# Capital Costs, Reinsurance, and the Price of Climate Risk

Benjamin L. Collier* Cameron M. Ellis† Po-Kang Hsiao‡ Anran Li§
Adam Solomon¶

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**Abstract**

In catastrophe-exposed insurance markets, a large share of what homeowners pay reflects the cost of correlated tail risk. Local insurers cannot diversify these losses on their own and instead transfer them to globally diversified reinsurers, but reinsurance is expensive and volatile. We study a regulatory reform that decreased the capital cost of reinsuring tail risk. After the cost decrease, reinsurance use expanded by 25%, the reinsurance market became less concentrated, and the price of reinsurance fell by a third. We find that these changes in the upstream reinsurance market flowed down to the homeowners insurance market in hurricane-exposed areas of Florida: additional insurers entered and homeowners' premiums for insuring wind risk fell sharply. By reducing required capital, the reform also increased insurers' exposure to reinsurer non-payment risk, but the magnitude of this effect is small overall because the reform led insurers to shift toward lower-risk reinsurers. To quantify welfare, we estimate an equilibrium model of the Florida wind-insurance market. The model implies that the reform reduced treated insurers' marginal cost of supplying wind coverage by about 10% and lowered equilibrium premiums by about 6%. Counterfactual simulations imply consumer-surplus gains of $95 per owner-occupied household per year, with about 60% attributable to lower marginal costs and the remaining 40% to expanded product availability. Both channels incorporate equilibrium markup adjustments. Our results imply that capital costs required to cover tail events, over and above average losses, meaningfully contribute to consumers' premiums and that policies that reduce those capital costs can decrease prices and improve availability for homeowners insurance.

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*University of Wisconsin-Madison, ben.collier@wisc.edu
†Tippie College of Business, University of Iowa, cameron-ellis@uiowa.edu
‡University of Wisconsin-Madison, phsiao4@wisc.edu
§University of Minnesota Twin Cities and NBER, anranli@umn.edu
¶NYU Stern, adam.solomon@nyu.edu

# 1 Introduction

The concentration and scale of natural catastrophes make them difficult to insure. In 2024 alone, 27 natural disasters in the U.S. each exceeded a billion dollars of damage (NOAA, 2026), often in a localized area. Insurers without national or international diversification do not have the capacity to pay billions of dollars of claims. Instead, insurers "purchase diversification" through the reinsurance market. Reinsurance – insurance for insurers – allows geographically concentrated insurers to offload tail risk to large, mostly foreign reinsurers that pool across countries and hazards. While reinsurance allows catastrophe risks to be insured by local insurers, it is a costly, volatile and opaque input. For example, reinsurance prices can more than triple year-on-year, and are the largest component of hurricane insurance premiums for homeowners in Florida, often four times larger than the expected loss (Boomhower et al., 2026). The capacity supplied by the reinsurance market, and the price at which it is supplied, are primary drivers of the ultimate cost homeowners pay to be protected against natural catastrophes.

Despite the central role of reinsurers in the U.S. property insurance market, we have remarkably little causal evidence on why reinsurance prices are so high, and on the extent to which they pass through to homeowners. Reinsurance premiums co-move with hurricane seasons, accumulated loss experience, and the balance sheets of global reinsurers, so simple regressions of price on any of these conflate supply shocks with demand shocks. Whether these price changes reflect updated risk assessments (Boomhower et al., 2026), shocks to reinsurer capital (Froot and O'Connell, 2008), or the exercise of market power (Froot, 2001) is hard to disentangle from observational data alone.

We provide such evidence by exploiting a policy reform that made foreign reinsurance less capital intensive. Reinsurers typically do not keep enough funds on hand to pay out 100% of potential claims, and instead rely on risk pooling. However, prior to 2009, insurance regulators across the U.S. required foreign reinsurers to post collateral to back their reinsurance agreements with U.S. insurers. Domestic reinsurers were never subject to this collateral requirement. From 2009 to 2021, U.S. states adopted NAIC Model Law 785, which reduced the collateral that foreign reinsurers were required to post by as much as 80% of the pre-reform level.[1] Posting collateral is expensive: it ties up funds that could otherwise be invested to earn returns elsewhere or used to pay claims on other accounts. Cutting collateral requirements, therefore, lowers the effective capital cost of supplying reinsurance to U.S. insurers, without changing the physical risks being

[1]While insurance markets are regulated at the state level, the National Association of Insurance Commissioners (NAIC) is the organizing entity that seeks to standardize insurance policy across states through model laws. States then choose whether (and when) to adopt these laws.

transferred. Adoption was staggered across states with primary insurers assigned to treatment through their domicile state, even when doing business in other states. As a result, we can compare reinsurance usage and, downstream, homeowner premiums and entry decisions across otherwise similar insurers who are exposed to the same risk but face different reinsurance capital costs.

We first examine how the collateral reform affected the U.S. reinsurance market. Using regulatory, contract-level data on all U.S. reinsurance contracts and Borusyak et al.'s (2024) imputed difference-in-differences design, we find that limiting collateral requirements increased foreign reinsurance use: U.S. primary insurers purchased 25% more post-reform. The effects are concentrated among insurers with substantial hurricane-corridor exposure, consistent with this group having the highest latent demand for catastrophe risk transfer. The reform also reduced concentration in the reinsurance market: state-level Herfindahl indices and (negative) Shannon entropy both declined, indicating that risk was reallocated across a broader set of reinsurers as the cost of accessing offshore capacity fell.

Next, we study the effect of the reform on reinsurance contracts and prices using a newly collected dataset. We extract reinsurance prices (premiums versus modeled expected losses) and contract structure (attachment points, coverage amounts) from free-text regulatory filings using text parsing and large language models. These data describe over 3,500 distinct reinsurance agreements with U.S. insurers from 2004 to 2025. We find that the capital cost reduction led to substantially lower reinsurance prices: the dollars of premium per dollar of modeled expected loss fell by 2.11 from a pre-reform mean of 6.85, a reduction of 31%. Moreover, four years after adoption, 27 percentage points more insurers bought a reinsurance contract covering extreme tail (1-in-500 year) events. Thus, collateral relief reduced the cost of reinsurance and made protection for tail risks economical to purchase when it was not before.

Collateral relief also carries a potential cost: a countervailing effect on credit risk. The post-reform schedule is graded by credit quality, with the largest collateral reductions going to the most highly rated foreign reinsurers. In principle, such a graded schedule could either lower system-wide credit risk—by steering cedents toward more highly rated counterparties—or raise it, by leaving a larger share of recoverables unsecured. We show that, on net, it raised it. Using implied impairment rates from rating agencies, we estimate that adoption increased expected impairment losses by approximately $57,600 per treated cedent-year. This increase does not reflect a shift toward weaker counterparties: the share of premium ceded to unrated reinsurers falls by 1.7 percentage points, and the premium-weighted impairment probability falls by 2.1 basis

points. Rather, the dollar measure reflects the combined effect of a larger ceded book and a larger unsecured share of reinsurance claims obligations.

We then study whether the reduction in reinsurer capital costs reaches consumers in the primary market. We focus on wind insurance premiums in Florida, the most catastrophe-exposed state in the country. Wind insurance covers homes against severe storms such as hurricanes. In Florida, households typically purchase wind coverage as part of their homeowners insurance policy, but its price is separately itemized rather than bundled with the other risks covered by the policy (e.g., theft). We use an imputed triple-difference approach to isolate the catastrophe-risk channel, comparing the price of wind-inclusive coverage to non-wind coverage, by the same insurer and in the same county, before versus after the reform. We find that the reform lowers the wind-inclusive premium rate by about \$0.50 per \$1,000 of insured exposure, roughly 10–16% of a typical pre-treatment rate. This premium effect is about twice as large in high-wind-risk counties as in low-risk counties, consistent with cheaper catastrophe reinsurance driving the result. Lower reinsurance costs also expand availability: treated insurers gain roughly one percentage point of county total-insured-value market share and become about five percentage points more likely to write wind-inclusive coverage relative to non-wind coverage. In summary, the reduced-form results establish that the reform changed reinsurance quantities, competition, primary-market premiums, and entry decisions.

To quantify welfare and study counterfactual policies, we build an equilibrium model of wind-coverage products in Florida. Consumers choose among homeowners insurance products with heterogeneous price sensitivity, while multi-product insurers set prices strategically. The supply side of the model links the collateral reform to the cost of providing wind coverage, allowing us to estimate how much cheaper reinsurance lowers primary insurers' marginal costs. We identify this cost channel using insurers' pre-reform share of foreign reinsurance purchases as a Bartik instrument and the staggered adoption of the collateral reform across states.

Model estimates imply that the collateral reform reduced the marginal cost of supplying wind coverage in Florida by approximately 10%. By encouraging entry, the reform reduced oligopolistic market power, magnifying the cost shock in pricing: the implied premium reduction is about 6%. Counterfactual simulations imply that the reform raises consumer surplus by about \$95 per owner-occupied household per year, or roughly \$470 million annually across Florida homeowners. These gains operate through two channels: lower marginal costs, accounting for about 60% of the total, and expanded product availability, accounting for about 40%. Both channels incorporate insurers' equilibrium pricing responses, including adjustments in markups. The cost-side benefits are greater in high-risk coastal counties where reinsurance constitutes the largest compo-

nent of premiums, whereas the variety gains are larger in thinner inland markets where additional products have the greatest value.

Our first contribution is to the literature on the costs of catastrophe risk in consumer insurance markets. Keys and Mulder (2024) document a rapid rise in U.S. homeowners' insurance premiums since 2017, especially in high disaster risk areas, and cite increasing reinsurance costs as a key explanation. Oh et al. (2026) study how state-level premium rate regulation can lead to cross-subsidization between states. Boomhower et al. (2024) illustrate how risk classification strategies vary among insurers and can generate a winner's curse that reduces insurance availability in high-risk segments. Sastry et al. (2024) show that climate losses destabilize insurer balance sheets with spillovers into mortgage markets. These papers take reinsurance costs as a fixed input to high insurance prices and study the broader implications of those prices for households. We complement them by isolating a clean, policy-driven shock to reinsurer capital costs themselves – the key mechanism that sits between disaster losses and consumer premiums – and measuring its pass-through to homeowner prices and its spatial incidence across risk levels.

Our second contribution is to the literature on market power, capacity, and entry in (re)insurance. Gron (1994), Winter (1994), and Froot (2001) argue that slow-moving capital and reinsurer market power generate high prices and restricted supply, and Koijen and Yogo (2016) show that regulation shapes how insurance liabilities are ultimately placed across captives and affiliated entities. In concurrent work, Anand et al. (2026) contrast insurers' internal capital management with regulatory capital requirements, using the collateral reform as a natural experiment. We show that reinsurance collateral rules operate as barriers to entry that depress competition and misallocate risk.[2] When these barriers are lowered, prices decline and competition increases as risk moves to foreign reinsurers that can more efficiently bear it. This, in turn, expands the set of homes that are economical for primary insurers to cover.

Our third contribution is to the finance literature on how the cost of capital at financial intermediaries feeds through to the prices that end-users pay. Froot and Stein (1998) and Froot (2007) show theoretically that costly external capital at financial institutions—and especially at insurers and reinsurers—should be priced into the contracts they sell; Zanjani (2002) models how marginal capital requirements for insuring tail risk push catastrophe insurance prices above expected losses; and Koijen and Yogo (2015) estimate the shadow cost of statutory capital for U.S. life insurers during the financial crisis and tie it to observed markups. Reinsurers are capital-intensive financial intermediaries whose required return on capital shapes the price of the risk transfer they sell, and collateral requirements are effectively a tax on that capital. By tracing a quasi-experimental reduc-

[2] As predicted, pre-reform, by Cole et al. (2010).

tion in that capital tax from reinsurer cost functions to homeowners insurers' supply decisions to consumer premiums, we provide a direct estimate of the cost-of-capital channel in property insurance pricing. The magnitudes matter well beyond this particular reform: they say how much of any such cost reduction – whether delivered by alternative-capital providers (Cummins and Weiss, 2009; Bloomberg News, 2026), new modeling technology, or new entry – reaches consumers rather than being retained as rents upstream.

A direct policy implication is that the case for a government reinsurer of last resort for property catastrophe risk is stronger than the standard market-failure argument suggests. The usual case, going back to Lewis and Murdock (1996) and summarized by Kousky and Cooke (2012), rests on private reinsurers' inability to diversify correlated catastrophe losses and on capacity constraints in extreme events, while Cummins (2006) cautions that government provision can crowd out private and alternative-capital solutions. Solomon (2026) shows that public reinsurance provided at cost in Australia does substantially reduce prices and increase availability. Our estimates add a separate channel, rooted in Arrow and Lind (1970) and formalized recently by Jia et al. (2025): a government reinsurer would fund its risk-bearing at a lower cost of capital than private reinsurers face, and our pass-through numbers imply that a large share of that cost-of-capital difference would reach homeowners as lower premiums rather than being absorbed as rents upstream. This channel operates even in quiet years when private capacity is not binding, and it is concentrated in exactly the high-risk counties where property insurance is becoming unaffordable.

# 2 The Collateral Reform: Institutional Details and Data

## 2.1 Institutional Details

### 2.1.1 Reinsurance

Reinsurance is insurance for insurers. It transfers risk between primary insurers ("cedents") and reinsurers: the cedent pays a premium and the reinsurer pays a contractually-specified share of incurred losses. Insurers purchase reinsurance for three primary reasons: capital relief, earnings smoothing, and access to capacity for tail risk that is too correlated to retain on their own balance sheet. For homeowners insurance exposed to natural catastrophes, the third is most important.

The United States is the largest single source of ceded premium in the global reinsurance market, and most of that premium is assumed by reinsurers domiciled outside the United States. In 2014, over 62% of premium ceded by U.S. insurers went to foreign-domiciled reinsurers (Federal

Insurance Office, U.S. Department of the Treasury, 2014). Reinsurance supply is concentrated in a small set of large, mostly non-U.S. firms: Munich Re, Swiss Re, Hannover Re, SCOR, the Lloyd's market, Berkshire/Gen Re, and Bermuda specialists such as RenaissanceRe, Everest, Arch, and Axis. Bermuda alone is the largest single supplier of catastrophe reinsurance to U.S. insurers (Cummins, 2008).

The offshore concentration of catastrophe reinsurance supply is driven primarily by tax and regulatory incentives. First, prior to the 2017 Tax Cuts and Jobs Act, the U.S. combined federal and state corporate tax rate was approximately 39%. In contrast, Bermuda had no corporate income tax and European rates ranged from 12.5% (Ireland) to 38% (France) (Organisation for Economic Co-operation and Development, 2024), generating a substantial after-tax return differential on the equity capital reinsurers must hold against tail losses. Second, U.S. risk-based capital is calibrated primarily to traditional liability lines and gives limited credit for diversification across catastrophe zones, while Bermuda's Solvency Capital Requirement and the European Solvency II regime credit cross-zone diversification more generously through internal catastrophe models (CAS Risk-Based Capital Dependencies and Calibration Working Party, 2012).

Since catastrophe risk-bearing capital is most efficiently held internationally, the relevant question for U.S. insurers is the cost of transferring risk offshore. Prior to the reform we study, foreign reinsurers were required to post 100% collateral against ceded liabilities. That collateral ties up capital that could otherwise earn risk-adjusted returns elsewhere. This is the policy lever changed by NAIC Model Law 785.

### 2.1.2 Reinsurance Regulation and Reform

Insurance in the United States is regulated at the state level. State insurance departments oversee prices and policy forms, market conduct, and the capital adequacy of insurers domiciled in the state. Reinsurance is regulated as part of this framework, primarily through its effect on the ceding insurer's required capital. Whether a reinsurance contract reduces a cedent's required capital—the "credit for reinsurance"—is determined by the cedent's state insurance code, which in nearly all states tracks the NAIC's Credit for Reinsurance Model Law.

Authorized reinsurers—those licensed in the cedent's state—are subject to the same solvency regulation as primary insurers, and a primary insurer that cedes to an authorized reinsurer automatically gets credit against required capital for all the catastrophe risk transferred. Unauthorized reinsurers, in particular, foreign-domiciled reinsurers, fall outside any state regulator's direct authority. For decades, the credit-for-reinsurance rule required that reinsurance ceded to a foreign

reinsurer be backed by collateral equal to 100% of the ceded liabilities, held in a U.S. trust in eligible high-quality, liquid, but low-return assets (Federal Insurance Office, U.S. Department of the Treasury, 2014). This functions as a cost wedge between reinsurance ceded to domestic and to foreign reinsurers: the latter must hold costly collateral, increasing the price for U.S. insurers to transfer risk offshore.

**Policy Reform.** We study a reform to this collateral requirement for foreign reinsurers. Starting in Florida in 2009, and proceeding across other states over the next decade[3], the collateral required to be posted by foreign reinsurers was reduced. Instead of 100%, foreign reinsurers now had to post collateral ranging from 0% (AAA) to 50% (A-) to 100% (BB and below) depending on their credit rating (National Association of Insurance Commissioners, 2019). This reduction in the cost of reinsurance benefited all U.S. insurers, but particularly those writing policies in catastrophe-exposed regions such as Florida, where the correlation of risk makes reinsurance critical. Once the implementing state law was passed, foreign reinsurers had to get certified to be eligible to post reduced collateral. That certification process often took one to three years, meaning the impacts of the law were often not instant.

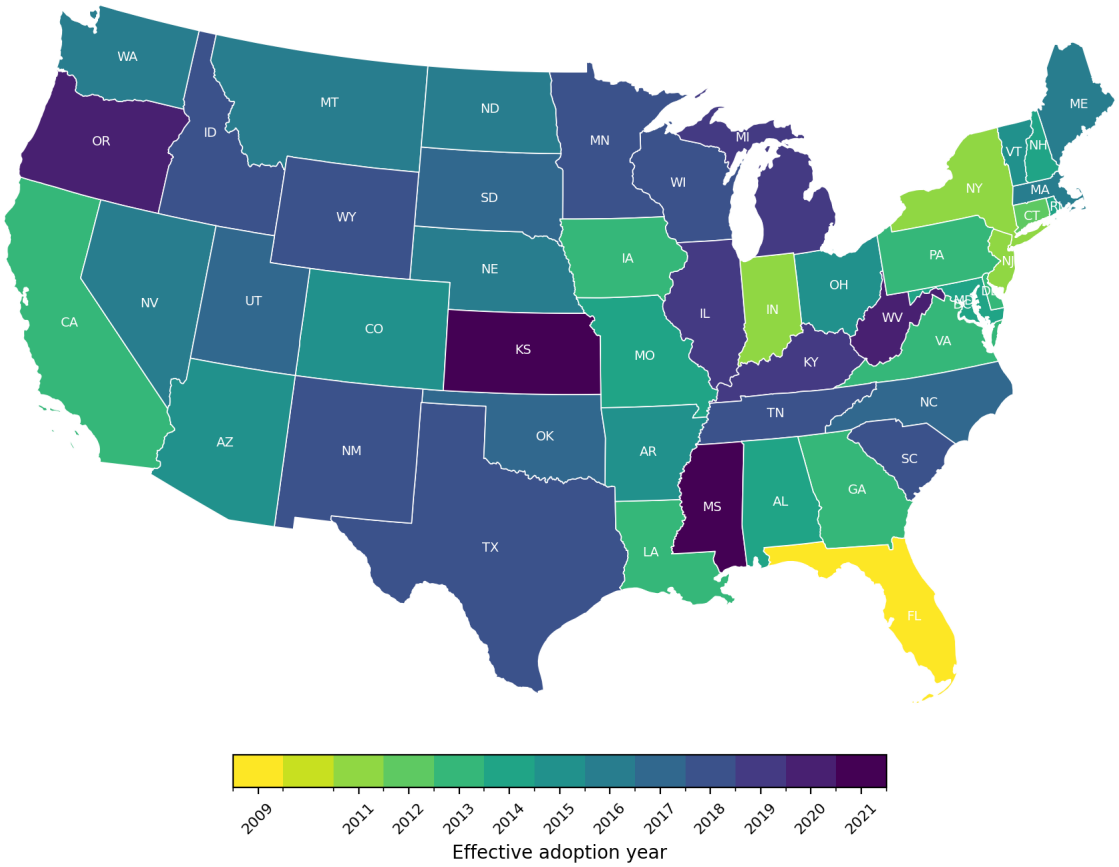
Treatment is assigned by domicile state—where the insurer is incorporated and subject to primary oversight. When a domicile state adopts Model Law 785, the insurer gains access to reduced-collateral reinsurance regardless of where it writes policies. For our Florida analysis, this creates within-market variation: all sample insurers write Florida policies, but they are domiciled in different states with different adoption dates.

Collateral is costly, but it also secures a cedent's claim if a reinsurer becomes impaired. The reform therefore trades lower capital costs against a larger unsecured exposure to reinsurer solvency. Because the post-reform collateral schedule depends on financial-strength ratings, it may also affect which counterparties cedents select. We examine this counterparty-risk margin in Section 3.2.5.

**Implications of the Collateral Rules.** By trapping funds in low-yielding, U.S. trust accounts, the pre-reform collateral rules imposed constraints on insurers' balance sheets and cash holdings that increased operating costs. The size of collateral holdings prior to the reform was large: collateral for unauthorized foreign reinsurers equaled approximately 87% of ceded premium (Appendix B).

[3] A full timeline of adoption is provided in Appendix D.

Figure 1: Policy Reform Timing




*Note:* Effective date of state adoption of NAIC Model Law 785, which replaced the 100% collateral requirement on foreign reinsurers with a sliding scale (0–100%) tied to credit ratings. Source: hand-collected state administrative codes and statutes; see Appendix Table 9.

