the CBR can vary over time, taking two possible values: the initial value CBR_0 and a new value, CBR_n, arriving in the $t = 1$ period. The capital ratio of the bank in the initial period, $C_{t=0}$, is assumed to be sufficient to meet its initial requirements ($CBR_0 + CR$), exactly, so that: $C_{t=0} = CR + CBR_0$. While this appears to be a fairly strict assumption, looser definitions would involve adding an arbitrary management buffer on top of the capital buffer and requirement, after which the model would proceed with the same logic. The bank distributes all dividends above its capital ratio, where dividends in period ($t = 1, 2$) are defined as follows:

$$\begin{cases} D_{t+h} = C_{t-1} - CBR_n - CR + R_t^i & \text{if } C_{t-1} > CBR_n - CR + R_t^i \\ D_t = 0 & \text{if } C_{t-1} < CBR_n - CR + R_t^i. \end{cases} \quad (2)$$

Note that the model assumes that the CBR is insufficient to cover the losses; $CBR_n < |R_{t=2}^L|$ and $CBR_0 < |R_{t=1}^L|$. This ensures the setup incorporates default risk. If the CBR was sufficient to cover all losses, creditors would not be running any default risk. Additionally, should the bank fail to meet the regulatory minimum requirement (CR), its banking license is withdrawn, after which bank creditors incur losses. The bank will enter insolvency, and creditors will be confronted with deadweight losses, N, with $N < 0$, that wipe out part of the credit investor's claim, implying $CBR_n + CR + N < 0$. [2]

A.3 The equity investor

The equity investor prices equity securities as the sum of discounted cash flows, i.e. dividends, where the investor discounts expected dividends at rate r. The assumption is that r is constant across periods. Summarizing the discount factor as $\delta = \frac{1}{1+r}$, yields the following equation for the equity price:

$$P_t = \delta E[D_{t=1}] + \delta^2 E[D_{t=2}]. \quad (3)$$

The two expected dividend terms, $E[D_{t=1}]$ and $E[D_{t=2}]$, can be easily defined when the equity investor is risk-neutral. For the first period's expected dividend, multiply the probability of

[2]Such deadweight losses are common in insolvency procedures, as liquidation of the estate typically results in liquidation values below book value, and even when creditors recover a large share of the notional value of their debts, the time it takes to obtain those recoveries reduces the value further in economic terms due to opportunity costs.

the good state by the probability of the good state. This follows from equation 2.

$$E[D_{t=1}] = (1-p)[R^H_{t=1} + C_{t=0} - CBR_n - CR]. \tag{4}$$

For the second period ($t = 2$) dividend, the probability of the good-good state multiplied by the amount of dividend distributed is equal to the dividend paid out in the good-good state, since in the bad-good state or in the good-bad state, there are no dividends to distribute. This gives:

$$E[D_{t=2}] = (1-p)(1-p)[R^H_{t=2} + C_{t=0} - CBR_n - CR]. \tag{5}$$

A.4 The credit investor

The credit investor holds a credit claim on the bank, and prices default risk on that claim. As mentioned in section A.2, the bank fails if it cannot meet its regulatory minimum requirement, CR, resulting in deadweight losses to the credit investor's claim. The risk-neutral credit investor prices the credit security through expected loss, charging a spread equal to the expected credit loss. To obtain an explicit definition of that expected loss, note that there are two cases in which the bank defaults: in the bad state in the second period ($t = 1$), as losses exceed the buffer, and in the good-bad state in the third period ($t = 2$), as losses will then also exceed the buffer. The credit spread can then be defined as follows:

$$CS_t = -[pL_{t=1} + p(1-p)L_{t=2}], \tag{6}$$

where $L_{t=1} = min(0; C_{t=0} + R^L_{t=1} + N)$ and $L_{t=2} = min(0; CBR_n + CR + R^L_{t=2} + N)$. The credit spread pricing equation states that the credit spread is equal to the expected loss in the first period, defined as the capital ratio from the first period minus the negative return and deadweight losses, and the expected loss in the third period, equal to the capital buffer requirement and capital requirement in the second period minus the low return and the deadweight losses occurring in insolvency. Note that the negative sign implies that as losses increase, the (positive) credit spread, CS_t, increases.

A.5 Results

The model has a straightforward solution, as the bank invests in a single asset, and credit and equity spreads result from changes in the buffer requirement. The implications can therefore be derived by differentiating with respect to the macroprudential buffer requirements to examine the responses of market variables.

