Capital Costs, Reinsurance, and the Price of Climate Risk

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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 model-implied price of reinsurance fell by a third. By reducing required capital, the reform also increased insurers' exposure to reinsurer non-payment risk. We show this dramatically impacted the primary insurance market in hurricane-exposed areas of Florida: additional insurers entered and the cost of insuring wind risk fell sharply. 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 14%. Counterfactual simulations imply consumer-surplus gains of $186 per household per year, split roughly equally across direct cost pass-through, strategic markup adjustment, and expanded product availability. 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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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 and earn returns elsewhere or 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. Because adoption was staggered across states and primary insurers are assigned to treatment through their domicile state, even when doing business in other states, we can compare reinsurance usage and, downstream, homeowner premiums and entry decisions across otherwise similar insurers exposed to the same risk but facing different reinsurance capital costs.

We first examine how the reform, which lowered the collateral requirements for foreign reinsurers, affected national reinsurance markets. 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. Using proxies for reinsurance prices – reinsurance premiums relative to realized reinsurance claims payments – we estimate that limiting collateral requirements cuts foreign reinsurance prices roughly in half.

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 property against severe storms such as hurricanes and is separately priced in Florida rather than bundled with traditional homeowners insurance. 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 rate by about $0.50 per $1,000 of insured exposure, which is roughly 12% of the pre-treatment average. 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 14%, close to the 12% decline estimated in the reduced-form exercises. Counterfactual simulations imply that the reform raises consumer surplus by about $186 per household per year, or roughly $912 million annually across Florida homeowners. These gains operate through three channels, of roughly equal importance: direct pass-through of lower marginal costs, strategic markup adjustment as insurers compete and re-price, and expanded product availability. The cost-side benefits are greater in high-risk coastal counties where reinsurance constitutes the largest component 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) documents a rapid rise in U.S. homeowners' insurance premiums since 2017, especially in high disaster risk areas, and cites increasing reinsurance costs as a key explanation. Oh et al. (2026) studies how state-level premium rate regulation can lead to cross-subsidization between states. Boomhower et al. (2024) illustrates 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) shows 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) contrasts 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 reduction 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

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

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

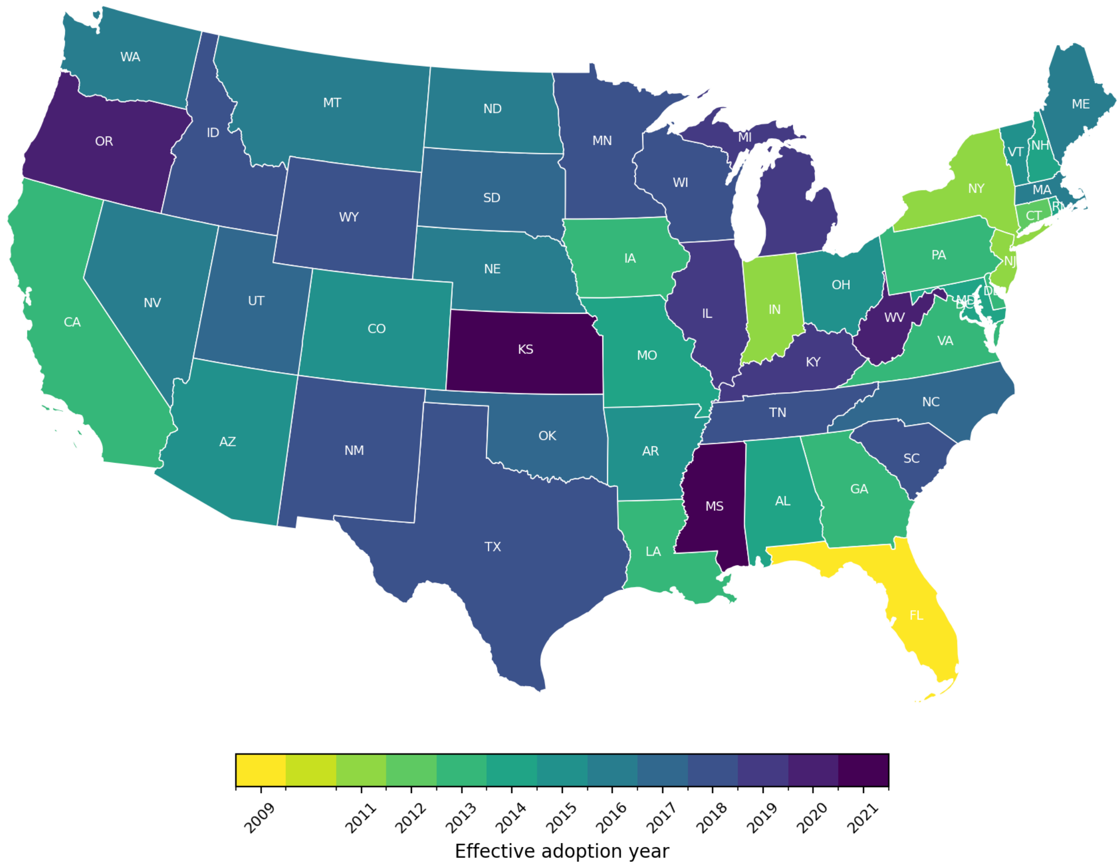
Insurance in the United States is regulated at the state level. State insurance departments oversee prices and policy forms, market conduct, and the capital adequacy of insurers domiciled in the state. Reinsurance is regulated as part of this framework, primarily through its effect on the ceding insurer's 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