The collateral trusts have three main implications for insurers' profitability. First, the reinsurer forgoes the yield it would earn on an intermediate-duration, investment-grade portfolio, holding instead the restricted, short-duration Treasuries admitted as collateral.[4] Applied to collateral of $0.87 per dollar of ceded premium, this implies a direct opportunity cost on the order of 2 to 3 cents per year, which the reform reduced to near zero. Second, the collateral rules can trap reinsurers' funds for long durations. Reinsurers immediately increase their reserves when a catastrophe (e.g., a major hurricane or wildfire) occurs in anticipation of future claims, and the collateral rules re-

[4]P&C insurers hold approximately two-thirds of their portfolios in high-grade bonds and earned industry book yields in the low-to-mid 3 percent range through the 2010s (S&P Global Market Intelligence; NEAM). Restricted trusts must hold ultra-liquid short Treasuries, which yielded well under 1 percent for much of that period, so the forgone spread is in the 2 to 3 percent range. Reinsurers satisfying the requirement via letters of credit avoid ring-fencing their own assets but pay annual fees of comparable magnitude.

quire immediate funding of trust accounts to correspond with this reserving. However, reinsurers' obligations are only settled after insurers have paid their claims, which sometimes takes several years to resolve, extending the time that reinsurers' funds are trapped in low-yield investments. Third, the collateral requirements can exacerbate currency and asset-liability mismatches on globally diversified books. They trap liquidity geographically: assets pledged against U.S. liabilities cannot meet a simultaneous shock elsewhere, so the reinsurer must hold additional unrestricted buffers. Removing these constraints lowered reinsurers' funding and frictional costs, expanding supply and reducing the equilibrium price of reinsurance. This is consistent with the declines in the Florida wind premium we analyze in Section 4.2.1.

### 2.1.3 Florida Residential Property Insurance

We study the effects of reinsurance cost changes on home insurance in Florida. Florida is the most catastrophe-exposed residential insurance market in the United States. The state has roughly $2.9 trillion of insured property value in its coastal counties that regularly experience hurricane landfalls (AIR Worldwide, 2013). Average annual insured catastrophe losses in Florida exceed those of any other state by a wide margin.

The structure of the Florida market compounds the effects of this concentration. National insurance groups, particularly State Farm and Allstate, have in recent decades sharply curtailed their Florida residential business (Carrillo et al., 2022). The void was filled by small, Florida-domiciled "specialist" insurers (Grace et al., 2005), which today write roughly 54% of Florida residential premium (Florida Office of Insurance Regulation, 2024). These domestic specialists are undiversified and thinly capitalized relative to their concentrated tail exposure.

As a result, Florida is particularly reliant on reinsurance. Florida insurers cede roughly 40% of direct premium to reinsurers, compared with a national average of about 16% (Gallagher Re, 2026). While a narrow portion of reinsurance coverage is provided by the public Florida Hurricane Catastrophe Fund (FHCF), insurers still rely on private reinsurance to offload their highly concentrated exposure to hurricanes. Cost shocks to reinsurance consequently feed directly into the prices consumers face in the primary homeowners insurance market in Florida.

## 2.2 Data Sources

We combine four data sources: (1) reinsurance transaction data from NAIC Schedule F filings, (2) hand-collected and validated state-level adoption dates for NAIC Model Law 785, (3) quarterly

data on Florida homeowners insurance from the Quarterly Supplemental Report (QUASR) of the Florida Office of Insurance Regulation, and (4) reinsurance layer prices and modeled expected losses from Florida homeowners rate filings.

### 2.2.1 Reinsurance Data

Our reinsurance data come from NAIC statutory annual statements and cover the period 2004 through 2024. Schedule F reports the premiums each insurer cedes to each of its reinsurers. In Table 1, we report summary statistics for these reinsurance flows, aggregated to the insurer-year across all lines of business.

Table 1: Summary Statistics for Reinsurance Data (Balanced Panel)

| | | Pre-treatment | | | | Post-treatment | | | |
|---|---|---|---|---|---|---|---|---|---|
| Outcome | Unit | Mean | SD | P25 | P75 | Mean | SD | P25 | P75 |
| Direct premiums written (DPW) | $M | 5,460.96 | 9,350.58 | 453.98 | 5,085.46 | 4,900.16 | 10,517.07 | 548.52 | 4,286.85 |
| Premium ceded to reinsurance | $M | 1,645.74 | 2,843.00 | 181.02 | 2,064.80 | 2,482.37 | 4,137.98 | 266.72 | 2,732.35 |
| Number of reinsurers | | 56.50 | 91.29 | 3.00 | 67.00 | 62.92 | 103.80 | 2.00 | 79.00 |

*Note:*
All quantities cover all lines of business. Premium ceded is an annual flow of ceded premium written. Dollar amounts are in $M. Statistics are DPW-weighted within each period. Pre = event time $< 0$; Post = event time $\geq 0$.

Premium ceded to reinsurance rises by about half after treatment (from $1,646M to $2,482M on average), and the number of reinsurers rises from about 57 to 63. These changes in the summary statistics – a higher quantity of reinsurance and more reinsurance competition – are investigated causally in subsequent sections.

Finally, Schedule F of the NAIC filings records the reinsurance premium ceded separately for each insurer-reinsurer pair. Schedule F notes whether the reinsurer is foreign, affiliated, etc. We use these data to analyze the quantity of reinsurance ceded to foreign reinsurers; the gross reinsurance in/reinsurance out flows from the income statement do not allow for that distinction.

### 2.2.2 Reinsurance Prices from Florida Rate Filings

NAIC Schedule F filings contain the universe of reinsurance premium ceded but no information on the expected losses those premiums are paying for protection against. We gather data on those expected losses, and hence on the economic price of reinsurance, from the raw regulatory filings

that insurers filed with the Florida Office of Insurance Regulation. Florida requires homeowners insurers to document the cost of their reinsurance program in each rate filing. For each layer of the catastrophe reinsurance tower, the filing reports the limit, attachment, ceded premium, and the modeled expected loss to the layer. We collect these exhibits for the 109 insurers in QUASR over 2004 through 2025 that we can match, giving 3,576 priced private excess-of-loss layers. Summary statistics are reported in Table 2.

Table 2: Summary Statistics (Prices) for Reinsurance Data

| | Pre-Treatment | | Post-Treatment | | |
|---|---|---|---|---|---|
| Implied occurrence period | Reinsurance Price / Expected Loss Multiple | N of Reins. Layers | Reinsurance Price / Expected Loss Multiple | N of Reins. Layers | Diff. |
| $<$ 1-in-10 | 2.99 | 237 | 2.30 | 719 | $-0.69$ |
| 1-in-10 to 25 | 3.81 | 208 | 3.09 | 367 | $-0.72$ |
| 1-in-25 to 50 | 6.16 | 219 | 4.64 | 368 | $-1.52$ |
| 1-in-50 to 100 | 7.87 | 225 | 6.24 | 477 | $-1.63$ |
| 1-in-100 to 250 | 10.41 | 237 | 9.60 | 302 | $-0.81$ |
| 1-in-250 to 500 | 15.34 | 52 | 16.31 | 103 | $+0.97$ |
| 1-in-500+ | 17.50 | 17 | 16.42 | 45 | $-1.08$ |
| All | 6.85 | 1,195 | 5.37 | 2,381 | $-1.48$ |

*Note:* Sample is priced private excess-of-loss layers in Florida homeowners rate filings, 2004–2025. Multiple is ceded premium divided by modelled expected loss. Implied occurrence period describes the modeled annual loss exceedance probability of the insurer's portfolio. Pre and Post are relative to Model Law 785 adoption in the carrier's domicile state; the adoption year counts as Post. Layers with a multiple of 50 or more are excluded (73 layers dropped). N counts layer rows; sibling filings by group members are included.

Our primary price measure is the ratio of premium to expected loss. We call this the reinsurance price multiple. We record this separately for each layer in the reinsurance "tower". For each layer in the tower, we observe the total possible loss and the expected loss. By dividing these, we approximate the probability of loss. From that, we compute an implied occurrence period: a layer that has a 10% probability of exhaustion is expected to occur once in every 10 years. We categorize reinsurance layers according to these calculated occurrence periods.

Table 2 shows that the average pre-reform price multiple was approximately 6.85: for every dollar of expected loss reinsurance was purchased against, the insurer paid \$6.85 in premium. After the reform, the price multiple fell to 5.37. This reduction is evident at almost every implied occurrence period, with both the pre-treatment price and the reduction generally larger for reinsurance further into the tail.

### 2.2.3 Florida Homeowners Insurance Data

The Quarterly Supplemental Report (QUASR) from the Florida Office of Insurance Regulation covers all 67 Florida counties from 2010 Q1 through 2020 Q4. We aggregate to the insurer-county-quarter level. The data include policy counts (total policies in force, with wind/non-wind breakdowns), premiums (total and by wind coverage status), exposures (total insured value), and policy flows (new policies, cancellations, nonrenewals, transfers).

*Sample Restrictions.* We restrict to standard personal residential homeowners policies (PR-HO). All "wind-only" standalone products (typically written by Citizens or the Florida Hurricane Catastrophe Fund) are dropped at the aggregation stage so that the unit of observation is a single bundled homeowners contract. Within the retained rows, premiums are decomposed into wind and non-wind components based on whether the policy itself includes wind coverage; this is the split reported in the "Wind share of DPW" row of Table 3.

We correct a small set of insurer-quarters where per-policy premiums appear mis-scaled by a factor of 100 and drop observations with non-finite log premium or log rate per \$1,000 of exposure. We then restrict to insurers whose domicile state has a known reinsurance collateral policy change adoption date within the QUASR years of observation (2010 through 2020). In particular, this removes Florida-domiciled insurers, who were treated in 2009. Since they are the most reliant on reinsurance, our results are likely conservative. Summary statistics for the Florida QUASR data are in Table 3 below.

As their states of domicile are treated, more insurers begin writing business in Florida. Those that do so write more policies, charge lower premiums, and move toward wind-heavy policies. This is consistent with the effects of a reinsurance price decrease passed through to lower homeowners insurance prices specifically in hurricane-exposed policies, and the offering of new policies that were previously uneconomical.

## 3 Effects on Upstream Reinsurance Access, Use, and Cost

This section measures the effects of the collateral reform on the national reinsurance market. We examine, in turn, the quantity of reinsurance ceded to foreign reinsurers, the concentration of the reinsurance market, reinsurance prices, purchases of tail-risk reinsurance, and cedents' exposure to counterparty risk.

Table 3: Florida Homeowners Insurance Summary Statistics (DPW-weighted): Pre vs. Post Treatment

| | | Pre-treatment | | | | Post-treatment | | | |
|---|---|---|---|---|---|---|---|---|---|
| Outcome | Unit | Mean | SD | P25 | P75 | Mean | SD | P25 | P75 |
| Policies in force | policies | 2,541.65 | 2,321.34 | 805.00 | 3,469.00 | 16,737.07 | 27,488.29 | 1,590.00 | 16,781.00 |
| Total direct premium written (DPW) | $M | 9.09 | 11.20 | 2.00 | 11.53 | 52.05 | 102.44 | 3.59 | 41.67 |
| Avg. premium per policy | $ | 4,129.77 | 4,294.15 | 1,634.12 | 4,084.28 | 2,940.88 | 2,508.55 | 1,669.39 | 3,200.48 |
| Total insured exposure | $M | 2,505.24 | 3,468.73 | 537.32 | 2,765.58 | 6,776.41 | 9,882.66 | 841.69 | 7,665.95 |
| Rate per $1,000 of exposure | $ | 6.81 | 83.16 | 3.06 | 4.79 | 6.30 | 169.57 | 3.56 | 6.98 |
| Wind share of DPW | | 0.91 | 0.16 | 0.90 | 1.00 | 0.97 | 0.10 | 0.98 | 1.00 |
| New policies written | policies | 34.46 | 125.54 | 2.00 | 31.00 | 1,293.68 | 2,739.50 | 27.00 | 809.00 |
| Net new policies written | policies | -41.87 | 163.72 | -69.00 | -4.00 | 81.18 | 3,238.13 | -130.00 | 201.00 |
| Insurers per county-quarter | insurers | 27.25 | 11.50 | 18.00 | 39.00 | 55.99 | 13.24 | 50.00 | 66.00 |

*Note:* Observations are at the insurer-county-quarter level except "Insurers per county-quarter" (county-quarter level). Pre vs. post defined by whether the insurer's domicile-state reinsurance collateral policy change adoption date has been reached. Average premium per policy is winsorized at the 1st and 99th percentiles to remove data-coding outliers (e.g., misreported exposure leading to implausible rates). Net new policies written = new policies written − policies canceled − policies nonrenewed − policies transferred out + policies received from other insurers. Statistics are DPW-weighted within each period.

## 3.1 Empirical Strategy

### 3.1.1 Two-Stage Imputation Difference-in-Differences

The natural reduced-form specification for the pass-through of the collateral reform onto reinsurance outcomes is the two-way fixed-effects (TWFE) event study:

$$Y_{it} = \alpha_i + \lambda_t + \sum_{\tau \neq -1} \gamma_\tau \, \mathbf{1}\{\mathrm{RelTime}_{it} = \tau\} + \varepsilon_{it}, \tag{1}$$

where $Y_{it}$ is the outcome of interest for insurer $i$ in year $t$, $\alpha_i$ is an insurer fixed effect, $\lambda_t$ is a calendar-year fixed effect, and $\mathrm{RelTime}_{it} = t - t^{\text{adopt}}_{s(i)}$ is years since the insurer's domicile state $s(i)$ adopted the reinsurance collateral policy change. The reference period is $\tau = -1$. Under parallel trends, no anticipation, and homogeneous treatment effects, $\gamma_\tau$ identifies the average treatment effect on the treated (ATT) at event time $\tau$.

Applied directly, equation (1) is unreliable in our setting. The literature on staggered-adoption designs has shown that when treatment timing varies across units and treatment effects are het-

erogeneous over event time—both of which apply here—earlier-treated insurers enter as controls in comparisons of later-treated insurers, and the estimated $\{\gamma_\tau\}$ become a contaminated weighted average of cohort-specific effects, sometimes with negative weights (de Chaisemartin and D'Haultfœuille, 2020; Goodman-Bacon, 2021; Borusyak et al., 2024). Pre-trends estimated from the same regression inherit the same contamination, making the test for parallel trends unreliable.

To avoid these issues, we use a two-stage imputation estimator following Gardner (2022) and Borusyak et al. (2024). In the first stage, we estimate

$$Y_{it} = \alpha_i + \lambda_t + \varepsilon_{it} \tag{2}$$

using only observations with $D_{it} = 0$ (years before insurer $i$'s domicile state adopted the reinsurance collateral policy change). Insurer fixed effects absorb time-invariant differences across insurers; year fixed effects absorb common shocks to the reinsurance market, including hurricane seasons, accumulated loss experience, and aggregate movements in global reinsurer capital.

In the second stage, we use these estimates to predict counterfactual outcomes $\hat{Y}_{it}(0)$ for all observations, including treated insurer-years. The residual $Y_{it} - \hat{Y}_{it}(0)$ is the treatment effect for treated observations and noise for controls. Regressing these residuals on the treatment indicator,

$$Y_{it} - \hat{Y}_{it}(0) = \tau \cdot D_{it} + \eta_{it}, \tag{3}$$

yields $\tau$, which consistently estimates the ATT under parallel trends. Standard errors are clustered by state of domicile.

### 3.1.2 Insurer-Level Slope Difference-in-Differences

While most outcomes we analyze are levels, several reinsurance outcomes in Section 3 are most naturally interpreted as ratios—for example, the dollars of reinsurance ceded relative to direct premiums written, or claims recovered relative to premiums ceded. We adapt the methodology of Section 3.1.1 to estimate how these ratios change with treatment.

The naive TWFE analog is

$$Y_{it} = \beta_0 X_{it} + \beta_1 (X_{it} \times D_{it}) + \alpha_i + \lambda_t + \varepsilon_{it}, \tag{4}$$

where $Y_{it}$ is the numerator (e.g., reinsurance ceded), $X_{it}$ is the denominator (e.g., direct premiums written, DPW), $D_{it}$ is the treatment indicator, and $\beta_1$ captures the post-treatment change in the slope. The event-study analog replaces $D_{it}$ with $\sum_\tau \mathbf{1}\{\mathrm{RelTime}_{it} = \tau\}$. This specification suffers from the same staggered-adoption pathologies as equation (1).

We therefore implement the imputation analog. In the first stage, using only untreated observations ($D_{it} = 0$), we estimate

$$Y_{it} = \beta_0 X_{it} + \alpha_i + \lambda_t + \varepsilon_{it}. \tag{5}$$

In the second stage, we construct counterfactuals $\hat{Y}_{it}(0)$ and regress residuals on event-time-interacted denominators:

$$Y_{it} - \hat{Y}_{it}(0) = \sum_{\tau=-L}^{R} \gamma_\tau \cdot X_{it} \cdot \mathbf{1}\{\mathrm{RelTime}_{it} = \tau\} + \nu_{it}, \tag{6}$$

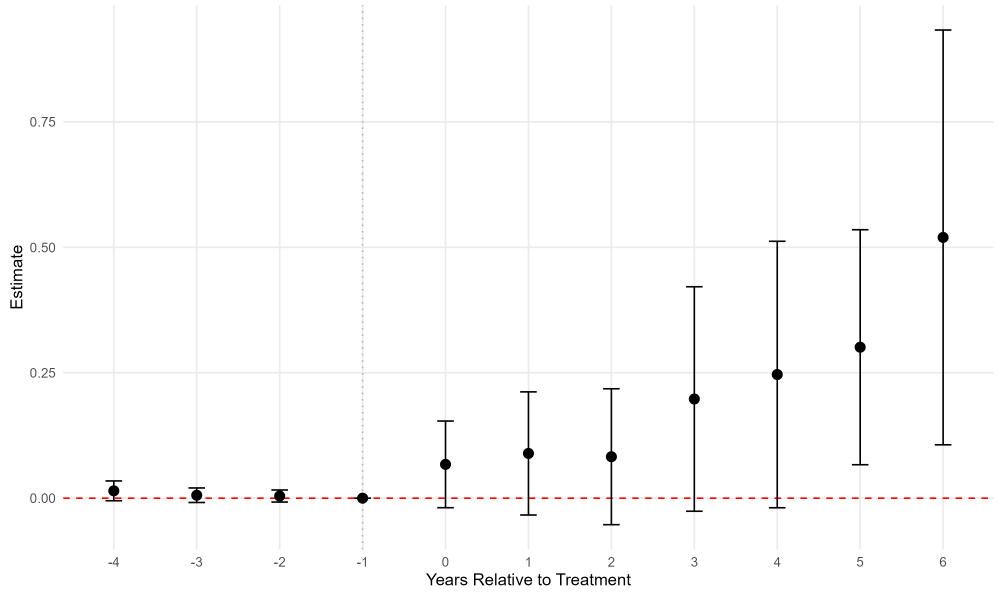
where $\mathrm{RelTime}_{it} = t - t^{\text{adopt}}_{s(i)}$ and $\tau = -1$ is the reference period. The slope coefficients $\{\gamma_\tau\}$ capture how the marginal relationship between $Y$ and $X$ shifts in event time. In both stages we weight by insurer-year direct premiums written. Standard errors are clustered by state of domicile.

## 3.2 Results

### 3.2.1 Reinsurance Quantity

We first analyze how the regulatory changes affected ceding behavior to foreign reinsurers—the category most directly affected by collateral and credit requirements under the reinsurance collateral policy change. Figure 2 displays the impact on reinsurance ceded to foreign reinsurers as a function of insurer size.

Figure 2: Quantity of Reinsurance Ceded: Foreign Reinsurers



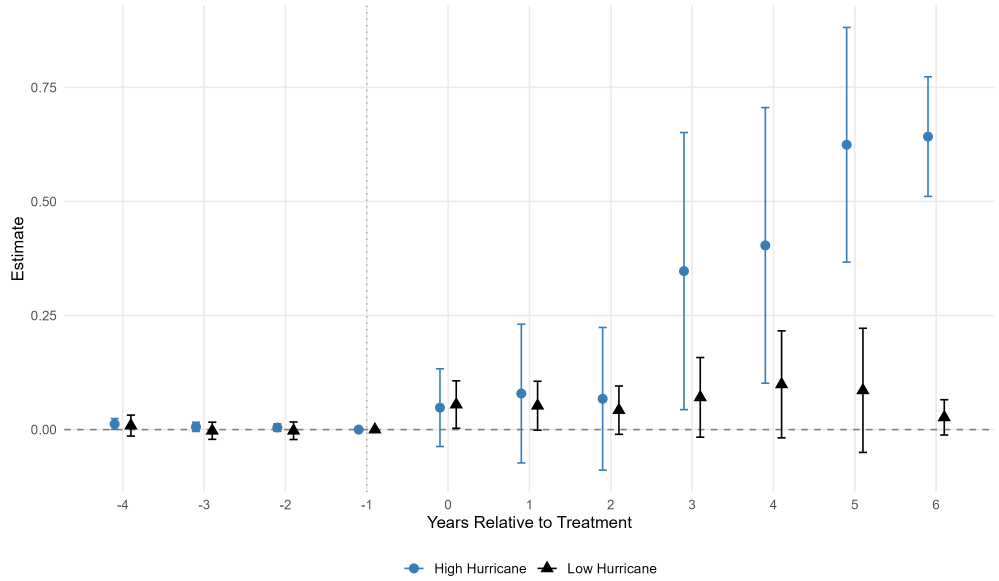

*Notes:* The figure reports an event study where the outcome is dollars of reinsurance ceded to foreign reinsurers; the exposure variable is direct premiums written (DPW). Event time $\tau$ is relative to the insurer's domicile state adoption. Coefficients $\{\gamma_\tau\}$ are from the two-stage specification in equations (5) and (6), with reference period $\tau = -1$. The event window is $\tau \in [-5, 6]$ with observations beyond $\tau = 5$ binned to $\tau = 6$. Estimation uses insurer DPW weights and clusters standard errors by state of domicile. State-of-domicile and year fixed effects are included in Stage 1.

The results show that treated insurers substantially increased reinsurance ceding to foreign reinsurers following the regulatory reform. Since the cost of doing so was directly reduced by the reform, this is intuitive.

To understand where the latent demand for reinsurance was highest, we split the sample of insurers into those that were above or below median in the share of their business, prior to treatment, that was written in hurricane-exposed states.[5] The results are in Figure 3.

[5]The states are: Florida, Texas, Louisiana, Mississippi, Alabama, South Carolina, North Carolina, and Georgia.