Proposition 1 *Announcement of a CBR change causes a concordant move of the equity price and credit spread.* When the capital buffer requirement is increased (decreased), the equity price declines (increases) $\frac{\partial P_t}{\partial CBR_n} < 0$. When the capital buffer requirement is increased (decreased), the credit spread declines (increases), $\frac{\partial CS_t}{\partial CBR_n} < 0$.

This proposition states that a change in buffer requirements will produce a concordant change in the equity price and the credit spread. Such a concordant move, therefore, identifies cases in which the other key variable, such as risk (i.e., the bad-state probability), does not change. These concordant changes can be used in an empirical setting to identify the announcements that are not affected by other events, such as major changes in risk. Intuitively, an increase in the buffer requirement is bad for shareholders because it reduces dividends, but good for creditors because it increases the amount of protection if the bank fails. Therefore, this identification logic is applicable specifically to buffer requirements, and not necessarily to capital requirements more generally.

Proof

Determining the response of the equity price, P_t, with respect to a change in the CBR_t, can be done by differentiating the equity price equation, as per equation 3, with respect to the CBR_n. This gives:

$$\frac{\partial P_t}{\partial CBR_n} = \frac{\partial}{\partial CBR_n}[\delta(1-p)[R^H_{t=1} + C_{t=0} - CBR_n - CR] + \delta^2(1-p)(1-p)[R^H_{t=2} + C_{t=0} - CBR_n - CR]], \quad (7)$$

$$\frac{\partial P_t}{\partial CBR_n} = -\delta(1-p) - \delta^2(1-p)^2 < 0. \quad (8)$$

Given that $p > 0$ and $\delta = \frac{1}{1+r} > 0$ are strictly positive, equation 8, is strictly negative.

Similarly, the change in the credit spread can be obtained by differentiating 6 with respect to CBR_n, which gives:

$$\frac{\partial CS_t}{\partial CBR_n} = -(1-p)p < 0. \quad (9)$$

As $p > 0$, the outcome of this equation is strictly negative. Both $\frac{\partial P_t}{\partial CBR_n}$ and $\frac{\partial CS_t}{\partial CBR_n}$ are strictly lower than zero, and therefore carry an identical (negative) sign.

Proposition 2 *An increase of risk causes a discordant move of the equity price and credit spread.* When the probability of the bad state increases (decreases), the equity price decreases (increases) $\frac{\partial P_t}{\partial p} < 0$. When the probability of the bad state increases (decreases), the credit spread increases (decreases); that is, $\frac{\partial CS_t}{\partial p} > 0$.

The proposition states that a change of the bad-state probability, p, which is also referred to as ”risk”, produces discordant moves in the equity price and the credit spread. Essentially, both credit (i.e., bond) and equity securities decline in value when exogenous risk increases. However, this produces a discordant move in the equity return (decrease) and the credit spread (increase), because of the relationship between yields and bond prices.

Intuitively, the proposition states that an increase in risk is bad for both shareholders and creditors, as it increases overall risk. In an empirical setting, it suggests that when there is discordance between credit spreads and equity returns, the bank's risk profile changes. As such, discordant moves after an announcement of a macroprudential buffer change suggest a change in risk, either due to the announcement by the macroprudential authority or to an alternative event occurring at the same time. Such discordant responses to a macroprudential buffer announcement cannot be used as exogenous changes in the policy variables in an empirical setting.

Together, these two propositions can be used to separate buffer announcements into those in which new risk information dominates and those in which the change in the policy variable dominates. However, when implementing such an approach, the policy variable is measured not in nominal terms but relative to market expectations. That is, *not* increasing the buffer to the expected level is measured as a buffer *decrease*. Similarly, keeping a buffer constant while market participants expected a *decrease* is measured as an increase. This matters when interpreting the empirical results, see Section 4.2.

Proof The response of the equity price to a change in risk can be obtained by differentiating the equity price equation, as specified in equation 3, with respect to the probability of a project failure, p. This gives:

$$\frac{\partial P_t}{\partial p} = -\delta D_{t=1} - \delta^2 2(1-p) D_{t=2} < 0. \tag{10}$$

This is strictly below zero, since all terms are negative.

Differentiating 6 with respect to p gives:

$$\frac{\partial CS_t}{\partial p} = -[L_{t=1} + (1-2p)L_{t=2}] > 0. \tag{11}$$

This is strictly positive, since $L_{t=1}$ and $L_{t=2}$ are both negative, yielding a negative term inside the brackets and a strictly positive overall term.

The term $\frac{\partial P_t}{\partial p}$ is strictly negative and $\frac{\partial CS_t}{\partial p}$ is strictly positive, resulting in opposite signs on the response to a change in the probability of a bad state, p.

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