[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.

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 3.2.3.

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 equalled to approximately 87% of ceded premium (Appendix B).

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 require immediately funding 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. It traps 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.

[4]P&C insurers hold approximately two-thirds 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.

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., 2007), 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 three 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, and (3) quarterly data on Florida homeowners insurance from the Florida Office of Insurance Regulation (QUASR).

2.2.1 Reinsurance Data

Our reinsurance data come from NAIC statutory annual statements and cover the period 2004 through 2024. The income statements report reinsurance ceded and losses reimbursed at the insurer-line-of-business level. In Table 1, we report summary statistics for these reinsurance flows for homeowners insurance. We do this for two groups: all insurers, and for insurers who are buying tail-risk reinsurance.[5]

Table 1: Summary Statistics for Reinsurance Data

		Pre-treatment				Post-treatment			
Outcome	Unit	Mean	SD	P25	P75	Mean	SD	P25	P75
All HO cedents									
Direct premiums written (DPW)	$M	3,347.20	4,314.87	549.75	4,563.46	3,444.57	5,020.37	676.13	4,439.79
HO premium ceded to reinsurance	$M	268.74	583.88	10.39	262.55	588.45	1,324.17	17.61	504.26
HO losses paid by reinsurers	$M	141.46	351.46	1.95	100.54	326.76	756.16	5.73	278.35
Ceded loss ratio (price)		0.53		0.20	0.65	0.56		0.34	0.65
Number of reinsurers		75.17	106.50	5.00	82.00	79.40	119.68	3.00	91.00
Tail-Risk Exposed									
Direct premiums written (DPW)	$M	2,394.34	2,167.62	613.00	4,501.42	2,840.00	3,068.49	512.35	3,913.67
HO premium ceded to reinsurance	$M	84.75	136.73	8.72	97.36	167.59	214.83	19.69	283.60
HO losses paid by reinsurers	$M	30.37	94.83	0.49	25.51	69.23	140.28	0.64	88.21
Ceded loss ratio (price)		0.36		0.04	0.58	0.41		0.07	0.52
Number of reinsurers		58.27	36.35	38.00	81.00	123.45	191.03	46.00	96.00

Note: Sample restricted to firm-years with positive HO ceded premium. All reinsurance quantities are annual FLOWS (no year-end stocks); dollar amounts in $M. The ceded loss ratio (price) is the DPW-weighted pooled ratio (total HO losses paid / total HO premium ceded); its P25/P75 are the firm-year distribution. Tail-Risk Exposed = cedents with $\geq$ 25% of classified ceded premium in catastrophe / right-tail reinsurance pairs (AEA right-tail classifier). DPW-weighted within period. Pre = event time < 0; Post = event time ≥ 0.