Figure 3: Quantity of Reinsurance Ceded: High vs. Low Hurricane Corridor Exposure




*Notes:* The figure reports a slope event study where the outcome is dollars of reinsurance ceded to foreign reinsurers; the exposure variable is direct premiums written (DPW). Insurers are split into high and low hurricane exposure groups based on their pre-treatment DPW-weighted average share of premiums written in hurricane corridor states (FL, TX, LA, MS, AL, SC, NC, GA), with the split at the DPW-weighted median. The high-hurricane-exposure estimates are in blue, the low-hurricane-exposure estimates are in black.

We see that the increase in reinsurance was driven entirely by insurers with high pre-treatment exposure to hurricane corridor states. This shows that demand for reinsurance is concentrated where losses are correlated. Consequently, reductions to the cost of offloading that tail risk accrue mostly to insurers who operate in hurricane-exposed areas.

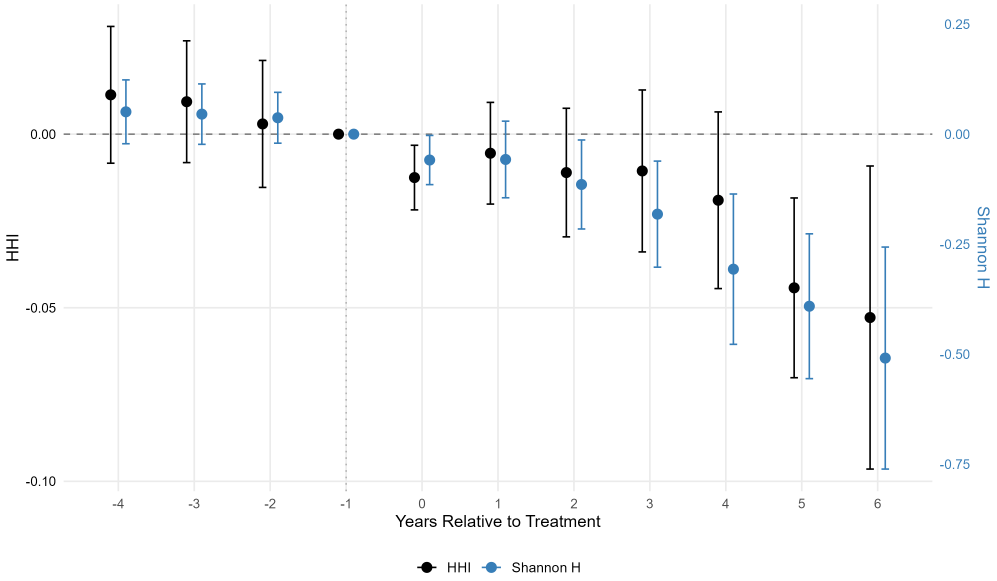
### 3.2.2 Market Concentration

To assess whether the regulatory reform affected market structure, we examine reinsurer concentration at the state-of-domicile level. Unlike the insurer-level analyses in Sections 3.2.1 and 3.2.3, which examine individual insurer ceding behavior, here we aggregate across all insurers domiciled in a given state to measure overall market concentration. This aggregation captures whether the reform changed the competitiveness of the reinsurance market accessible to insurers in treated states, as opposed to individual portfolio diversification patterns.

We measure concentration using two standard indices. The Herfindahl-Hirschman Index (HHI) is computed as $\sum_j s_j^2$ where $s_j$ is reinsurer $j$'s share of total ceded premiums in a state-year; higher values indicate greater concentration. Shannon entropy, computed as $H = -\sum_j s_j \log(s_j)$, naturally increases with market dispersion. For consistency of interpretation with HHI, we analyze the negative of Shannon entropy ($-H$) so that higher values indicate greater concentration and lower values indicate more competitive markets for both measures.

These outcomes are estimated with the two-stage framework in Section 3.1.1 (equations (2) and (3)), adapted to the state-year level with state-of-domicile and year fixed effects. Figure 4 shows that reinsurer concentration declined following adoption of the reinsurance collateral policy change. Both measures indicate that reinsurance markets became less concentrated—equivalently, more competitive—in treated states post-reform.

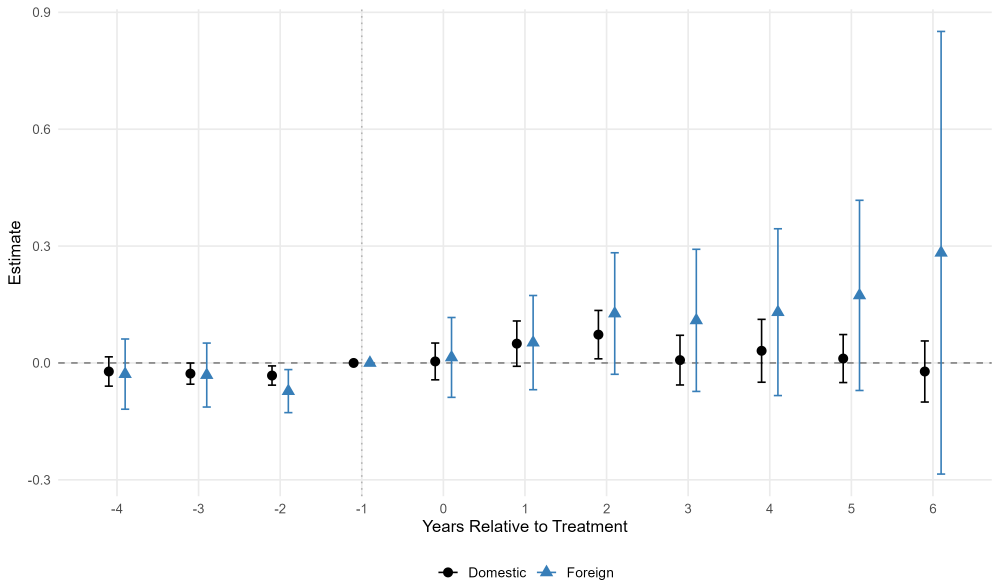
Figure 4: Reinsurance Market Concentration: HHI and Shannon Entropy




*Notes:* This figure reports a state-level event study where the outcomes are two measures of market concentration. For each state-year, we aggregate ceded premiums across all insurers domiciled in that state and compute concentration measures from the resulting distribution of premiums across reinsurers. Lower HHI (left axis) and (negative) Shannon entropy (right axis) indicate less concentrated (more competitive) markets. The sample includes only state-years constructed from insurers in the balanced panel (observed continuously from 4 years before through 5 years after their state's adoption). Event time $\tau$ is relative to state adoption. Coefficients $\{\beta_\tau\}$ are from the two-stage specification described in Section 3.1.1, adapted to the state-year context with state and year fixed effects in Stage 1 (equations (2) and (3)), with reference period $\tau = -1$. The event window is $\tau \in [-4, 6]$ with observations beyond $\tau = 5$ binned to $\tau = 6$. Estimation uses state-level aggregate DPW weights and clusters standard errors by state of domicile.

The decrease in market concentration was largely due to insurers contracting with foreign reinsurers they did not contract with before. We illustrate this in Figure 5 below. The outcome is the log of the number of reinsurers an insurer contracts with. Five years after treatment, the typical insurer contracts with 50% more foreign reinsurers, but no additional domestic reinsurers, consistent with the intended impact of the policy change.

Figure 5: Number of Reinsurer Relationships: Domestic vs. Foreign (Log Counts)




*Notes:* This figure reports an event study where the outcome is the log number of distinct reinsurers an insurer cedes to in a given year, split by counterparty type: domestic (NAIC-licensed) reinsurers and foreign (non-NAIC) reinsurers. The sample includes only insurers in the balanced panel (observed continuously from 4 years before through 5 years after their state's adoption). Event time $\tau$ is relative to state-of-domicile adoption of the reinsurance collateral policy change. Coefficients $\{\beta_\tau\}$ are from the two-stage specification described in Section 3.1.1 (equations (2) and (3)), with state-of-domicile and year fixed effects in Stage 1 and reference period $\tau = -1$. The event window is $\tau \in [-4, 6]$ with observations beyond $\tau = 5$ binned to $\tau = 6$. Estimation uses insurer DPW weights and clusters standard errors by state of domicile.

### 3.2.3 Reinsurance Prices

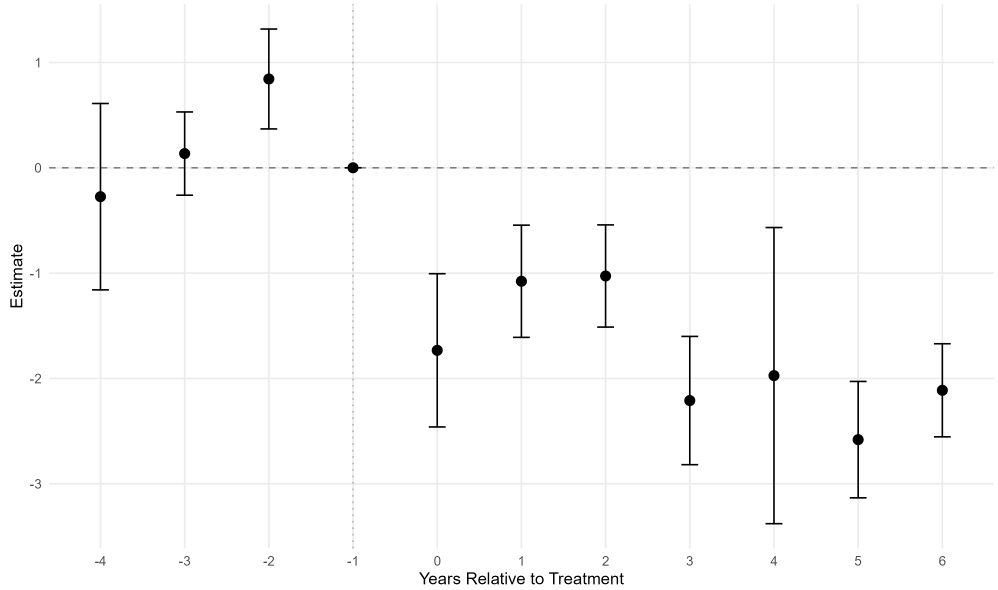
The Florida rate filings described in Section 2.2.2 let us measure the price of reinsurance directly. For each layer of a carrier's catastrophe tower, the filing reports the premium ceded and the modeled expected loss to the layer. We define the ratio of the two as our primary measure of price:

the premium paid per dollar of expected loss transferred. We call this ratio the "price multiple." Table 2 shows that the multiple rises steeply with the implied occurrence period of the layer.

We run the same event study specified in equations (2)–(3). The outcome is the carrier-year multiple, ceded premium over expected loss summed across the carrier's priced layers. A negative coefficient means cheaper reinsurance.

Figure 6 reports the estimates. The multiple falls by 1.73 in the adoption year and by 2.58 five years after adoption. The pooled post-adoption effect is $-2.11$ (s.e.: 0.18). Against the pre-adoption mean multiple of 6.85 in Table 2, the pooled effect is a 31% decline in the premium charged per dollar of expected loss, and the five-year effect is a 38% decline. The specification excludes second-event, top-and-drop and drop-down layers, whose modeled expected losses are small relative to premium; Appendix Figure 20 reports the same specification with these layers retained, where the pooled effect is $-1.82$ (s.e.: 0.17). This effect is consistent with the model-implied price decrease we observe in Section 5 below.

Figure 6: Reinsurance Prices: Premium per Dollar of Expected Loss



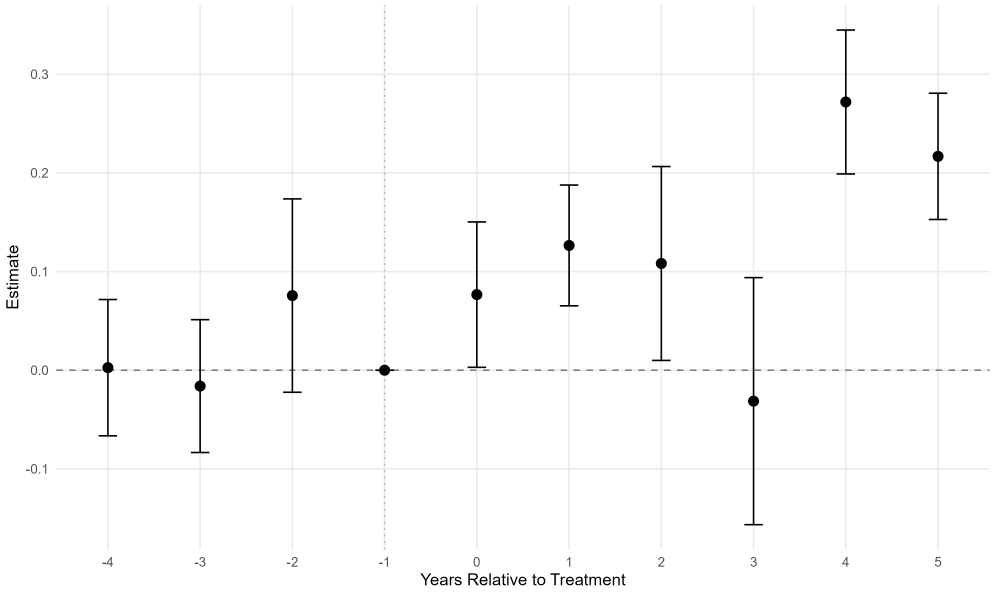

*Notes:* Event study of the carrier-year reinsurance price multiple, ceded premium divided by modeled expected loss over priced private excess-of-loss layers in Florida homeowners rate filings, 2004–2025. Coefficients are from the two-stage imputation specification in equations (2) and (3) with state-of-domicile and year fixed effects in Stage 1, unweighted, and reference period $\tau = -1$. The event window is $\tau \in [-4, 6]$ with observations beyond $\tau = 5$ binned to $\tau = 6$. Standard errors are clustered by state of domicile, and bars are 95% confidence intervals. The sample is 517 carrier-years, 100 carriers, and 16 domicile states. Negative coefficients denote cheaper reinsurance.

### 3.2.4 Tail-Risk Reinsurance Purchases

Reinsurance is more expensive when it covers less frequent tail events. As the price falls, certain tail risks that were not economical to buy reinsurance against might become so. To study this, we run the same event study where the outcome is whether a reinsurance program included protection against a 1-in-500 (or less frequent) loss. Before adoption, 11% of treated carrier-years buy cover this far into the tail.

Figure 7 reports the estimates. After adoption the probability of buying 1-in-500 cover rises by 0.08 (0.04) in the adoption year, 0.27 (0.04) four years out, and 0.22 (0.03) five years out. The pooled post-adoption effect is 0.11 (0.02), an 11 percentage point increase against a pre-adoption share of 11%.

Figure 7: Probability of Buying Reinsurance Above a 1-in-500 Year Loss




*Notes:* Event study where the outcome is an indicator for the carrier-year holding at least one priced private excess-of-loss layer with an implied occurrence period of 1-in-500 years or longer, from Florida homeowners rate filings, 2004–2025. Coefficients are from the two-stage imputation specification in equations (2) and (3) with state-of-domicile and year fixed effects in Stage 1, unweighted, and reference period $\tau = -1$. Standard errors are clustered by state of domicile, and bars are 95% confidence intervals. The sample is 424 carrier-years, 90 carriers, and 14 domicile states.

### 3.2.5 Counterparty Exposure and Expected Impairment Losses

While costly, the collateral requirement does protect cedents against the possibility that a reinsurer becomes impaired before paying its obligations. Reducing collateral therefore lowers the cost of supplying reinsurance but leaves a larger portion of the cedent's recoverable unsecured. We quantify this countervailing effect by combining cedent–reinsurer recoverables with rating-implied impairment probabilities and the statutory collateral schedule.

For cedent $i$ in year $t$, we define expected impairment loss as

$$\mathrm{EL}_{it} = \mathrm{LGD} \sum_{j} p(r_j^*)\, \lambda_{ijt}\, R_{ijt}, \tag{7}$$

where $R_{ijt}$ is the recoverable owed by reinsurer $j$, $p(r_j^*)$ is the one-year impairment probability associated with the reinsurer's financial-strength rating, and $\lambda_{ijt}$ is the uncollateralized share of the recoverable under the rules applicable to the cedent. The baseline assumes a $50\%$ recovery rate on unsecured claims (a loss given default, LGD, of $50\%$) and assigns unrated counterparties the median impairment probability among the observed rated tiers, $1.17\%$. Appendix A.3 describes the construction.

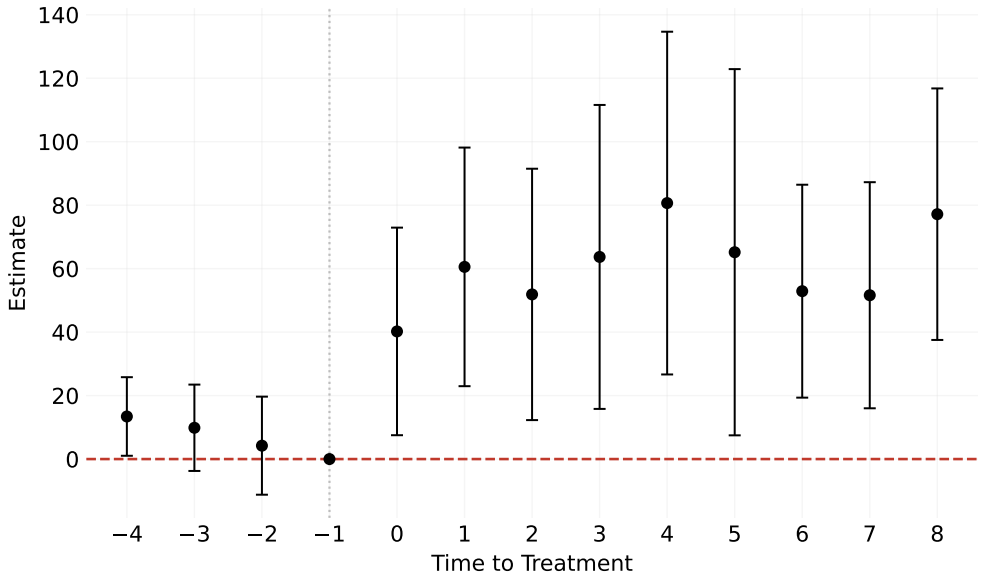
Figure 8 reports the event-study estimates. Expected impairment losses show little systematic movement before adoption and rise following adoption. The pooled estimate implies an increase of approximately \$57,600 per treated cedent-year, with a standard error of \$19,700. Scaling this per-cedent effect by the $1{,}509$ treated cedents observed in the final sample year implies an aggregate increase in expected impairment losses on the order of \$87 million per year. Because the outcome is measured in dollars, it captures the total change in cedents' counterparty exposure, including both the expansion of the ceded book documented above and the reduction in collateral. Scaled instead by the cedent's total reinsurance premium, expected loss per dollar ceded is essentially unchanged—an imprecisely estimated $-\$0.21$ per \$1,000 of premium ceded (Appendix A.3)—so the dollar increase reflects a larger, less-collateralized book rather than each dollar of reinsurance becoming riskier.

The composition of cedents' counterparties moved in the opposite direction. Following adoption, the share of premium ceded to unrated reinsurers declines by approximately $1.7$ percentage points. Under the baseline mapping from ratings to impairment probabilities, the premium-weighted average impairment probability also declines by $2.1$ basis points from a pre-reform mean of $47$ basis points. Cedents therefore appear to shift toward rated counterparties after adoption, consistent with the graded post-reform schedule: because the largest collateral reductions accrued

to the most highly rated foreign reinsurers, cedents had an incentive to move business toward them. The increase in dollar expected losses arises because the larger and less-collateralized exposure more than offsets this safer counterparty composition.

The magnitude is far less certain than the sign, since it depends on the recovery rate and especially on the impairment rate assigned to unrated reinsurers. We read it as a directional result on unsecured exposure rather than a loss forecast.

Figure 8: Expected Loss from Reinsurer Impairment




*Notes:* The figure reports an event study where the outcome is the model-implied expected impairment loss of equation (7), in thousands of dollars per cedent-year. Event time $\tau$ is relative to the cedent's domicile-state adoption of Model Law 785. Coefficients are from the two-stage imputation specification in equations (2) and (3), with cedent and year fixed effects in Stage 1 and reference period $\tau = -1$. The event window is $\tau \in [-4, 8]$. Standard errors are clustered on cedent domicile state, and bars are $90\%$ confidence intervals. The sample is unaffiliated third-party property-casualty reinsurance, a national book broader than the homeowners cessions in Figures 2–5. Baseline calibration: $50\%$ recovery on unsecured claims, the statutory collateral schedule, and unrated reinsurers assigned the median rated-tier one-year impairment probability ($1.17\%$); each reinsurer's end-of-sample financial-strength rating is applied to all years. Appendix A.3 reports the construction and additional estimates.

# 4 Effects on Downstream Home Insurance Rates and Coverage

## 4.1 Empirical Strategy: Triple-Difference with Wind vs. Non-Wind Coverage

The Florida QUASR panel is at the insurer-county-quarter level, which enables two extensions to the imputation framework of Section 3.1.1. First, we replace the year fixed effect $\lambda_t$ with a county-by-quarter fixed effect $\lambda_{ct}$, absorbing any shock common to all insurers operating in a given Florida county-quarter (local hurricane realizations, housing-market conditions, regulatory actions). We use this specification for the total insured value (TIV) market-share results in Section 4.2.2.

Second, QUASR separately reports wind-inclusive and wind-excluded policies within the same cell. Because the reinsurance collateral policy change relaxes collateral specifically for catastrophe-backing capital, it should primarily affect wind-exposed policies. For two outcomes we use the within-cell wind-minus-non-wind contrast directly. For pricing (Section 4.2.1),

$$\Delta^{r}_{ict} = r^{\text{wind}}_{ict} - r^{\text{nonwind}}_{ict}, \tag{8}$$

where $r^{\text{wind}}_{ict}$ and $r^{\text{nonwind}}_{ict}$ are the insurer's rate per \$1,000 of exposure. For entry (Section 4.2.3),

$$\Delta^{e}_{ict} = \mathbf{1}\{\text{wind present}\}_{ict} - \mathbf{1}\{\text{non-wind present}\}_{ict} \in \{-1, 0, +1\}. \tag{9}$$

Differencing within cell subtracts off any shock hitting wind and non-wind equally—common demand movements, firm-wide scale adjustments, local loss experience unrelated to reinsurance cost—leaving the wind-specific response on which the reform should load. Identification therefore requires parallel trends in the wind-minus-non-wind gap rather than in each level separately, a weaker requirement than in Section 3.1.1.

We estimate equations (8) and (9) using the same two-stage imputation procedure as before: Stage 1 fits $\Delta^{\bullet}_{ict} = \alpha_i + \lambda_{ct} + \varepsilon_{ict}$ for $\bullet \in \{r, e\}$ on not-yet-treated observations; Stage 2 regresses the residuals on event-time dummies. We weight by direct premiums written and cluster standard errors by state of domicile.

## 4.2 Results

### 4.2.1 Intensive Margin: Premium

We examine effects on wind premium coverage using Florida QUASR data. This analysis observes catastrophe exposure directly (wind coverage) and exploits treatment variation from insurer domicile state. Since QUASR covers Florida policies only, variation comes from insurers headquartered in different states with different reinsurance cost policy change adoption timing.

We estimate a triple-difference comparing wind-inclusive versus wind-excluded policy pricing. For each insurer-year, we calculate the rate per $1,000 of coverage separately for each type and take the difference—the *wind premium gap*. A negative coefficient indicates the reinsurance cost policy change narrowed this gap, consistent with improved catastrophe reinsurance access reducing the cost of wind coverage. The results are in Figure 9.

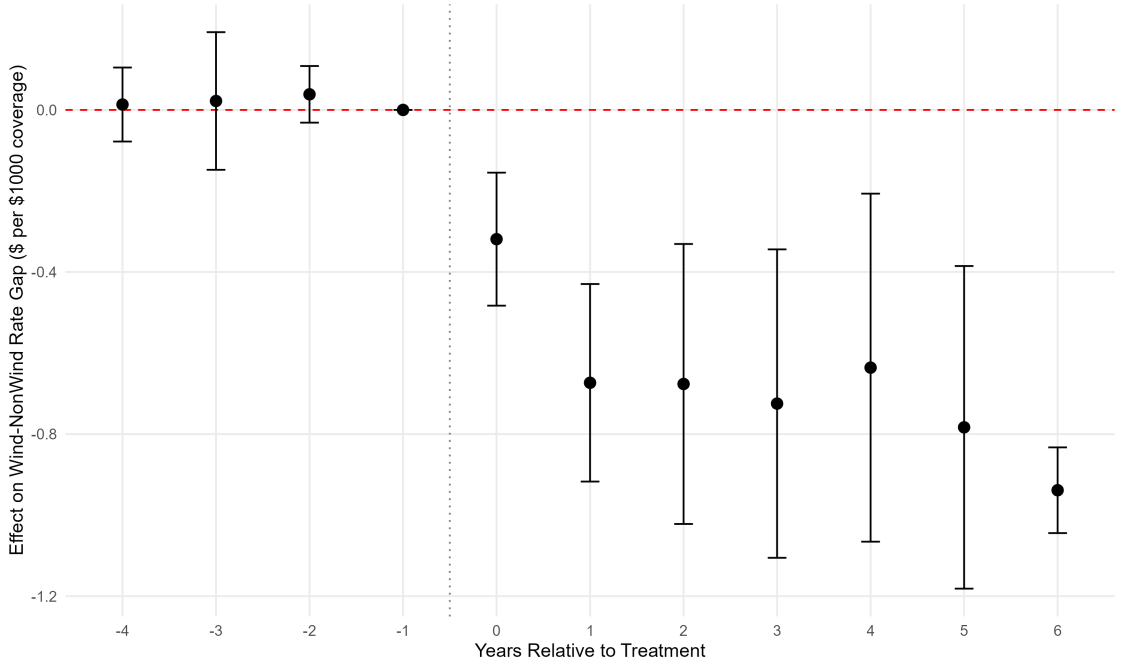
We see a clear drop in the pricing of wind-exposed homeowners insurance policies after treatment, relative to non-wind-exposed policies. Relative to the pre-treatment interquartile range of the rate per $1,000 of exposure ($3.15–$4.94; Table 3), a drop of 0.5 is a rate decrease of roughly 10–16%. As a placebo, the non-wind rate level falls by far less, and only at longer horizons (Appendix Figure 23). Moreover, the effect is similar among incumbent firms (Appendix Figure 24), ruling out an entrant-composition explanation. Finally, the decline is a direct price effect rather than a change in the composition of the insured portfolio: coverage per policy, wind-exposure per policy, and the rates of new business, cancellations, nonrenewals, and transfers show no movement around adoption (Appendix Figure 25).

Moreover, in Appendix C.1, we estimate heterogeneous treatment effects based on the average wind-risk in each county. The triple-difference estimates are about twice as large in high-wind-risk counties relative to low. This confirms the mechanism: reinsurance price decreases are passed through to insurance price decreases specifically in hurricane-exposed locations.

### 4.2.2 Insurance Coverage

We next examine whether the reinsurance cost policy change shifted the quantity of insurance sold, not just prices. For each insurer-county-quarter we compute the insurer's share of total insured value (TIV) in that county. This TIV-based share captures both the extensive margin (number of policies written) and the intensive margin (coverage depth per policy): if treated insurers expand by writing more policies, by raising coverage limits per policy, or both, market share will

Figure 9: Policy Pricing: Wind-Included vs. Wind-Excluded Coverage

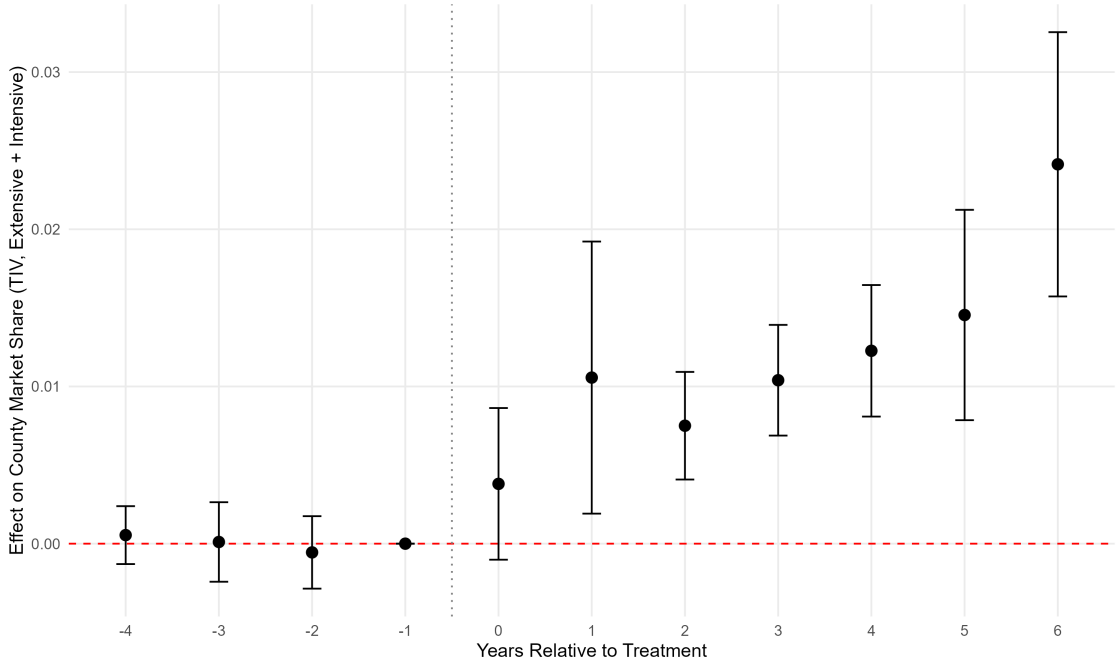



*Notes:* The figure reports a triple-difference event study estimated on insurer-county-quarter observations from Florida QUASR data. The outcome is the *wind premium gap*—the within-insurer-county-quarter difference in the rate per $1,000 of coverage between wind-inclusive and wind-excluded policies, following the structure of equation (8). Differencing wind from non-wind within insurer-county-quarter nets out any insurer- or market-level shock common to both coverage types, so identification requires parallel trends in the wind-minus-non-wind gap rather than in each level separately. Event time $\tau$ is measured in years relative to the adoption of the reinsurance collateral policy change by the insurer's state of domicile. Coefficients are estimated using the two-stage imputation procedure described in Section 4.1: Stage 1 fits insurer and county-by-quarter fixed effects on not-yet-treated observations, and Stage 2 regresses the imputed residuals on event-time dummies with reference period $\tau = -1$. The event window is $\tau \in [-4, 6]$ years with observations beyond $\tau = 5$ binned to $\tau = 6$. Negative coefficients indicate that wind-inclusive coverage became cheaper relative to wind-excluded coverage post-treatment. Estimation uses insurer direct-premiums-written weights and clusters standard errors by state of domicile.

rise. The event study in Figure 10 plots the reduced-form effect of the reinsurance cost policy change on the TIV market share of an insurer whose domicile state has adopted the reform, relative to insurers from not-yet-adopting states.

Treated insurers' share of county-level TIV expands post-reform. The post-period coefficients correspond to roughly a one-percentage-point gain on average across the post-reform period, rising to about two percentage points by the sixth year after adoption. Because TIV share co-moves with both the number of policies and the size of each policy, the effect is consistent with entry into new county-quarter markets, within-market expansion among incumbent firms, and upward adjustment of coverage limits—all symptoms of capacity expansion that one would expect if the

Figure 10: Insurance Market Share: Total Insured Value




*Notes:* The figure reports an event study estimated on an insurer-county-quarter panel constructed from Florida QUASR data. For each cell, the outcome is the insurer's share of total insured value (TIV) in the county-quarter, which co-moves with both the extensive margin (number of policies written) and the intensive margin (coverage depth per policy). Event time $\tau$ is measured in years relative to the adoption of the reinsurance collateral policy change by the insurer's state of domicile. Coefficients $\{\beta_\tau\}$ are from the two-stage specification in equations (2) and (3), with insurer fixed effects $\alpha_i$ and county-by-quarter fixed effects $\lambda_{ct}$ in Stage 1, and reference period $\tau = -1$. The event window is $\tau \in [-4, 6]$ years with observations beyond $\tau = 5$ binned to $\tau = 6$. Positive coefficients indicate that treated insurers expanded their share of county TIV post-reform. Standard errors are clustered by state of domicile.

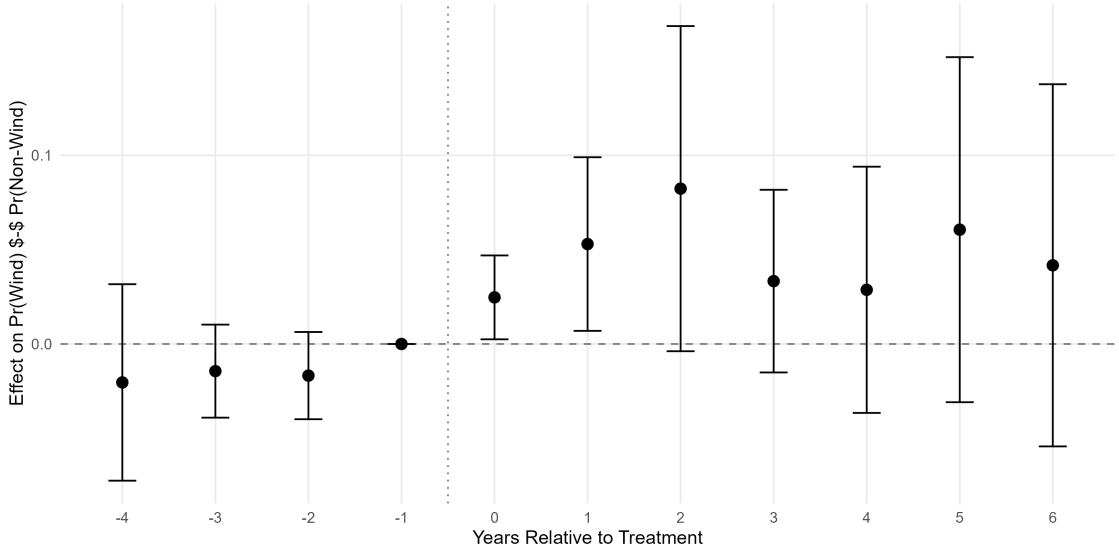
reinsurance cost policy change reduced the marginal cost of supplying wind-risk coverage. We study the extensive margin—entry into new areas—below.

### 4.2.3 Extensive Margin: Insurance Offerings

Cheaper reinsurance means that certain areas that were previously uneconomical to insure might become viable. We analyze whether insurers began writing wind-exposed policies in counties that they previously did not operate in. We define an indicator for whether the insurer is present in wind-inclusive coverage and an indicator for whether it is present in non-wind coverage, and then take the difference: $\Delta^e_{ict} = \mathbf{1}\{\text{wind present}\}_{ict} - \mathbf{1}\{\text{non-wind present}\}_{ict} \in \{-1, 0, +1\}$. This is the entry analogue of the pricing triple difference—it differences out any common entry/exit shock

that affects both coverage types and any firm- or market-level scale changes. Figure 11 reports the event-study estimates.

Figure 11: Extensive Margin Triple Difference: Presence in Wind − Presence in Non-Wind




*Notes:* The figure reports a triple-difference event study estimated on an insurer-county-quarter panel constructed from Florida QUASR data. The outcome is the within-cell difference between an insurer's presence in wind-inclusive coverage and its presence in non-wind coverage, $\Delta^{e}_{ict} = \mathbf{1}\{\text{wind present}\}_{ict} - \mathbf{1}\{\text{non-wind present}\}_{ict} \in \{-1, 0, +1\}$, as defined in equation (9). Differencing within insurer-county-quarter removes any common entry or exit shock affecting both coverage types and any firm- or market-level scale change, so identification requires parallel trends in the wind-minus-non-wind entry gap rather than in each margin separately. Event time $\tau$ is measured in years relative to the adoption of the reinsurance collateral policy change by the insurer's state of domicile. Coefficients are estimated using the two-stage imputation procedure described in Section 4.1: Stage 1 fits insurer and county-by-quarter fixed effects on not-yet-treated observations, and Stage 2 regresses the imputed residuals on event-time dummies with reference period $\tau = -1$. The event window is $\tau \in [-4, 6]$ years with observations beyond $\tau = 5$ binned to $\tau = 6$. Positive coefficients indicate treated insurers became relatively more likely to write wind-inclusive coverage post-treatment. Standard errors are clustered by state of domicile.

With imperfect pre-trends, the probability of insurers writing wind-exposed policies increases by approximately five percentage points after treatment. Combined with the intensive-margin pricing triple difference in Section 4.2.1 and the TIV market-share result in Section 4.2.2, these findings cohere: the reinsurance collateral policy change lowered the cost of catastrophe reinsurance, and treated insurers responded both by cutting the price of wind coverage and by selectively expanding their presence in exactly those segments where the catastrophe-reinsurance channel matters most.

# 5 An Equilibrium Model of the Wind Insurance Market

In the previous sections, we document how the reinsurance collateral policy change affected insurers' reinsurance purchases, product offerings, and homeowners' insurance premiums in Florida. Building on these findings, we develop a structural model of demand and pricing in the Florida wind insurance market. The goal is to quantify the share of premium variation attributable to reinsurance costs and enhanced competition, and subsequently, evaluate the welfare effects of the reform.

Our model is a simple equilibrium demand and supply framework. Homeowners in Florida counties choose among wind-coverage products based on heterogeneous preferences over premiums and other product characteristics. Insurers simultaneously set premiums for their products at the county level, taking into account demand, marginal claims costs, and the cost of ceding risk to reinsurers. In each period, insurers move first, followed by consumers making their product choices. We analyze the Nash-Bertrand equilibrium of this game.

Our focus on pricing is informed by the reduced-form findings in Section 4. In our primary specification, we hold market structure fixed and do not model insurer entry or exit.[6] We treat the reinsurance market as exogenous and do not model reinsurer pricing or contract design. This is plausible because Florida-focused primary insurers are typically small relative to global reinsurers, and are better described as price-takers in the reinsurance market than as Nash bargainers.

## 5.1 Model Set-up

**Consumer Choices.** Let $i$ denote consumers, $f$ denote insurers, $j \in \mathcal{J}_{ft}$ denote products of $f$ in period $t$, $m$ denote counties, and $t$ denote year-quarters. A market is a county-quarter pair. We focus on wind-peril homeowner products, which are the products most directly exposed to reinsurance costs. Market size $M_{mt}$ equals the number of owner-occupied housing units in county $m$ at time $t$.[7]

[6] In ongoing work, we extend the framework to include a supply-side entry stage, in which insurers decide whether to operate in a given county-year based on expected variable profits net of a fixed operating cost. The counterfactuals reported hold the set of active insurers fixed.

[7] We define market size as owner-occupied housing units in ACS, linearly interpolated from annual estimates to quarterly frequency. The implied mean outside share is $0.39$. A natural alternative is to restrict the denominator to mortgaged owner-occupied units, since mortgaged households are contractually required to carry hazard insurance. We reject this on empirical grounds: in our wind-insurance estimation sample, policy counts exceed mortgaged owner-occupied housing units in 1,477 of 2,935 county-quarters (50.3%), with the excess concentrated in retirement and seasonal-home counties. The excess reflects voluntary purchase by free-and-clear owners and insurance on seasonal homes outside the ACS principal-residence universe. Across counties, HO take-up rises with mortgage penetration, validating the uninsured homeowner margin as the outside option.

The demand for insurance products follows a standard Berry et al. (1995, BLP) framework. Consumer $i$ chooses product $j$ that maximizes her flow utility, which is

$$u_{ijmt} = -\alpha_i \, p_{jmt} + \beta_i^0 + \mathbf{x}'_{jmt}\boldsymbol{\beta} + \xi_{jmt} + \varepsilon_{ijmt}, \quad j \neq 0, \tag{10}$$

$$\alpha_i = \bar{\alpha} + \exp(\sigma_\alpha \, \nu_i^\alpha), \quad \nu_i^\alpha \sim N(0,1), \tag{11}$$

$$\beta_i^0 = \bar{\beta}_0 + \sigma_0 \, \nu_i^0, \quad \nu_i^0 \sim N(0,1), \tag{12}$$

where $p_{jmt}$ is the annual premium (in \$1,000s), $\xi_{jmt}$ is unobserved product quality, and $\varepsilon_{ijmt}$ is an i.i.d. type-I extreme value error. The price coefficient $\alpha_i$ is lognormally distributed, with $\bar{\alpha}$ governing the mean disutility of premiums and $\sigma_\alpha$ governing the dispersion of price sensitivity across the population. A separate normally distributed random coefficient $\beta_i^0$ on the inside constant allows for heterogeneous valuations of the inside option relative to the outside option of not purchasing wind insurance. The set of covariates $\mathbf{x}_{jmt}$ controls for firm, county, and year-quarter fixed effects.[8] We normalize the outside option of not purchasing a wind-coverage product so that $u_{i0mt} = \varepsilon_{i0mt}$.[9]

**Insurers' Pricing.** Insurer $f$ simultaneously sets premiums to maximize variable profits,

$$\max_{\{p_{jmt}\}_{j \in \mathcal{J}_{ft}}} \sum_{j \in \mathcal{J}_{ft}} \sum_{m} \underbrace{(p_{jmt} - mc_{jmt})}_{\text{margin}} \cdot \underbrace{M_{mt} \cdot s_{jmt}(\mathbf{p}_t; \boldsymbol{\theta})}_{\text{quantity}}, \tag{13}$$

where $s_{jmt}(\cdot)$ is the predicted market share implied by the demand system in equation (10). Nash-Bertrand pricing yields the familiar first-order conditions

$$\mathbf{p}_t = \mathbf{mc}_t + \boldsymbol{\Delta}_t^{-1}\mathbf{s}_t, \tag{14}$$

where $\boldsymbol{\Delta}_t$ is the matrix of share derivatives with off-diagonal blocks zeroed out across firms.

We parameterize marginal costs as:

$$\ln(mc_{jmt}) = \gamma_0 + \gamma_1 \times \text{post785}_{ft} + \gamma_2 \times \text{post785}_{ft} \times \text{I}_f + \gamma_3 \times \text{post785}_{ft} \times \text{I}_f \times \text{AAL}_m + \bar{\mathbf{w}}'_{fmt}\boldsymbol{\gamma}_m + \omega_{jmt}. \tag{15}$$

$\text{post785}_{ft}$ is an indicator for whether insurer $f$'s state of domicile has adopted the reinsurance collateral policy change by period $t$; its coefficient $\gamma_1$ summarizes how reinsurance-market reform passes through to insurers' marginal costs. $\text{I}_f$ is the insurer's pre-reform share of ceded premium to foreign unauthorized reinsurers. $\text{post785}_{ft} \times \text{I}_f$ measures the insurer's exposure to the collat-

[8] Empirically, we implement the high-dimensional fixed effects using Mundlak-Chamberlain group means of premiums by firm, county, and year-quarter. Including these group means as linear regressors is econometrically equivalent to absorbing firm, county, and year-quarter fixed effects (Mundlak, 1978; Chamberlain, 1984).

[9] We model Florida's state-sponsored product, Citizens Property Insurance, as an inside good with regulated prices.

eral reform, and $\gamma_2$ measures the differential effect of the collateral reform by treatment intensity. $\text{AAL}_m$ is the county average annual loss and captures differences in catastrophe risk across Florida. The coefficient on the interaction term between reform intensity and $\text{AAL}_m$, $\gamma_3$, allows the exposure gradient to vary with county risk. $\mathbf{w}_{fmt}$ includes log exposure per policy, which proxies for the scale of insured structures; the Citizens indicator, which absorbs the fact that Florida's government-sponsored residual-market insurer operates under a distinct cost structure; $\text{AAL}_m$ and $\text{I}_f$ separately as controls; county-by-year fixed effects, which control for the riskiness level of the geographic market, and firm fixed effects. $\omega_{jmt}$ is the cost residual.

### 5.2 Identification and Estimation

**Demand-Side.** Premiums are endogenous because insurers observe preference shocks $\xi_{jmt}$ when setting prices, inducing correlation between $p_{jmt}$ and $\xi_{jmt}$, which is unobservable to the econometrician. We address this endogeneity problem with three complementary instruments. First, we construct standard Berry et al. (1995) instruments: the number and summed characteristics of competing wind-coverage products in each market. The identifying assumption is that the density of nearby rivals shifts equilibrium markups without directly affecting demand for a given product, conditional on firm, county, and year-quarter fixed effects. Second, we add Gandhi and Houde (2019) differentiation instruments, which measure the degree of local isolation of each product in characteristic space and provide additional identifying variation when BLP-style sums are nearly collinear.

Third, we exploit variation in insurers' loss experience outside Florida. Specifically, we use the lagged logarithm of non-Florida homeowners and non-homeowners losses, and lagged inverse loss ratios outside Florida. The identifying assumption is that a multi-state insurer's losses in, e.g., Texas, shift its financial position and reinsurance demand, and hence its Florida premiums, but are mean-independent of the unobserved quality of its Florida products. This is plausible because homeowners' insurance demand is highly local. Following Borusyak and Hull (2023), we recenter these shocks by subtracting within-market means to avoid biases from mechanical correlation between own and rival shares of the same shock.

Consumers' mean price preference, $\bar{\alpha}$, is identified by the correlation between the BLP, differentiation, and loss-based cost-shifter instruments and market shares: markets in which rival markups are higher or in which multi-state insurers face larger out-of-state losses see higher equilibrium prices, and the share response to these price shifts pins down mean price sensitivity. The dispersion of the price random coefficient, $\sigma_\alpha$, is identified by second-moment variation: substi-

tution patterns that deviate from the IIA pattern implied by a pure logit—in particular, differential within-market substitution between similarly priced and differently priced wind products—identify dispersion in price sensitivity (Berry et al., 1999; Berry and Haile, 2014). The dispersion of the inside good random coefficient, $\sigma_0$, is identified by variation in the total inside share relative to what the mean utility alone would predict.

**Supply-Side.** The demand estimates determine how market shares respond to premiums. Given observed ownership and the Nash-Bertrand pricing assumption, equation (14) maps these demand derivatives into insurers' markups, and marginal costs are then recovered as observed premiums net of the implied markups. Decomposing premiums into markups and marginal costs therefore relies on the demand estimates and the pricing assumption; it does not require an excluded supply-side instrument, and we use none. The supply moments use the included cost variables as instruments and the supply side is exactly identified.

The reform effect, $\gamma_1$, is identified by variation in the timing of Model Law 785 adoption across insurers' states of domicile. The domicile state is the relevant treatment assignment because its credit-for-reinsurance law determines the collateral treatment of the insurer's reinsurance recoverables. Intuitively, after removing the markups implied by the demand system, differences in adoption timing across insurers trace out how the reform changes marginal costs. The identifying assumption is that, conditional on the observed cost variables and Mundlak controls, domicile-state adoption is mean-independent of other changes in insurers' unobserved Florida marginal costs:

$$E\left[\omega_{jmt} \mid \text{post785}_{ft}, \mathbf{x}^{c}_{jmt}\right] = 0. \tag{16}$$

Since $\mathbf{x}^{c}_{jmt}$ includes year-quarter fixed effects and the firm means of the reform terms, $\gamma_1$ compares a treated insurer's recovered costs before and after adoption with those of not-yet-treated and never-treated insurers in the same quarter. This is the parallel-trends assumption, analogous to the reduced form.

We also allow the reform effect to vary with the insurer's pre-reform reliance on foreign unauthorized reinsurers and with county catastrophe risk. The exposure interaction is identified by differences in post-adoption cost changes across insurers with different pre-reform portfolios. The interaction with county average annual loss asks whether this exposure gradient varies with catastrophe risk, identified by differences in post-adoption cost changes across insurers with similar pre-reform portfolios but different catastrophe-risk counties. With the intensity terms included, $100\left[\exp(\gamma_1 + \gamma_2 \mathrm{I}_f + \gamma_3 \mathrm{I}_f \mathrm{AAL}_m) - 1\right]$ is the percentage cost change for an insurer with exposure $\mathrm{I}_f$ in a county with risk $\mathrm{AAL}_m$.

**Estimation** The primitives to be estimated include consumers' price sensitivity $\bar{\alpha}, \sigma_\alpha$, preferences for wind-product and other product characteristics $\bar{\beta}_0, \sigma_0, \boldsymbol{\beta}$, and insurers' cost components $\boldsymbol{\gamma}$. We estimate these primitives with a generalized method of moments estimator, jointly using demand-side and supply-side moments:

$$\mathbb{E}\Big[\xi_{jmt}(\boldsymbol{\theta}) \cdot \mathbf{z}^d_{jmt}\Big] = \mathbf{0}, \qquad \mathbb{E}\big[\omega_{jmt}(\boldsymbol{\theta}) \cdot \mathbf{z}^s_{jmt}\big] = \mathbf{0}. \tag{17}$$

Here, $z^d_{jmt}$ contains the demand instruments described above, while $z^s_{jmt}$ contains all regressors in the marginal-cost equation (15), which serve as their own instruments. We use no excluded supply-side instruments.

### 5.3 Estimation Results

Table 4 reports our estimates. We first estimate a common reform-related cost shift in Column (1), and then add the interaction terms to allow marginal cost shifts to vary with the insurer's exposure to foreign unauthorized reinsurance and with county catastrophe risk in Columns (2) and (3). All columns yield economically similar magnitudes in both estimates and derived statistics. Therefore, we will focus our discussion on column (3).

**Consumer Preferences and Elasticities.** Panel (a) of Table 4 reports consumer preference estimates. The mean price coefficient $\bar{\alpha}$ is 2.69 (in \$1,000s), confirming that consumers dislike increases in the annual premiums.[10] The estimated scale parameter for the price random coefficient is 0.86, indicating substantial heterogeneity in price sensitivity.

We report two additional statistics to characterize the substitution patterns implied by these estimates. First, we calculate product-level own-price elasticities at observed premiums, $\varepsilon_{jj} = \partial \log s_j / \partial \log p_j$. The mean and median are $-4.31$ and $-3.45$, respectively, with an interquartile range of $-4.41$ to $-2.82$.

Second, on the extensive margin, we calculate the partial-equilibrium response of take-up to a uniform shock to the premium of each inside product. Consumers who own their homes outright can make this extensive margin decision while households with a mortgage are often required to insure against wind damages by their lenders. The baseline take-up of wind coverage, weighted by owner-occupied housing units, averages $60.5\%$ across county-quarter markets ($57.9\%$ in high-risk counties; $68.0\%$ in low-risk counties). A proportional 10% price decrease raises take-up to

[10] The price coefficient is $\alpha_i = [\beta_{-p} + \exp(\sigma_p \nu_i)]$, where $\nu_i \sim N(0,1)$. Thus, $E[\alpha_i] = \beta_{-p} + \exp(\sigma_p^2/2)$. The estimates $\hat{\beta}_{-p} = 1.242$ and $\hat{\sigma}_p = 0.858$ imply a mean of 2.69.

Table 4: Joint BLP Estimates: Demand and Cost Function

| | (1) Post785 | (2) + Reform intensity | (3) + AAL and interaction |
|---|---|---|---|
| *(a). Demand: Linear Parameters and Random Coefficients* | | | |
| Fixed component of price sensitivity | $1.242^{***}$ | $1.242^{***}$ | $1.242^{***}$ |
| | (0.159) | (0.159) | (0.159) |
| Dispersion of price RC | $0.858^{***}$ | $0.858^{***}$ | $0.858^{***}$ |
| | (0.047) | (0.047) | (0.047) |
| Dispersion of constant RC | 0.000 | 0.000 | 0.000 |
| *(b). Cost Function:* $\log mc = w'\gamma + \omega$ | | | |
| Constant | $-7.271^{***}$ | $-7.352^{***}$ | $-8.835^{***}$ |
| | (0.311) | (0.312) | (0.340) |
| log(exposure/policy) | $0.394^{***}$ | $0.393^{***}$ | $0.383^{***}$ |
| | (0.026) | (0.026) | (0.027) |
| Citizens indicator | $0.321^{***}$ | $0.334^{***}$ | $0.349^{***}$ |
| | (0.009) | (0.009) | (0.009) |
| Post Model Law 785 | $-0.088^{***}$ | $-0.099^{***}$ | $-0.101^{***}$ |
| | (0.004) | (0.005) | (0.004) |
| Reform intensity | – | $0.032^{***}$ | $0.033^{***}$ |
| | | (0.004) | (0.008) |
| County AAL main effect | – | – | $0.083^{***}$ |
| | | | (0.003) |
| Reform intensity × AAL | – | – | 0.002 |
| | | | (0.002) |
| *(c). Implied Statistics* | | | |
| Mean own-price elasticity | −4.306 | −4.305 | −4.306 |
| Median own-price elasticity | −3.450 | −3.450 | −3.450 |
| Mean marginal cost ($1,000s) | 1.801 | 1.801 | 1.801 |
| Mean observed price ($1,000s) | 2.296 | 2.296 | 2.296 |
| Mean Lerner index $(p - mc^{\mathrm{FOC}})/p$ | 0.290 | 0.290 | 0.290 |
| Reinsurance / marginal cost (aggregate) | 0.613 | 0.613 | 0.613 |
| Mean own-cost pass-through ($\partial p/\partial mc$) | 1.065 | 1.065 | 1.065 |
| Products with $\lvert\varepsilon_{jj}\rvert < 1$ (%) | $< 0.01$ | $< 0.01$ | $< 0.01$ |
| Demand FE-demeaned first-stage $F$ | 617.6 | 617.6 | 617.6 |
| Supply FE-demeaned first-stage $F$ | – | – | – |
| GMM objective | 3,763.6 | 3,763.6 | 3,763.6 |
| Observations | 152,507 | 152,507 | 152,507 |
| Markets | 2,935 | 2,935 | 2,935 |

*Notes:* Robust standard errors are in parentheses; $^{***}p < 0.01$, $^{**}p < 0.05$, and $^{*}p < 0.10$. All columns jointly estimate demand and log marginal costs, including Mundlak controls and cost-equation year fixed effects. Column (2) adds reform intensity. Column (3) further adds county AAL and its interaction with reform intensity. The price coefficient on $-p$ is $\alpha_i = \beta_p + \exp(\sigma_p \nu_i)$, with $\nu_i \sim N(0,1)$; panel (a) reports the fixed component $\beta_p$ and log-scale $\sigma_p$. Prices and marginal costs are in $1,000s of annual premium per policy.

Panel (c) reports unweighted product-market statistics, including Citizens unless indicated otherwise. Marginal cost is structural, $mc = \exp(\mathbf{w}'\widehat{\gamma} + \widehat{\omega})$; the Lerner index uses FOC-based costs, $(p - mc^{\mathrm{FOC}})/p$. Reinsurance/MC is $\sum_j R_j / \sum_j mc_j$ over observations with nonmissing $R_j$ and positive MC, where $R_j$ is non-affiliated reinsurance premium allocated per policy using raw firm-year cession-to-DPW ratios. Own-cost pass-through averages $\partial p_j/\partial mc_j$ across private products, holding Citizens prices fixed. Sample: wind-only Florida homeowners insurance, 2010Q1–2020Q4, covering 131 firms.

66.7%, implying a log arc elasticity of take-up with respect to price of $-0.93$; a symmetric $10\%$ price increase reduces take-up to 54.3%, implying an elasticity of $-1.13$. In additive terms, a $100 reduction in annual premiums raises take-up by roughly 3.2 percentage points, while a $500 reduction raises it by 15.2 points.

**Cost Functions and Markup.** Panel (b) of Table 4 reports the cost function estimates. The log exposure elasticity is $\hat{\gamma}$ = 0.38, indicating that a 10% increase in the insured value per policy raises marginal cost by roughly 3.7%. The Citizens indicator coefficient is $\hat{\gamma}_c = 0.35$, implying approximately 42% higher marginal cost than for an otherwise comparable private product. This difference is consistent with Citizens' role as Florida's residual-market insurer.[11]

Each unit of county AAL raises cost by 8.6 percent, so a county at the 75th percentile of hurricane risk (AAL 4.65) has marginal costs about 35 percent above one at the 25th (AAL 0.99). The constant $\hat{\gamma}_0$ reflects the cost level after Mundlak partialling and is not directly interpretable; we instead report the implied marginal cost in panel (c).

The reinsurance-reform coefficient is the parameter of central interest. Column (1) estimates a common reform effect, $\hat{\gamma}_r = -0.088$, implying that adoption of the collateral-policy change in an insurer's state of domicile reduces its marginal cost of supplying Florida wind coverage by $1 - \exp(-0.088) \approx 8.4\%$. This coefficient captures the reform's effect on total recovered marginal cost. The reinsurance-price results presented earlier provide separate evidence that lower reinsurance costs are an important mechanism.

Table 5 converts the column (3) coefficients into proportional cost effects at three county-risk percentiles and three exposure levels. An insurer with no pre-reform unauthorized exposure sees its marginal cost fall by 9.6 percent (95 percent interval $[-10.4, -8.8]$) in every county, since the AAL main effect cancels from the contrast. At the median exposure, the reduction is 8.8 percent in a 25th-percentile county and 8.7 percent in a 75th-percentile county; at full exposure, it is 6.4 and 5.7 percent. Comparing across rows, the proportional cost reduction does not vary with catastrophe risk: the reform's dollar effect is larger in high-risk counties only because their costs are higher, so the same 9.6 percent is worth about 35 percent more per policy in a 75th- than in a 25th-percentile county. Comparing across columns, the reduction is smaller for insurers with greater pre-reform reliance on foreign unauthorized reinsurers.

The positive exposure coefficient on reform intensity does not mean the reform mattered less for insurers that relied more on foreign unauthorized reinsurance. The cost function measures

[11] For a 10% increase in exposure, $mc_{\text{new}}/mc_{\text{old}} = (1.1)^{\hat{\gamma}}$. For the Citizens indicator, $mc_{\text{Citizens}}/mc_{\text{non-Citizens}} = \exp(\hat{\gamma}_c)$. In each case, subtracting one from the cost ratio and multiplying by 100 gives the percentage change.

Table 5: Implied Marginal-Cost Effects of Model Law 785

| County risk (AAL percentile) | Nearest county | $AAL_c$ | No exposure $s = 0$ | Median exposure $s = 0.229$ | Full exposure $s = 1$ |
|---|---|---|---|---|---|
| P25 | Gilchrist / Jackson (AAL = 0.960 / 1.017) | 0.988 | $-9.57\%$ $[-10.35, -8.77]$ | $-8.84\%$ $[-9.67, -8.01]$ | $-6.38\%$ $[-7.76, -4.97]$ |
| P50 | Hernando (AAL = 1.823) | 1.823 | $-9.57\%$ $[-10.35, -8.77]$ | $-8.81\%$ $[-9.59, -8.02]$ | $-6.21\%$ $[-7.33, -5.09]$ |
| P75 | Lee / Brevard (AAL = 4.639 / 4.655) | 4.647 | $-9.57\%$ $[-10.35, -8.77]$ | $-8.68\%$ $[-9.36, -8.00]$ | $-5.66\%$ $[-6.50, -4.81]$ |

*Notes:* Entries report the model-implied percentage change in marginal cost from switching Model Law 785 from zero to one, holding other cost determinants fixed, using the estimates in Table 4, column (3):

$$100\left[\exp\left\{\widehat{\gamma}_{785} + s\left(\widehat{\gamma}_I + \widehat{\gamma}_{I\times AAL} AAL_c\right)\right\} - 1\right].$$

Brackets report pointwise 95% confidence intervals, constructed on the log-cost scale using the full robust coefficient covariance matrix and then transformed into percentage changes. Exposure and county AAL are held fixed when constructing these intervals. The AAL main effect cancels from the reform contrast.
Exposure $s$ is the firm's share of ceded premiums going to foreign unauthorized reinsurers, constructed from Schedule F filings. Pre-adoption filings are used where available, including earlier pre-adoption years when necessary; otherwise, the first observed positive-cession years are used. Exposure is set to zero when no usable filings are available. Median exposure ($s = 0.229$, rounded) is calculated across the 113 firms observed after treatment; $s = 0$ and $s = 1$ are theoretical endpoints.
Rows correspond to the unweighted 25th, 50th, and 75th percentiles of county average annual loss (AAL) across Florida's 67 counties. County names and their associated AAL values provide geographic reference; effects are evaluated at the percentile values. The table reports marginal-cost effects, not equilibrium changes in premiums or welfare.

the reform's effect on marginal cost after insurers re-optimize their reinsurance programs, and the Schedule F evidence earlier shows that treated insurers responded to cheaper collateral by ceding more. For an insurer with little pre-reform exposure, the reform's main effect is a lower price on cessions it can now place with unauthorized reinsurers, and the saving shows up as lower marginal cost. For an insurer already reliant on unauthorized reinsurers, the reform lowers the price of the marginal layer it was previously constrained from buying, and part of the saving is spent on additional coverage: ceded premium per policy falls by less, or rises, even though the price of risk transfer has fallen. Consistent with this reading, treated insurers with higher pre-reform exposure increased cessions per policy by more after adoption. The cost coefficient for exposed insurers is therefore net of their reinsurance response: it understates the reform's effect on the price of risk transfer, and the consumer-surplus gains in the counterfactual section exclude the value of the additional coverage these insurers bought, which accrues as lower insolvency risk rather than lower premiums.

Taking the unweighted average of the product-market Lerner indices, $\overline{(p_j - mc_j^{\text{FOC}})/p_j}$, yields a mean of 0.29.[12] This magnitude is close to the direct-business underwriting margin of 0.27 that we calculate from insurers' NAIC filings.

The non-affiliated reinsurance premium ceded per policy averages 0.61 of marginal cost in our preferred specification, with values being the same in all three specifications. Multiplying by one minus the mean Lerner index, $(1 - \bar{L}) = 0.71$, gives an approximate conversion to the more familiar reinsurance-share-of-direct-premium basis, giving $0.613 \times 0.710 \approx 0.44$ in our preferred specification. These line up almost exactly with the FL-specific benchmark we compute directly from Schedule F non-affiliated ceded premium (excluding fronting transactions) joined to NAIC state-page direct homeowners premium written: averaged across FL writers over 2010–2020, the pro-rata-weighted $C_R/p_{HO}$ is $0.51$ and the HO-dominant-firm-weighted ratio is $0.52$. Both ratios lie slightly above the upper bound of the $[0.30, 0.50]$ band typically cited for Florida homeowners in the catastrophe-reinsurance literature (Froot and O'Connell, 2008; Born and Klimaszewski-Blettner, 2013). The agreement between our structural estimate and the accounting share is independent evidence that the cost equation recovers the loss-and-reinsurance-loading component of marginal cost in the proportion implied by direct decomposition of insurer financials.

**Pass-Through of Cost Reduction to Prices.** We further calculate the structural pass-through rates implied by equation (14). The mean own-cost pass-through ($\partial p_j/\partial mc_j$) is 1.065 (median 1.063, interquartile range $[1.038, 1.094]$), with Citizens' prices held fixed. Pass-through above one can arise under imperfect competition because demand curvature affects the response of markups to costs, making the price overshoot uniform cost shocks (Weyl and Fabinger, 2013). In percentage terms, the mean own-cost elasticity of premiums among treated products is 0.748.

For illustration, applying this elasticity to a $9.6\%$ marginal-cost reduction gives a first-order own-cost contribution of approximately $7.2\%$ to premium reductions. The full equilibrium effect under a fixed market structure additionally incorporates cross-product cost effects and the heterogeneous cost reductions implied by the reform.

[12] Our estimated marginal cost is gross of reinsurance ceded premium and gross of reinsurance recoveries, so the Lerner index measures the wedge between premium and total expected claims and loading cost.

# 6 Welfare Effect of the Collateral Reform

We now use the estimated model to quantify how the collateral reform affected consumers in the Florida wind insurance market. The reform operates on two margins. First, the reform changes insurers' total marginal costs and induces private insurers to adjust their premiums. Second, the post-reform market contains additional product offerings, which alter the competitive environment.

We measure consumer welfare by the change in expected consumer surplus implied by the estimated demand model, evaluated at each counterfactual's equilibrium premiums and product offerings.[13]

We report gains in dollars per owner-occupied household per year, including households that choose the outside option. We average these gains across county-quarter markets using owner-occupied housing units as weights; statewide annual totals use the panel-average number of owner-occupied households, approximately 4.9 million.

**Counterfactual Design.** Holding the set of operating insurers and the residual market regulator fixed, we re-simulate Bertrand pricing equilibria in four combinations of marginal costs and product availability. Counterfactual (S4) is the no-reform equilibrium. Insurers face no-reform marginal costs, and the choice set excludes products classified as post-adoption additions. Counterfactual (S2) introduces the reform-related reduction in marginal costs while retaining the same restricted product set. Counterfactual (S3) retains no-reform marginal costs but introduces the full observed product set. Finally, Counterfactual (S1) combines reform marginal costs with the full observed product set.

We highlight two channels from the comparison across counterfactual scenarios. First, the difference between (S4) and (S2) changes marginal costs while holding the product set fixed. Because private insurers reoptimize their premiums, this comparison combines the direct effect of lower marginal costs with the equilibrium response of markups. It is therefore not a mechanical one-

[13] We compute expected consumer-surplus gains using the standard log-sum formula for the estimated random-coefficients demand model. For counterfactuals $a$ and $b$, the gain in county-quarter market $m$ is

$$\Delta \mathrm{CS}_m^{a,b} = \sum_i \frac{w_{im}}{\alpha_i} \log \left( \frac{1 + \sum_{j \in \mathcal{J}_m^a} \exp\left(v_{ijm} - \alpha_i p_{jm}^a\right)}{1 + \sum_{j \in \mathcal{J}_m^b} \exp\left(v_{ijm} - \alpha_i p_{jm}^b\right)} \right),$$

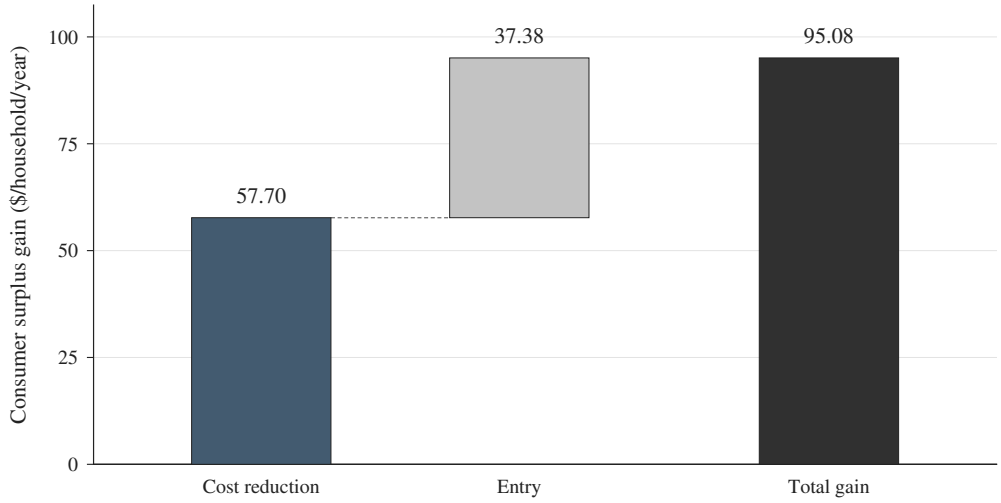
where $w_{im}$ are integration weights over consumer heterogeneity, $\alpha_i > 0$ is the marginal disutility of premiums, $p_{jm}^k$ is the equilibrium annual premium in scenario $k$, and $\mathcal{J}_m^k$ is its available product set. The term $v_{ijm}$ denotes utility excluding the price term and the idiosyncratic choice shock, and is held fixed across scenarios. The outside option has normalized utility zero. Premiums and surplus are expressed in dollars, with $\alpha_i$ scaled accordingly.

for-one pass-through calculation. Second, the difference between (S2) and (S1) adds the observed post-adoption products while holding the reform cost schedule fixed. Firms again reoptimize their premiums in the expanded product set. This comparison captures the effect of additional offerings on equilibrium surplus. Because the model does not include an entry stage, the product-set comparisons condition on the observed post-adoption offerings rather than predicting firms' entry decisions.

**Aggregate Effects.** Comparing the reform equilibrium (S1) with the no-reform equilibrium (S4), average annual premiums for incumbent private products fall by 5.7%, approximately 6%.[14]

Figure 12 shows that the reform raised consumer surplus by roughly $95 per household per year, about 4 percent of the average premium, or $0.47 billion per year statewide. Of these gains, lower costs account for 60% ($58) and entry for 40% ($37).

Figure 12: Welfare decomposition of the reinsurance collateral reform




The cost channel works through the price of risk transfer. Before the reform, a foreign reinsurer had to post collateral equal to its full obligations for the ceding insurer to receive statutory credit, and the cost of that trapped capital was priced into the reinsurance premium the primary insurer paid. The reform lowered the requirement for highly rated reinsurers, and the cost function shows that this lowered treated insurers' marginal cost of writing wind coverage by 9.6 percent for an in-

[14]We calculate this percentage change between policy-weighted mean premiums, restricting the comparison to private products present in both scenarios and using the same observed policy counts as weights. Citizens is excluded because its premiums are held fixed. Newly available products have no corresponding price in S4 and are therefore excluded from this price comparison.

surer with no pre-reform foreign-unauthorized exposure. Nash-Bertrand pricing then determines how much of the cost reduction reaches homeowners. The cost channel's $58 is the premium reduction homeowners would have received if no insurer entered a new county, combining the direct effect of lower marginal costs with the equilibrium pricing responses of treated insurers and their competitors.

The entry channel also relates to the price of risk transfer. Writing wind coverage in a catastrophe-exposed county requires reinsurance capacity, and collateral requirements raised costs most where exposure was largest. Cheaper reinsurance raises the variable profit an insurer expects from a county it does not yet serve and tips the marginal counties into profitability; the reduced form finds that adoption raised a treated insurer's probability of presence by 5 to 9 percentage points. Furthermore, additional products place competitive pressure on incumbents, which respond by adjusting their premiums. The entry channel's $37 is the value of new, lower-priced products to households, combining the direct effect of increased variety with the equilibrium pricing responses of incumbents and their competitors.

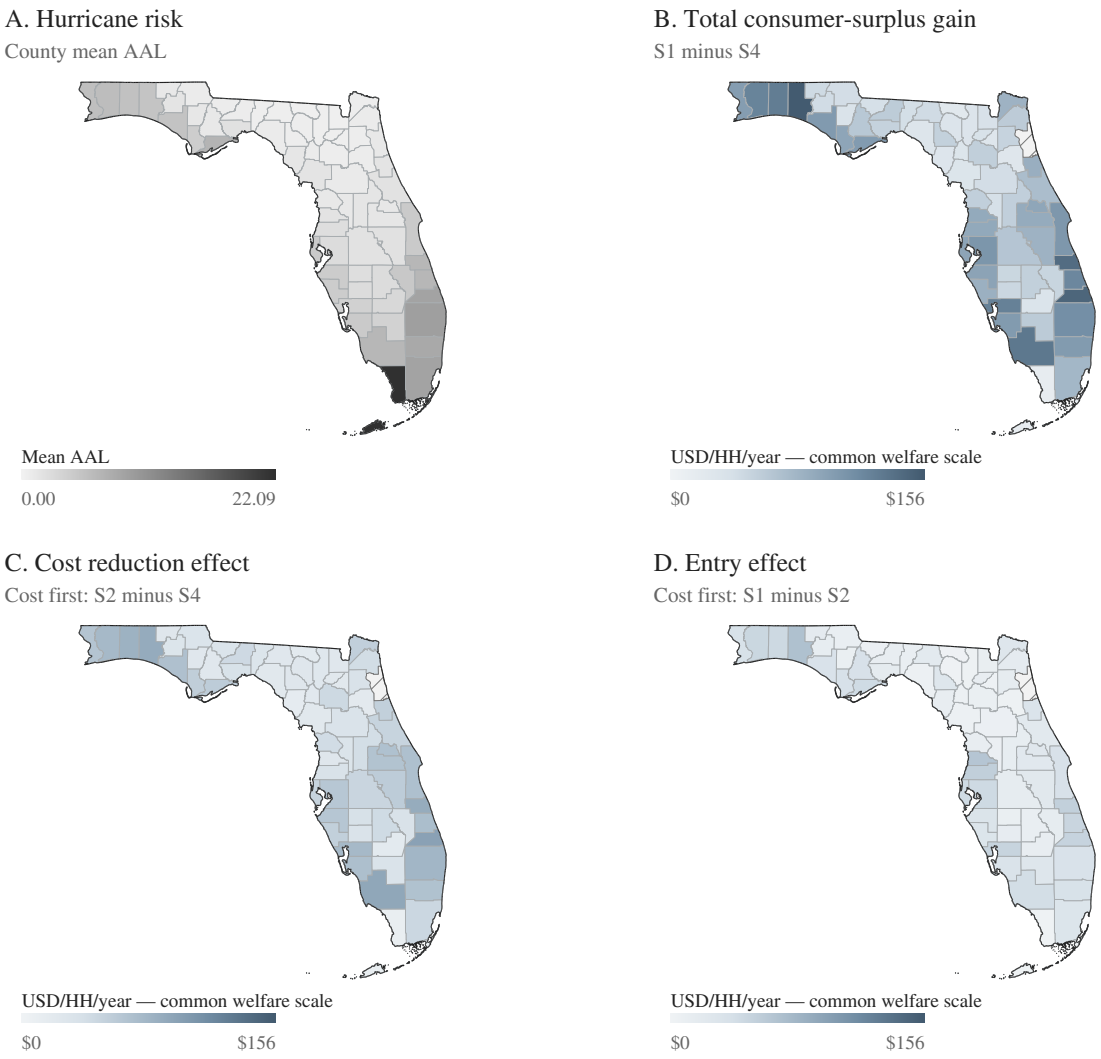
**Spatial Incidence.** We next examine how the welfare gain is distributed across Florida's counties. Figure 13 maps four quantities side by side: county risk exposure in panel (a), the aggregate welfare gain in panel (b), the welfare gain from the cost channel alone in panel (c), and the welfare gain from the entry channel in panel (d) (the difference between panels (b) and (c)).

The cost channel is especially large in Florida's coastal and high-premium markets. The proportional cost reduction is the same in every county, but marginal costs are about 35 percent higher in a 75th-percentile county than in a 25th-percentile county, so the dollar cost reduction and the premium response scale with catastrophe risk. Differences in dollar savings reflect baseline marginal costs, the presence and market shares of treated insurers, and equilibrium substitution among products.

The entry channel follows a slightly different spatial pattern. Its magnitude depends on the competitive importance of the additional products in each county. An additional product has a larger effect when it is a close substitute for incumbent offerings and generates a stronger pricing response. Thus, the importance of the expanded-product component depends on local market structure rather than catastrophe risk alone.

Taken together, the aggregate effect combines geographic variation in insurers' cost changes with variation in the competitive effect of additional products. We find that higher-risk areas, on average, benefit more from the collateral reform than lower-risk areas.

Figure 13: Spatial incidence of the reinsurance collateral reform




# 7 Conclusion

Reinsurance is a critical input to insuring homes against severe climate risks. While reinsurers diversify globally, they cannot fully smooth the large catastrophe risks in the U.S. market, and bearing those risks is capital-intensive. We study a regulatory reform that lowered the capital cost of accessing offshore reinsurance, exploiting the staggered adoption of the reinsurance collateral policy change across states. Because treatment is assigned by an insurer's state of domicile, we can compare otherwise similar insurers competing in the same markets but facing different effective costs of ceding catastrophe risk.

We find that the reform cut foreign reinsurance prices by about 30%, increased the quantity of foreign reinsurance ceded by 25%, and reduced concentration in the reinsurance market. These effects are concentrated among insurers with substantial hurricane-corridor exposure, consistent with a binding latent demand for catastrophe risk transfer. These reinsurance-market changes pass through to the primary market: in Florida, wind premiums fell roughly 10–16% in treated insurer-county-quarter cells, treated insurers gained county market share, and they became more likely to write wind-inclusive coverage. A structural model of the Florida wind market implies that the reform reduced treated insurers' marginal cost of supplying wind coverage by about 10%, with three-quarters of that reduction reaching homeowners as lower premiums. The aggregate consumer-surplus gain is roughly $95 per household per year, concentrated in the high-risk coastal counties where reinsurance is the largest component of premiums.

Two implications follow. First, the capital costs required to hold catastrophe risk, not just the expected losses themselves, are a first-order driver of homeowners insurance prices in disaster-prone markets. Policies that reduce those capital costs, whether through collateral reform, alternative-capital vehicles, or public reinsurance of last resort, can deliver meaningful affordability gains and should be evaluated against the cost-of-capital channel we identify, not only against the expected-loss channel that has dominated the policy discussion. Second, the pass-through rate we recover is itself useful beyond this particular reform: it characterizes how any future shock to the cost of bearing catastrophe risk, including new entrants and risk-transfer technologies, should translate into household premiums. Our estimates imply that most of any such cost reduction reaches consumers rather than being absorbed as rents upstream.

Several limitations qualify these results. Our structural welfare estimates are identified from Florida and from variation generated by the reinsurance collateral policy change, and external validity to other catastrophe-exposed states or to alternative cost-of-capital interventions is not guaranteed. Our counterfactuals also hold the set of insurers operating in each county-quarter fixed, so the reported gains likely understate the full benefit of the reform once entry is allowed to respond. Finally, we drop Florida-domiciled insurers from the QUASR analysis, since Florida adopted in 2009 before our QUASR window begins; because these insurers are the most reinsurance-reliant in the state, our Florida estimates are likely conservative.

Several extensions are already in progress. We are microfounding the marginal-cost function by replacing the reduced-form post-reform indicator with a structural reinsurance loading instrumented by the reinsurance collateral policy change, which generates testable cross-sectional restrictions on the cost-of-capital interpretation. We are also reconciling the reduced-form and structural magnitudes through a sequence of partial-equilibrium counterfactuals that incremen-

tally activate the cost reduction, markup re-equilibration, and post-reform product composition, and decomposing the estimated pass-through into demand-curvature, concentration, and multi-product internalization components.

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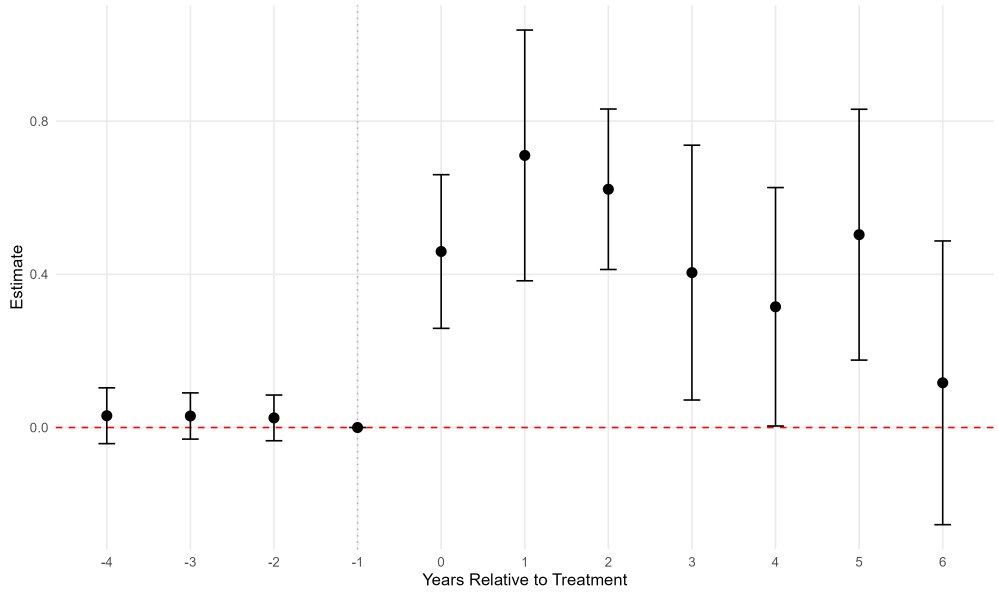
# A Additional National Reinsurance Market Results

## A.1 Domestic Reinsurer Results

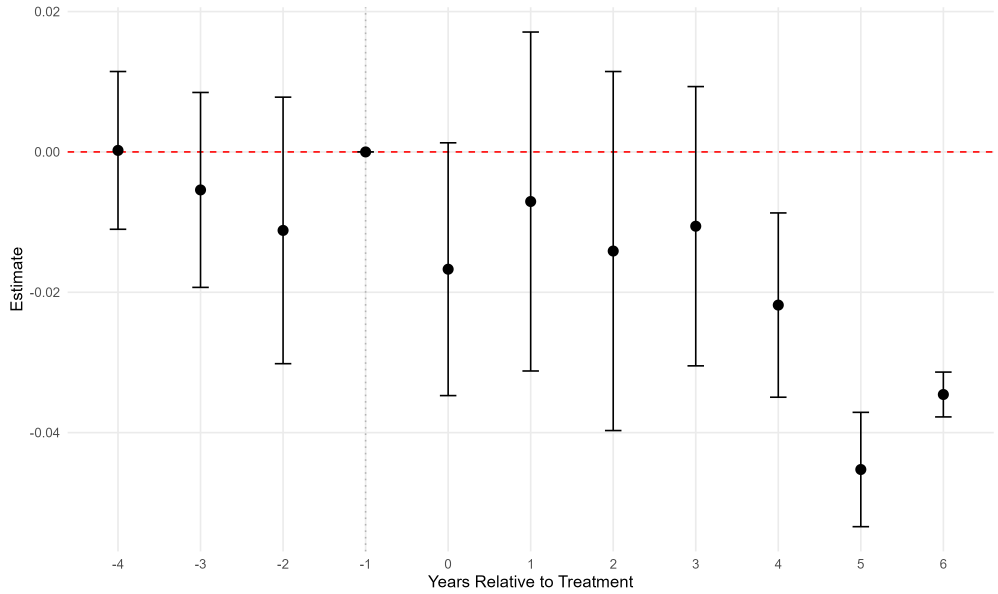
This appendix presents the national reinsurance market results for domestic reinsurers, complementing the foreign reinsurer results in Section 3.2.1.

Figure 14: Quantity of Reinsurance Ceded: Domestic Reinsurers




*Notes:* The figure reports a slope specification event study estimated on an insurer-by-year balanced panel restricted to insurer DPW $> 0$. The outcome is dollars of reinsurance ceded to domestic reinsurers; the exposure variable is direct premiums written (DPW). Event time $\tau$ is relative to the insurer's domicile state adoption. Coefficients $\{\gamma_\tau\}$ are from the two-stage specification in equations (5) and (6), with reference period $\tau = -1$. The event window is $\tau \in [-5, 6]$ with observations beyond $\tau = 5$ binned to $\tau = 6$. Estimation uses insurer DPW weights and clusters standard errors by state of domicile. State-of-domicile and year fixed effects are included in Stage 1.

Figure 15: Reinsurance Prices: Domestic Reinsurers

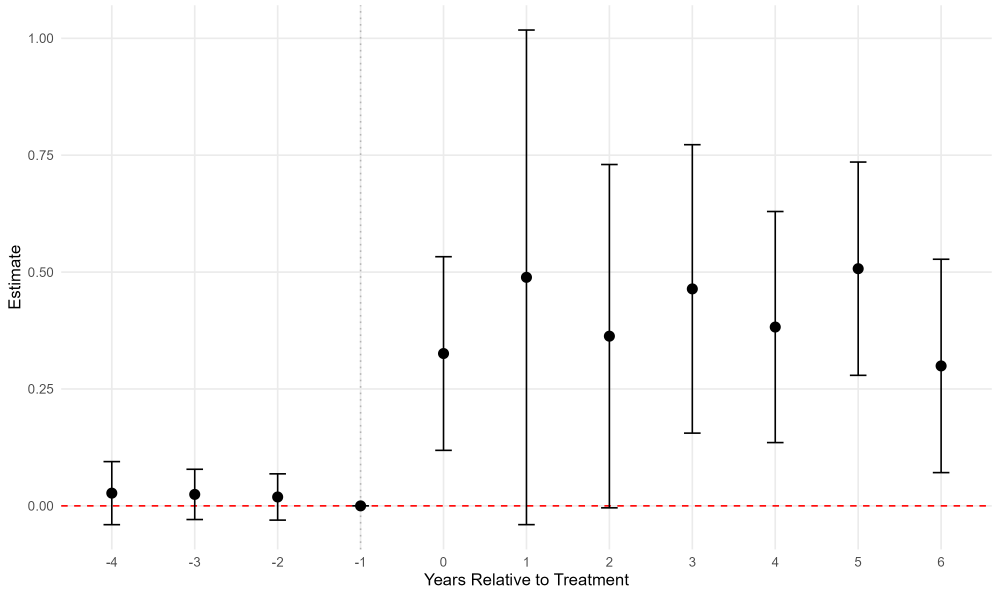



*Notes:* The figure reports a price slope specification event study estimated on an insurer-by-year balanced panel. The outcome is losses plus loss adjustment expenses (LAE) incurred on reinsurance ceded to domestic reinsurers; the exposure variable is premiums ceded to domestic reinsurers. The sample is restricted to insurer-years with positive ceded premiums to domestic reinsurers. Event time $\tau$ is relative to the insurer's domicile state adoption. Coefficients $\{-\gamma_\tau\}$ (sign-flipped) are from the two-stage specification in equations (5) and (6), with reference period $\tau = -1$. Negative coefficients indicate lower loss ratios (cheaper reinsurance) post-treatment. The event window is $\tau \in [-4, 6]$ with observations beyond $\tau = 5$ binned to $\tau = 6$. Estimation uses insurer DPW weights and clusters standard errors by state of domicile. State-of-domicile and year fixed effects are included in Stage 1.

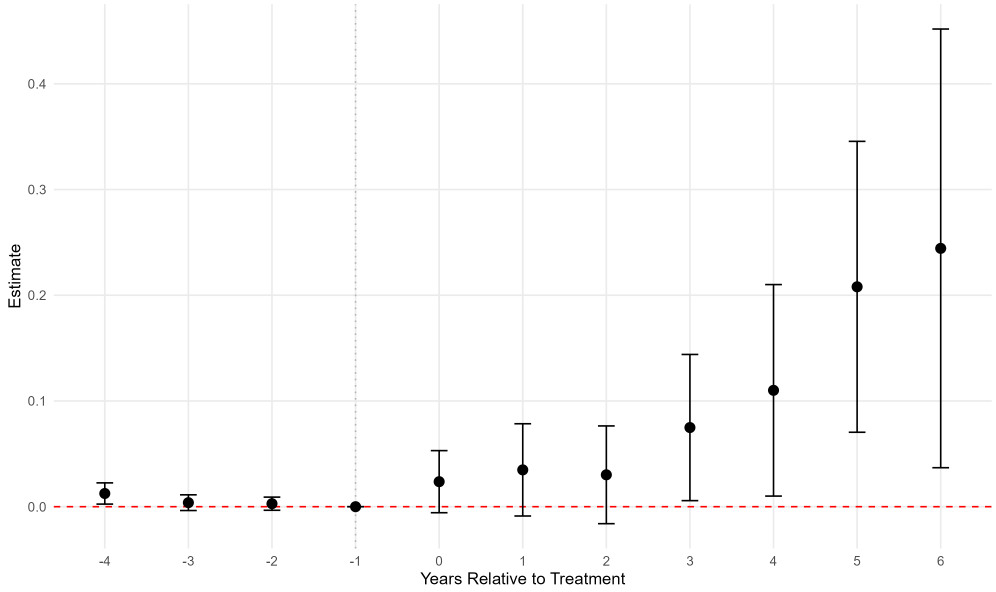
## A.2 Affiliated Reinsurer Decomposition

To understand the economic nature of the reinsurance quantity effects documented in Section 3.2.1, we decompose reinsurance flows by whether the reinsurer is affiliated (a captive entity created primarily to reinsure its parent or affiliated companies) versus non-affiliated (a traditional commercial reinsurer). Figure 16 reveals that the increase in reinsurance flows entirely to non-affiliated reinsurers. Ceding to affiliated reinsurers actually declined modestly post-reform, suggesting the regulatory change facilitated access to arm's-length commercial reinsurance markets rather than within-group risk transfer arrangements.

Figure 16: Quantity of Reinsurance Ceded: Affiliated vs. Non-Affiliated Reinsurers




(a) Premiums ceded to affiliated reinsurers as a function of DPW

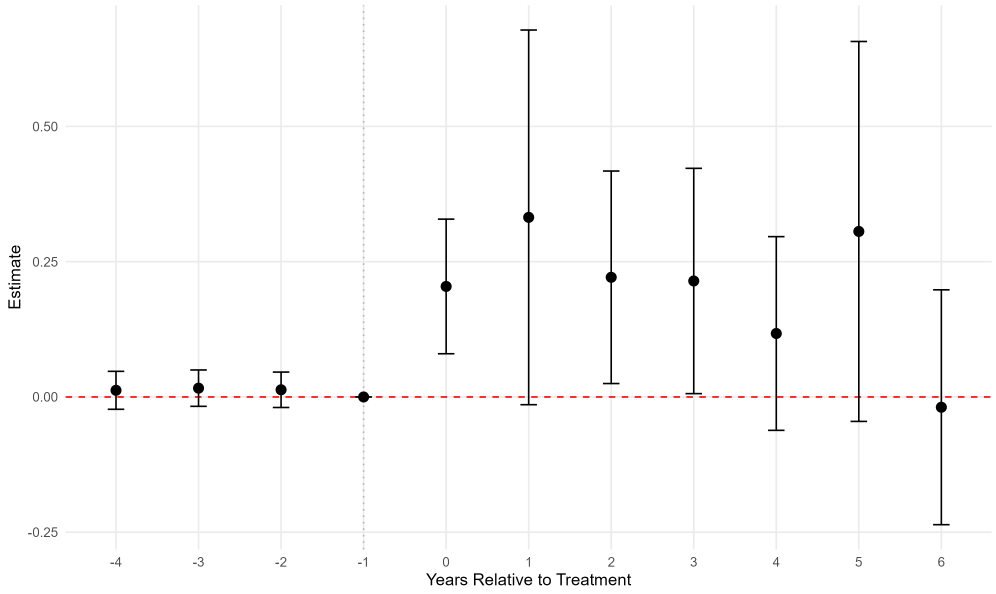



(b) Premiums ceded to non-affiliated reinsurers as a function of DPW

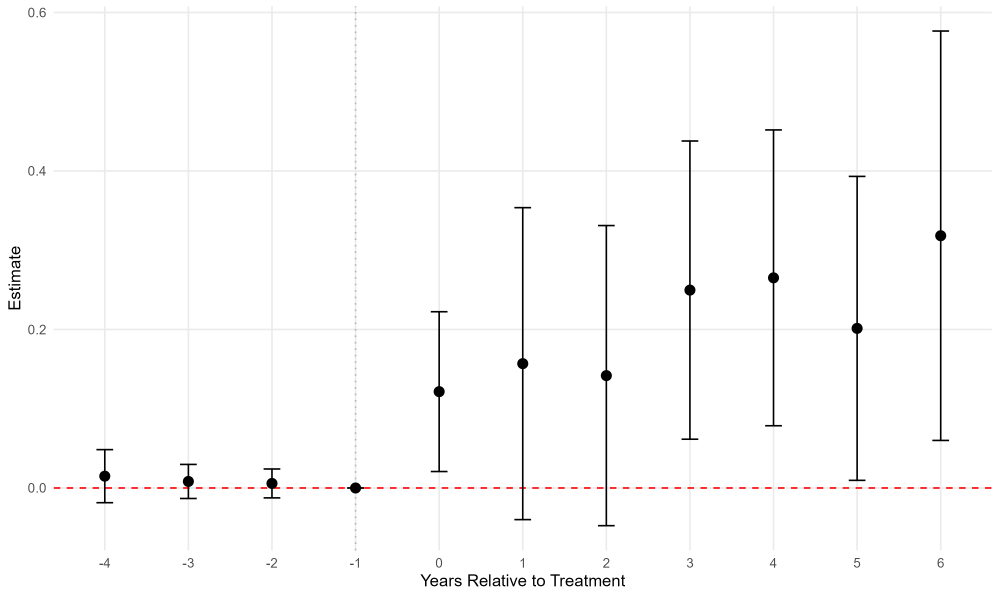
*Notes:* Each panel reports a slope specification event study estimated on an insurer-by-year balanced panel restricted to insurer DPW $> 0$. The outcome is dollars of reinsurance ceded to affiliated (panel a) or non-affiliated (panel b) reinsurers; the exposure variable is direct premiums written (DPW). Affiliated reinsurers are entities created primarily to reinsure affiliated companies; non-affiliated reinsurers are traditional commercial reinsurance companies. Event time $\tau$ is relative to the insurer's domicile state adoption. Coefficients $\{\gamma_\tau\}$ are from the two-stage specification in equations (5) and (6), with reference period $\tau = -1$. The event window is $\tau \in [-5, 6]$ with observations beyond $\tau = 5$ binned to $\tau = 6$. Estimation uses insurer DPW weights and clusters standard errors by state of domicile. State-of-domicile and year fixed effects are included in Stage 1.

Within affiliated reinsurance, we further examine whether the dollars ultimately end up onshore or offshore.

Figure 17: Quantity of Reinsurance Ceded: Affiliated Onshore vs. Offshore Reinsurers




(a) Premiums ceded to onshore affiliated reinsurers as a function of DPW




(b) Premiums ceded to offshore affiliated reinsurers as a function of DPW

*Notes:* Each panel reports a slope specification event study estimated on an insurer-by-year balanced panel restricted to insurer DPW $> 0$. The outcome is dollars of reinsurance ceded to affiliated reinsurers whose group is entirely domestic (panel a) or affiliated reinsurers whose group has some foreign element (panel b); the exposure variable is direct premiums written (DPW). Event time $\tau$ is relative to the insurer's domicile state adoption. Coefficients $\{\gamma_\tau\}$ are from the two-stage specification in equations (5) and (6), with reference period $\tau = -1$. The event window is $\tau \in [-5, 6]$ with observations beyond $\tau = 5$ binned to $\tau = 6$. Estimation uses insurer DPW weights and clusters standard errors by state of domicile. State-of-domicile and year fixed effects are included in Stage 1.

## A.3 Expected Reinsurer Impairment Losses

This appendix details the construction of the model-implied expected impairment loss in Section 3.2.5 and reports the pooled estimates and event studies.

**Construction.** For each cedent–reinsurer pair we observe the recoverable $R_{ijt}$ in Schedule F. We map each reinsurer's financial-strength rating to a one-year impairment probability $p(r_j^*)$ using the A.M. Best impairment table. We impute the uncollateralized share $\lambda_{ijt}$ from the statutory collateral schedule in every pair-year, including years with reported collateral. We set $\lambda = 1$ for domestic and U.S.-authorized reinsurers. Non-admitted foreign reinsurers have $\lambda = 0$ before the cedent's domicile state adopts Model Law 785. After adoption, rated foreign reinsurers receive the uncollateralized share implied by the statutory schedule for their rating tier. Unrated foreign reinsurers remain at $\lambda = 0$. We use reported post-2018 collateral only to assess this imputation.

Expected impairment loss is $\mathrm{EL}_{it} = \mathrm{LGD} \sum_j p(r_j^*)\lambda_{ijt}R_{ijt}$, with $\mathrm{LGD} = 1 - \rho$ and a baseline recovery rate $\rho = 0.50$. About $49\%$ of the unaffiliated recoverable carries a rating; rather than treat the remainder as default-free, the baseline assigns unrated reinsurers the median rated-tier impairment probability ($1.17\%$). Because ratings are observed only at end of sample, we apply each reinsurer's end-of-sample rating $r_j^*$ to all years, so the exercise is a calibration of credit exposure rather than a direct estimate of changing reinsurer solvency. The sample is unaffiliated third-party property-casualty reinsurance, $2{,}015$ cedents over 2004–2023, estimated with the imputation difference-in-differences estimator (cedent and year fixed effects) and standard errors clustered on cedent domicile state ($50$ clusters).

**Pooled estimates.** Table 6 reports the pooled treatment effects for expected impairment loss (Panel A) and for the premium-weighted impairment probability (Panel B).

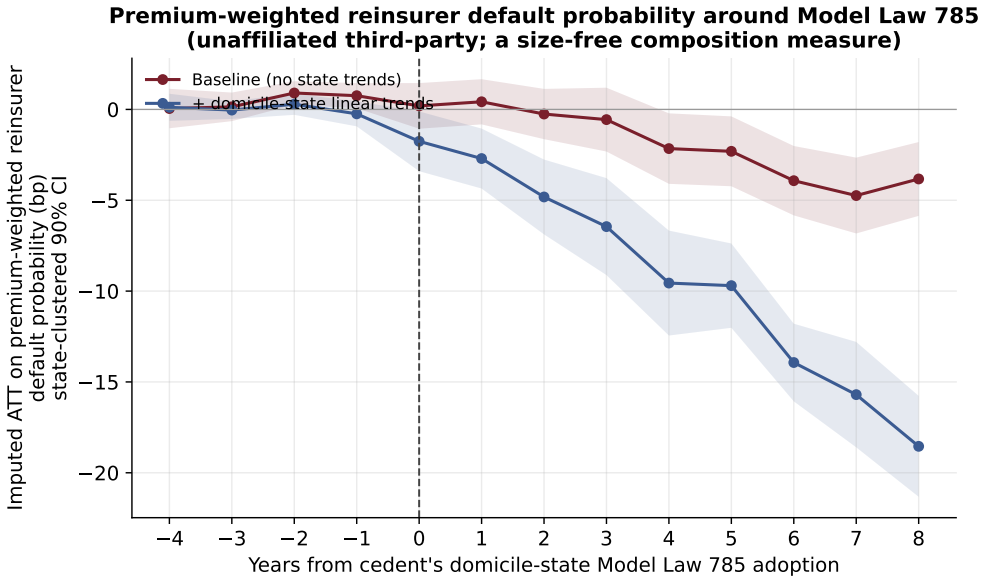
Table 6: Effect of Model Law 785 on cedents' reinsurer impairment exposure

| Specification | ATT | s.e. | $t$ | $p$ |
|---|---|---|---|---|
| *Panel A. Expected impairment loss ($/cedent/year)* | | | | |
| Estimate | $+57{,}584^{***}$ | $19{,}693$ | $2.92$ | $0.003$ |
| *Panel B. Premium-weighted impairment probability (bp)* | | | | |
| Estimate | $-2.14$ | $0.81$ | $-2.63$ | $0.009$ |

*Notes:* Imputed difference-in-differences, cedent and year fixed effects, standard errors clustered on cedent domicile state ($50$ clusters); unaffiliated third-party reinsurance. Baseline calibration: $50\%$ recovery, statutory collateral, unrated reinsurers at the median rated-tier impairment probability ($1.17\%$; pre-reform mean premium-weighted probability $47$ bp). $^{***}$ denotes statistical significance at the $1\%$ level for the Panel A dollar estimate.

**Event studies.** Figure 18 reports the event study for the premium-weighted impairment probability, which declines after adoption. The share of premium ceded to unrated reinsurers also falls. Cedents therefore move toward rated counterparties. The baseline dollar expected-loss estimate is positive and statistically significant at the $1\%$ level because the larger unsecured exposure more than offsets the safer counterparty composition.

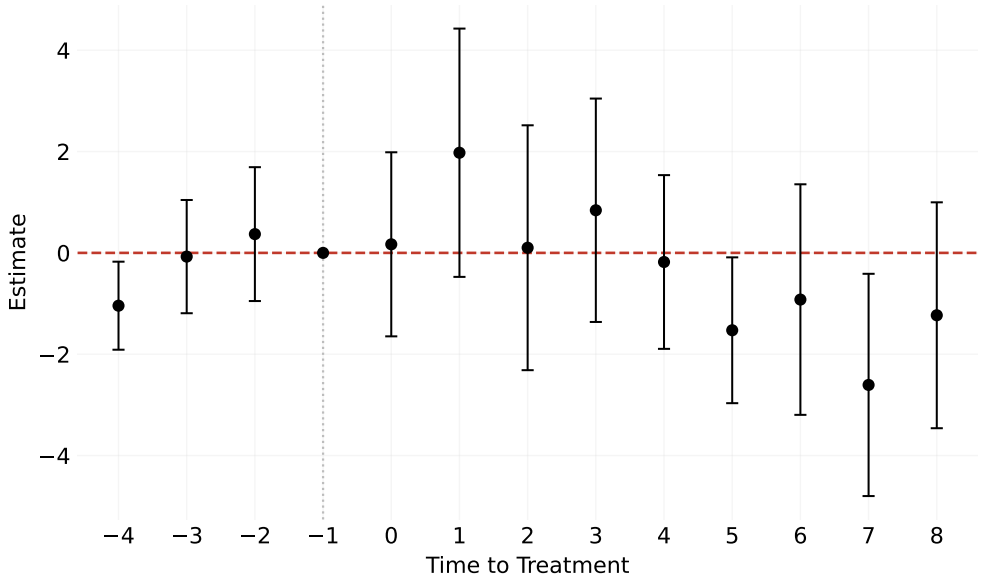
Figure 18: Premium-Weighted Reinsurer Impairment Probability




*Notes:* Event study of the premium-weighted mean reinsurer impairment probability (basis points), a size-free composition measure; unrated reinsurers at the median rated-tier rate. Bands are 90% confidence intervals, clustered on cedent domicile state.

**Riskiness per dollar of reinsurance.** To separate changes in book size from risk per dollar, we scale expected impairment loss by the cedent's total reinsurance premium ceded and re-estimate. The pooled estimate is $-\$0.21$ per \$1,000 of premium ceded and is imprecise. The data therefore do not show that each dollar of reinsurance became riskier (Figure 19).

Figure 19: Expected Impairment Loss per Dollar of Reinsurance Premium Ceded

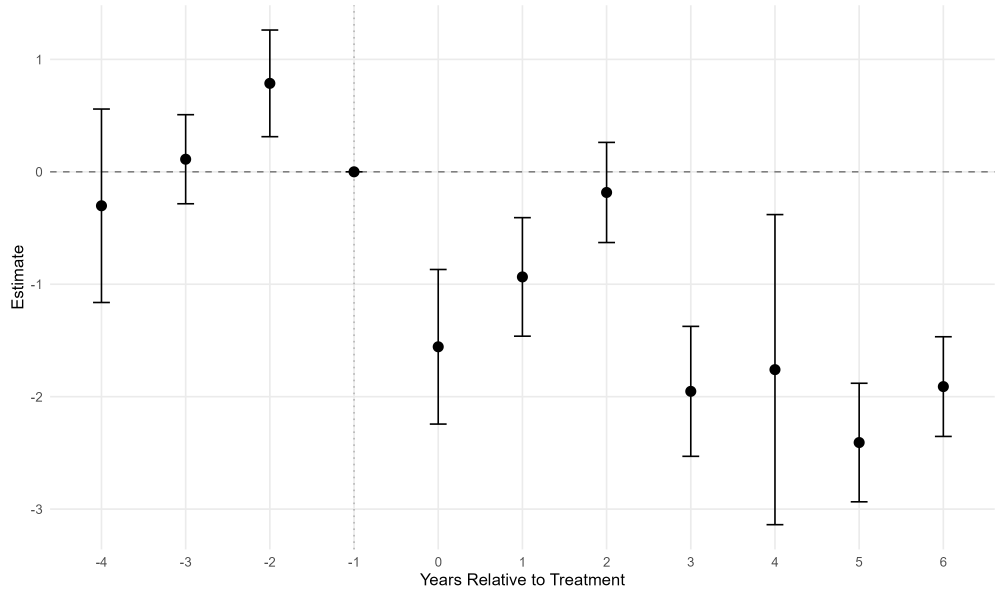



*Notes:* The figure reports an event study where the outcome is the model-implied expected impairment loss per dollar of total reinsurance premium ceded, displayed per \$1,000 of premium—a size-free riskiness-per-dollar measure. Event time $\tau$ is relative to the cedent's domicile-state adoption of Model Law 785, with reference period $\tau = -1$; the imputation estimator, sample, and calibration match Figure 8. Standard errors are clustered on cedent domicile state, and bands are 90% confidence intervals.

## A.4 Florida Reinsurance Prices: Second-Event Covers Retained

The specification in Section 3.2.3 excludes second-event, top-and-drop and drop-down layers, whose modeled expected losses are small relative to premium. Figure 20 repeats that specification with these layers retained.

Figure 20: Reinsurance Prices: Second-Event Covers Retained




*Notes:* Event study of the carrier-year reinsurance price multiple, estimated as in Figure 6 but retaining second-event, top-and-drop and drop-down layers. The pooled post-adoption effect is $-1.82$ (s.e.: 0.17). The sample is 518 carrier-years, 100 carriers, and 16 domicile states. Standard errors are clustered by state of domicile, and bars are 95% confidence intervals.

# B Reinsurance Collateral by Authorization Class

This appendix documents the collateral magnitudes underlying the cost-of-capital benchmark in the main text, for foreign unaffiliated reinsurers—the population targeted by the credit-for-reinsurance reforms (non-NAIC-licensed reinsurers with an unaffiliated Schedule F reinsurance class). Using NAIC Schedule F, we measure recognized credit-risk collateral—multiple-beneficiary trusts plus letters of credit—against ceded premium and against total recoverables, by authorization class. Table 7 reports 2018, the first year the collateral schedule is disclosed and the closest available snapshot to the pre-reform regime; Table 8 reports 2024, the most recent year.

Collateral on the unauthorized book—the regime the reform relieves—is first order: unauthorized reinsurers post collateral equal to 87% of premium (73% of recoverables) in 2018 and 98% (80%) in 2024. The reform ladder moves qualifying reinsurers toward certified status, where collateral falls well below the unauthorized requirement, and ultimately toward reciprocal-jurisdiction status at zero collateral (omitted from the tables). The recognized collateral measured here excludes funds held and other allowable offsets and so understates total posted security; it is the conservative basis for the benchmark.

Table 7: Reinsurance collateralization by authorization class (foreign unaffiliated reinsurers, 2018)

| Class | Premium ($B) | Collateral ($B) | Recoverables ($B) | Coll./Prem. | Coll./Recov. |
|---|---|---|---|---|---|
| Authorized | 6.5 | 2.0 | 12.2 | 31% | 16% |
| Certified | 1.5 | 1.2 | 2.7 | 79% | 44% |
| Mixed Filing | 9.6 | 6.7 | 18.7 | 70% | 36% |
| Unauthorized | 28.7 | 25.1 | 34.4 | 87% | 73% |
| Total | 46.2 | 34.9 | 68.0 | 76% | 51% |

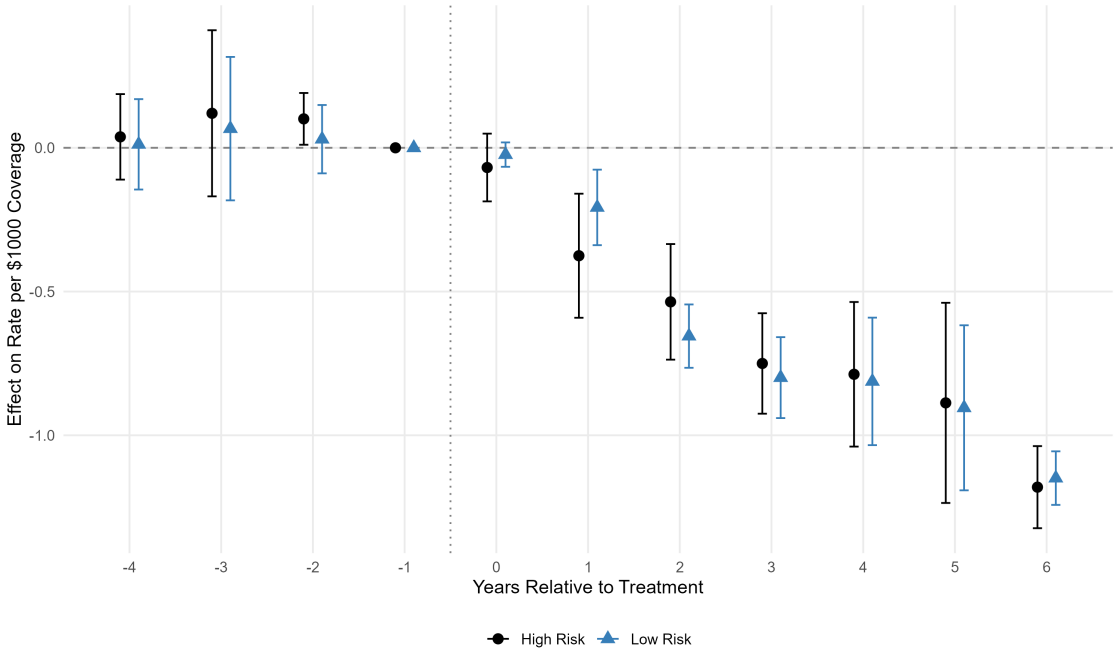
*Note:*
Collateral is recognized credit-risk collateral (multiple-beneficiary trusts plus letters of credit); it excludes funds held and other allowable offsets. Foreign = non-NAIC-licensed reinsurer; unaffiliated per the Schedule F reinsurance class. The authorization label understates the certified book, much of which files as "Mixed Filing." Source: NAIC Schedule F, statement year 2024. Dollar amounts in $ billions.

Table 8: Reinsurance collateralization by authorization class (foreign unaffiliated reinsurers, 2024)

| Class | Premium ($B) | Collateral ($B) | Recoverables ($B) | Coll./Prem. | Coll./Recov. |
|---|---|---|---|---|---|
| Authorized | 11.1 | 1.8 | 15.7 | 16% | 12% |
| Certified | 1.6 | 0.4 | 2.4 | 27% | 17% |
| Mixed Filing | 29.3 | 15.8 | 45.3 | 54% | 35% |
| Unauthorized | 55.0 | 54.1 | 67.2 | 98% | 80% |
| Total | 97.0 | 72.1 | 130.6 | 74% | 55% |

*Note:*
Collateral is recognized credit-risk collateral (multiple-beneficiary trusts plus letters of credit); it excludes funds held and other allowable offsets. Foreign = non-NAIC-licensed reinsurer; unaffiliated per the Schedule F reinsurance class. The authorization label understates the certified book, much of which files as "Mixed Filing." Source: NAIC Schedule F, statement year 2024. Dollar amounts in $ billions.

Figure 21: Policy Pricing: High vs. Low Risk

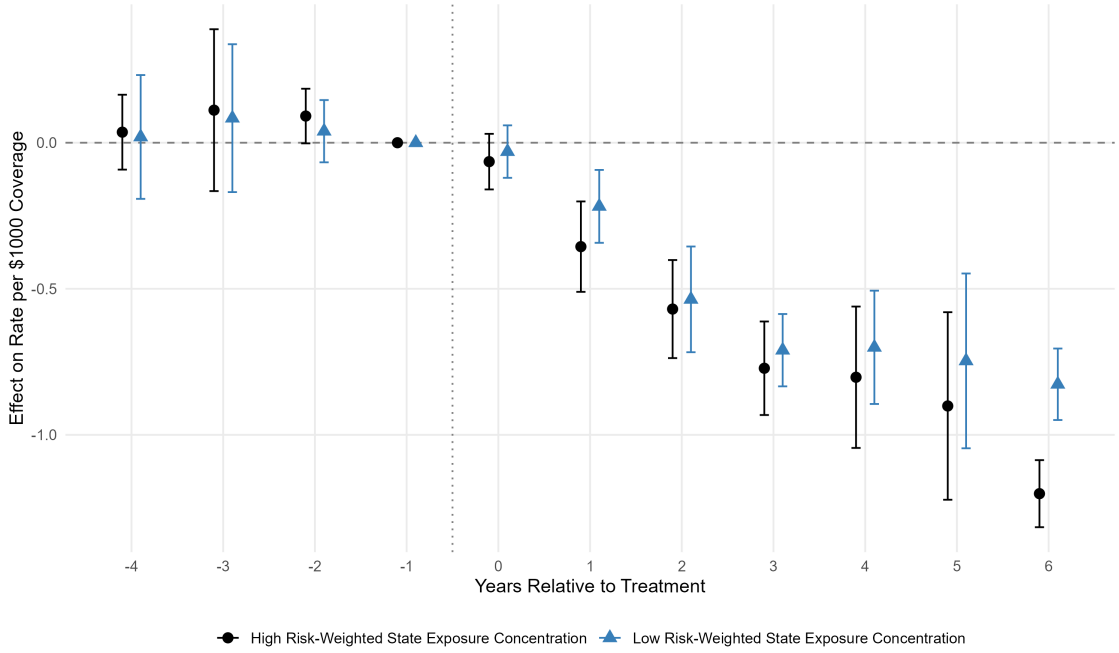



# C Additional Florida Results

## C.1 Hurricane Exposure Heterogeneity

This appendix presents additional Florida results examining heterogeneity in the effects of the reinsurance collateral policy change on policy pricing across risk categories (for this raw-rate outcome we drop observations above the 99th percentile of the rate), complementing the main wind premium results in Section 4.2.1.

Figure 22: Policy Pricing: High vs. Low Expected Loss Counties




## C.2 Wind vs. Non-Wind Price Effects

As a placebo for the wind triple difference, we re-estimate the same event study separately on the wind and non-wind rate levels: if the reinsurance cost relief operates through catastrophe coverage, the wind rate should fall sharply while the far-less-reinsurance-intensive non-wind rate should move less. Figure 23 bears this out. The wind rate declines steadily after adoption, whereas the non-wind rate is flat for the first two years and ultimately falls by only about a third as much.

## C.3 Price Effects: Incumbents versus Entrants

A natural concern is that the wind-rate decline reflects the entry of new insurers writing low-priced wind policies rather than repricing by established firms. We test this by restricting to incumbents: those present in a county in the four years prior to treatment. The results, in Figure 24 below, show that if anything the price effects are stronger among incumbents.

Figure 23: Placebo: Wind vs. Non-Wind Rate Levels

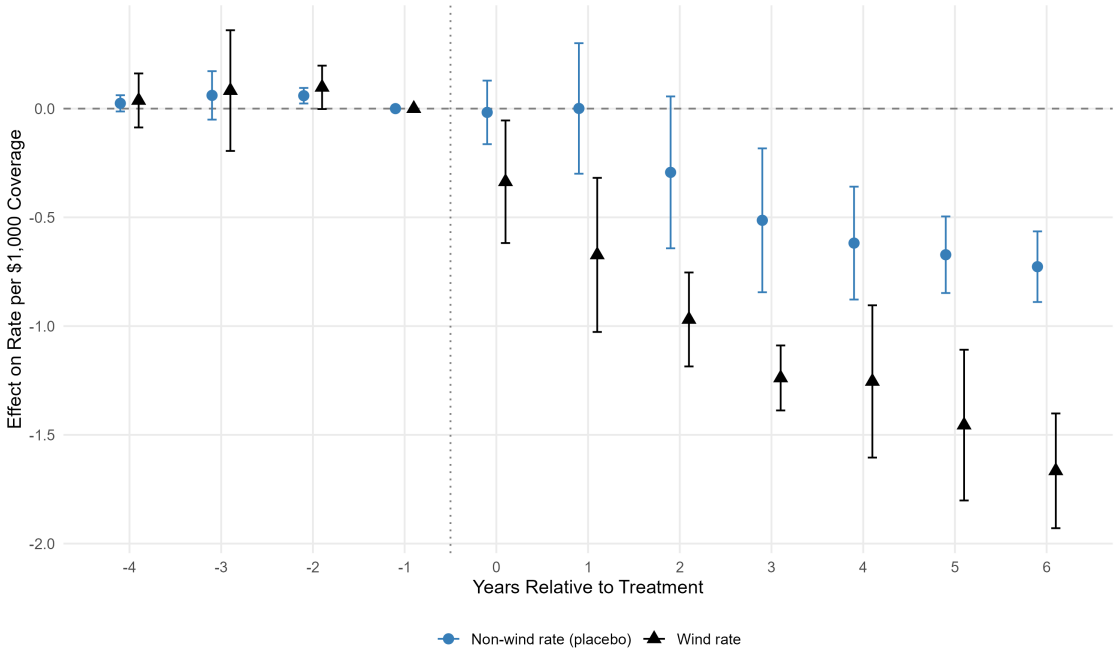
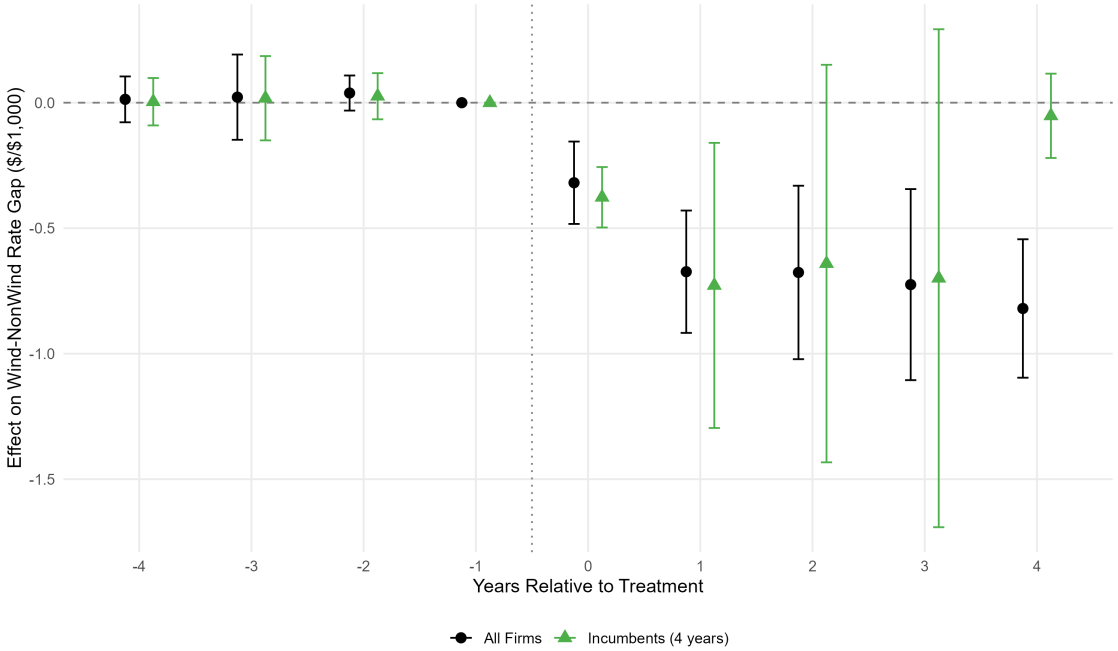



Figure 24: Wind Triple Difference: Incumbents vs. All Firms

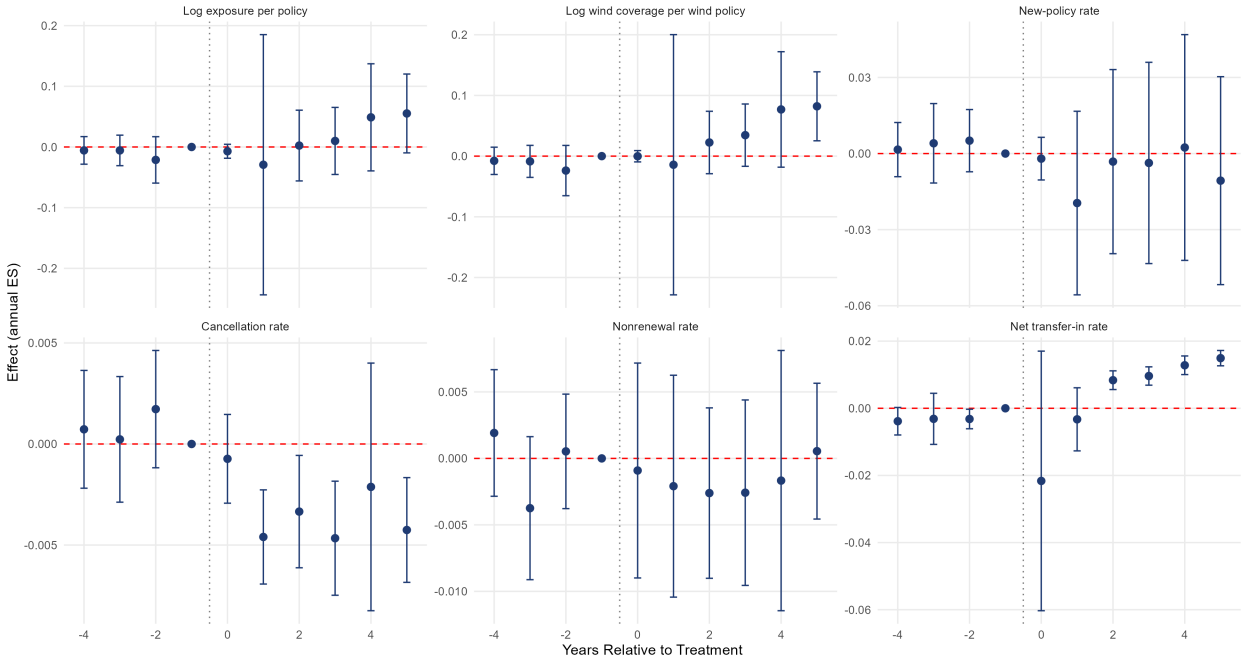


## C.4 Unchanged Policy Composition

A concern is that our price results in Florida represent a change in the composition of the risks being insured. For example, insurers could be writing safer risks or risks with a smaller wind component. This would generate the patterns in Figure 9 in the main text.

To test this, we study whether the risk (exposure per policy), wind exposure, or flow of new business changes with the reform. Figure 25 shows they do not. This confirms that our results are driven by lower prices on a comparable existing book of business, not a shift toward a structurally different set of risks.

Figure 25: Compositional-Change Robustness: Coverage and Policy-Flow Outcomes

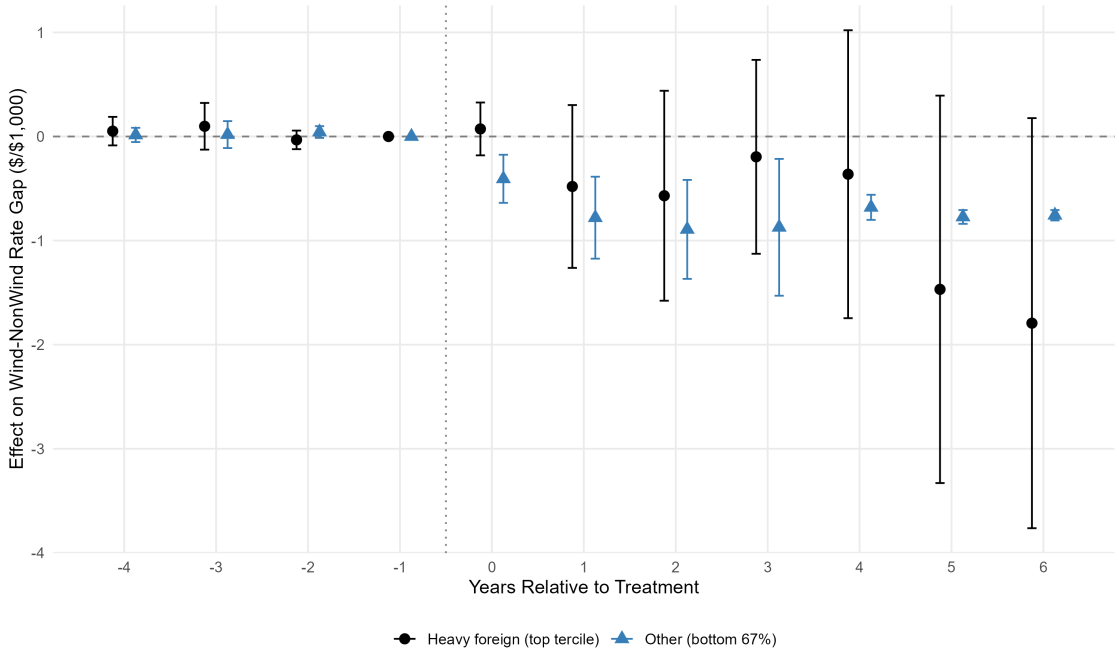



## C.5 Heterogeneity by Pre-treatment Exposure to Foreign Reinsurance

To probe the mechanism directly, we link each insurer to its pre-reform reliance on foreign reinsurance. We compute each insurer's 2008–2010 share of reinsurance premium ceded to foreign reinsurers. We re-estimate the headline price specification splitting insurers on that measure into the top tercile versus the bottom two terciles. Figure 26 shows the wind-rate decline concentrates

among heavy foreign-reinsurance users, consistent with the collateral reform specifically targeting the cost of foreign reinsurance.

Figure 26: Wind-Rate Effect by Pre-Reform Foreign-Reinsurance Exposure




# D Model Law 785 Treatment Timing Details

Table 9: NAIC Model Law 785 Effective Adoption Dates by State

| State | Effective Date | Citation |
|---|---|---|
| Florida | Jan 1, 2009 | Florida Administrative Code Rule 69O-144.007 "Credit for Reinsurance from Certified Reinsurers." |
| New York | Jan 1, 2011 | https://www.law.cornell.edu/regulations/new-york/11-NYCRR-125.4 |
| Indiana | Apr 6, 2011 | https://www.law.cornell.edu/regulations/indiana/760-IAC-1-56-7.5 |
| New Jersey | Jun 20, 2011 | https://www.law.cornell.edu/regulations/new-jersey/N-J-A-C-11-2-28-3 |

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| State | Effective Date | Citation |
|---|---|---|
| Connecticut | Oct 1, 2012 | Public Act 12-139 |
| California | Jan 1, 2013 | SB 1216 |
| Virginia | Jan 1, 2013 | https://law.lis.virginia.gov/admincode/title14/agency5/chapter300/section95/ |
| Georgia | Jun 12, 2013 | Rule 120-2-78-.05. Credit for Reinsurance Accredited Reinsurers |
| Pennsylvania | Jun 24, 2013 | https://www.pacodeandbulletin.gov/Display/pabull?file=%2Fsecure%2Fpabulletin%2Fdata%2Fvol43%2F43-21%2F950.html |
| Louisiana | Jul 1, 2013 | https://www.law.cornell.edu/regulations/louisiana/La-Admin-Code-tit-37-SS-XIII-3510 |
| Delaware | Aug 11, 2013 | 18 DE Admin. Code 1003 "Credit for Reinsurance" |
| Iowa | Oct 30, 2013 | https://www.legis.iowa.gov/docs/aco/arc/1111C.pdf |
| Alabama | Jan 1, 2014 | Ala. Admin. Code 482-1-156 |
| Missouri | Jan 1, 2014 | https://www.law.cornell.edu/regulations/missouri/20-CSR-200-2-100 |
| New Hampshire | Jan 1, 2014 | https://isiteplus.naic.org/crin-report/crinReportPublic.xhtml |
| Maryland | Aug 18, 2014 | https://www.law.cornell.edu/regulations/maryland/COMAR-31-05-08-24 |
| Rhode Island | Sep 2, 2014 | https://risos-apa-production-public.s3.amazonaws.com/DBR/7794.pdf |
| Colorado | Jan 1, 2015 | https://law.justia.com/codes/colorado/title-10/regulation-of-insurance-companies/article-3/part-7/section-10-3-701/ |
| Hawaii | Jan 1, 2015 | Act 234, Session Laws of Hawaii 2014 |
| Arkansas | Jul 4, 2015 | SB881 |
| Vermont | Aug 27, 2015 | https://www.law.cornell.edu/regulations/vermont/21-038-Code-Vt-R-21-020-038-X |
| Arizona | Oct 23, 2015 | Ariz. Admin. Register Vol. 21, Issue 43 |
| Ohio | Nov 6, 2015 | https://dam.assets.ohio.gov/image/upload/insurance.ohio.gov/Company/Documents/ORA/Assuming_insurers%20A-5%20list.pdf |
| North Dakota | Jan 1, 2016 | https://www.insurance.nd.gov/companies/company-licensing/certified-reinsurers |

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| State | Effective Date | Citation |
|---|---|---|
| Washington | Jan 2, 2016 | https://lawfilesext.leg.wa.gov/law/wsr/2015/24/15-24-126.htm |
| Maine | Jan 24, 2016 | https://www.maine.gov/pfr/insurance/sites/maine.gov.pfr.insurance/files/inline-files/740-adopted-amendment.pdf |
| Nebraska | Feb 15, 2016 | https://doi.nebraska.gov/rules-regulations-and-guidance-document-index |
| Massachusetts | Oct 1, 2016 | https://www.law.cornell.edu/regulations/massachusetts/department-211-CMR/title-211-CMR-130.00 |
| Nevada | Nov 2, 2016 | https://regulations.justia.com/states/nevada/chapter-681a/certified-reinsurers/section-681a-440/ |
| Montana | Nov 26, 2016 | https://www.law.cornell.edu/regulations/montana/Mont-Admin-r-6.6.3848 |
| Utah | Jan 10, 2017 | https://www.law.cornell.edu/regulations/utah/Utah-Admin-Code-R590-173-8 |
| North Carolina | Jul 20, 2017 | https://lrs.sog.unc.edu/bill/naic-modelsorsa-credit-reinsurance-ab-new |
| Oklahoma | Sep 15, 2017 | https://www.law.cornell.edu/regulations/oklahoma/OAC-365-25-7-73 |
| South Dakota | Oct 23, 2017 | https://dlr.sd.gov/insurance/companies/reinsurers.aspx |
| Texas | Jan 1, 2018 | https://www.legis.state.tx.us/tlodocs/85R/billtext/html/SB01070F.HTM |
| Wisconsin | Jan 1, 2018 | https://oci.wi.gov/Pages/Companies/CertifiedReinsurers.aspx |
| Wyoming | Feb 21, 2018 | https://doi.wyo.gov/companies/certified-reinsurers |
| Idaho | Mar 23, 2018 | IDAPA 18.01.75 "Credit for Reinsurance Rules" |
| Minnesota | May 10, 2018 | https://www.revisor.mn.gov/laws/2017/0/Session+Law/Chapter/6/ |
| South Carolina | May 25, 2018 | https://services.statescape.com/ssu/Regs/ss_8586743245995961190.pdf |
| Tennessee | May 31, 2018 | https://publications.tnsosfiles.com/rules/0780/0780-01/0780-01-63.20180531.pdf |
| New Mexico | Jul 24, 2018 | https://www.srca.nm.gov/wp-content/uploads/attachments/13.002.0008.pdf |
| Kentucky | Jan 1, 2019 | https://apps.legislature.ky.gov/record/18rs/hb464.html |

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| State | Effective Date | Citation |
|---|---|---|
| Michigan | Jan 2, 2019 | https://www.michigan.gov/difs/industry/insurance/certified-reinsurers/certified-reinsurers |
| Illinois | Nov 19, 2019 | https://www.law.cornell.edu/regulations/illinois/Ill-Admin-Code-tit-50-SS-1104.45 |
| Alaska | Dec 26, 2019 | https://www.commerce.alaska.gov/web/Portals/11/Pub/RA19-02.pdf |
| Oregon | Jan 1, 2020 | https://www.law.cornell.edu/regulations/oregon/Or-Admin-Code-SS-836-012-0046 |
| West Virginia | Jan 1, 2020 | https://apps.sos.wv.gov/adlaw/csr/ruleview.aspx?document=16966 |
| Mississippi | Jan 1, 2021 | https://www.mid.ms.gov/wp-content/uploads/2023/04/20201208ref.pdf |
| Kansas | Jul 1, 2021 | https://www.kslegislature.gov/li_2022/b2021_22/measures/documents/sb78_enrolled.pdf |