Homeowners premium ceded to reinsurance more than doubles after treatment (from $269M to $588M on average), and the number of reinsurers rises from about 75 to 79. As our measure of reinsurance price, tail-risk exposure insurers receive roughly 36 cents of losses back per dollar of homeowners premium ceded before treatment and 41 cents after. These changes in the summary statistics – a higher quantity of reinsurance, more reinsurance competition, and a lower reinsurance price – 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 that 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.

[5] We define this by categorizing each reinsurance contract as tail-risk exposed if it pays out no losses in more than half of the sample. We define an insurer as tail-risk exposed if more than a quarter of its reinsurance premium is tail-risk exposed.

2.2.2 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 2.

We correct a small set of insurer-quarters where per-policy premiums appear mis-scaled by a factor of 100, drop observations with non-finite log premium or log rate per \$1,000 of exposure, and winsorize the top and bottom 1% of rate per \$1,000 (cutoff ≈ 9.83). 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 likely the most reliant on reinsurance, our results are likely conservative. Summary statistics for the Florida QUASR data are in Table 2 below.

As their states of domicile are treated, more insurers begin writing business in Florida. Those that do 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, cedents' exposure to counterparty risk, and a proxy for reinsurance prices.

Table 2: 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	2,471.55	2,226.76	797.00	3,357.00	11,460.52	20,195.28	1,518.00	11,931.00
Total direct premium written (DPW)	$M	9.30	11.51	2.04	11.65	28.06	54.48	3.11	26.78
Avg. premium per policy	$	4,053.87	3,920.08	1,700.74	4,122.17	2,557.00	2,071.85	1,548.93	2,811.13
Total insured exposure	$M	2,503.69	3,541.40	530.47	2,666.84	4,773.87	7,076.41	758.92	5,301.12
Rate per $1,000 of exposure	$	4.13	1.41	3.15	4.93	4.89	1.92	3.41	6.11
Wind share of DPW		0.92	0.16	0.90	1.00	0.97	0.10	0.98	1.00
New policies written	policies	35.46	133.77	2.00	31.00	795.24	2,171.92	31.00	496.00
Net new policies written	policies	-41.20	172.04	-68.00	-4.00	164.95	1,906.59	-109.00	126.00
Insurers per county-quarter	insurers	24.94	10.71	17.00	32.00	59.89	9.56	52.00	68.00

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:

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

where Y_{it} is the outcome of interest for insurer i in year t, α_i is an insurer fixed effect, λ_t is a calendar-year fixed effect, and $\text{RelTime}_{it} = t - t^{\text{adopt}}_{s(i)}$ 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, γ_τ identifies the average treatment effect on the treated (ATT) at event time τ.

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 heterogeneous 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 τ, 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., DPW), D_{it} is the treatment indicator, and β_1 captures the post-treatment change in the slope. The event-study analog replaces D_{it} with $\sum_\tau \mathbf{1}\{\text{RelTime}_{it} = \tau\}$. 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:

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

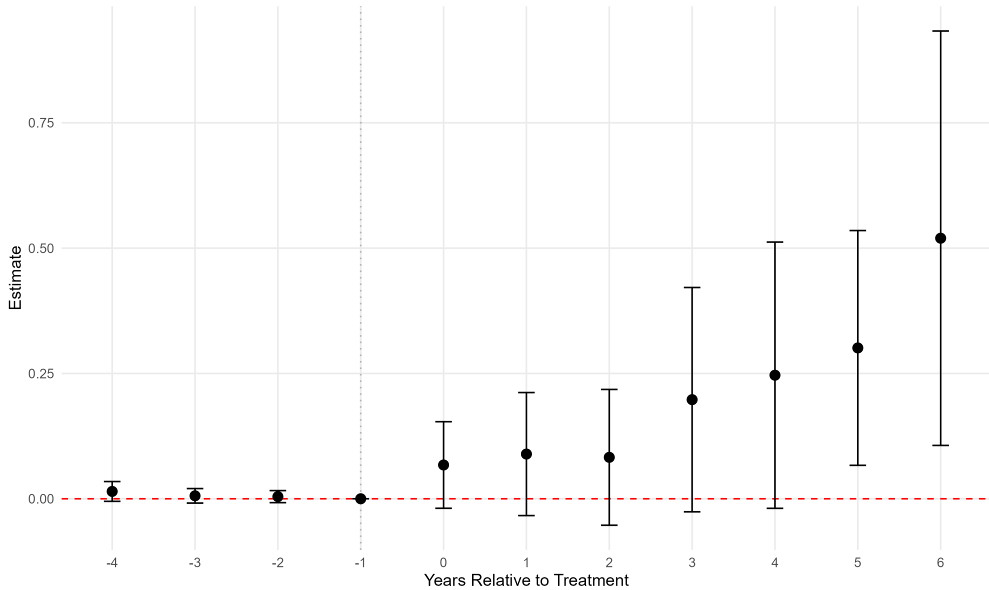
where $\text{RelTime}_{it} = t - t^{\text{adopt}}_{s(i)}$ and $\tau = -1$ is the reference period. The slope coefficients $\{\gamma_\tau\}$ capture how the marginal relationship between Y and X shifts in event time. In both stages we weight by insurer-year direct premiums written. Standard errors are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



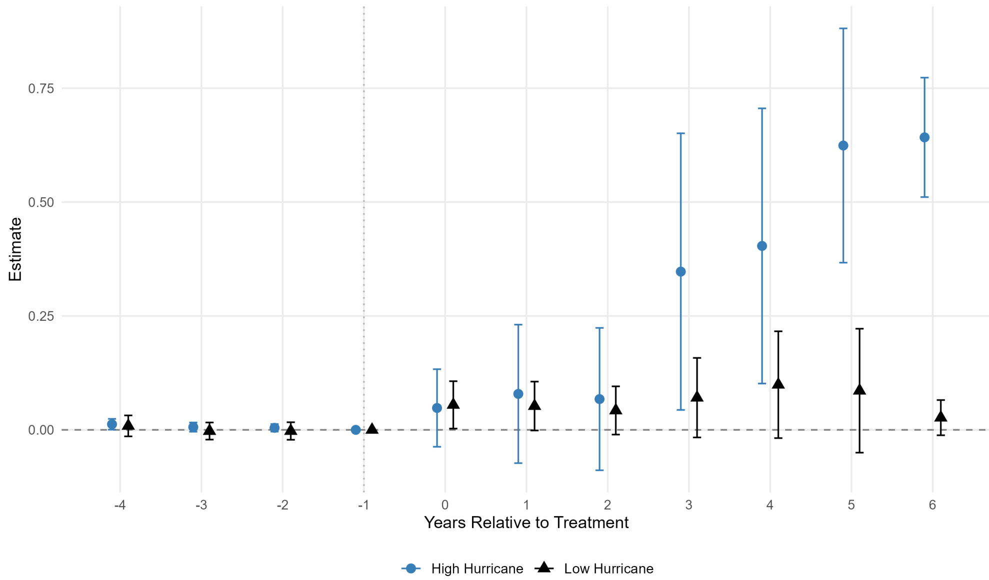
Notes: The figure reports an event study where the outcome is dollars of reinsurance ceded to foreign reinsurers; the exposure variable is direct premiums written (DPW). Event time τ is relative to the insurer's domicile state adoption. Coefficients $\{\gamma_\tau\}$ are from the two-stage specification 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[6]. The results are in Figure 3.

[6]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.

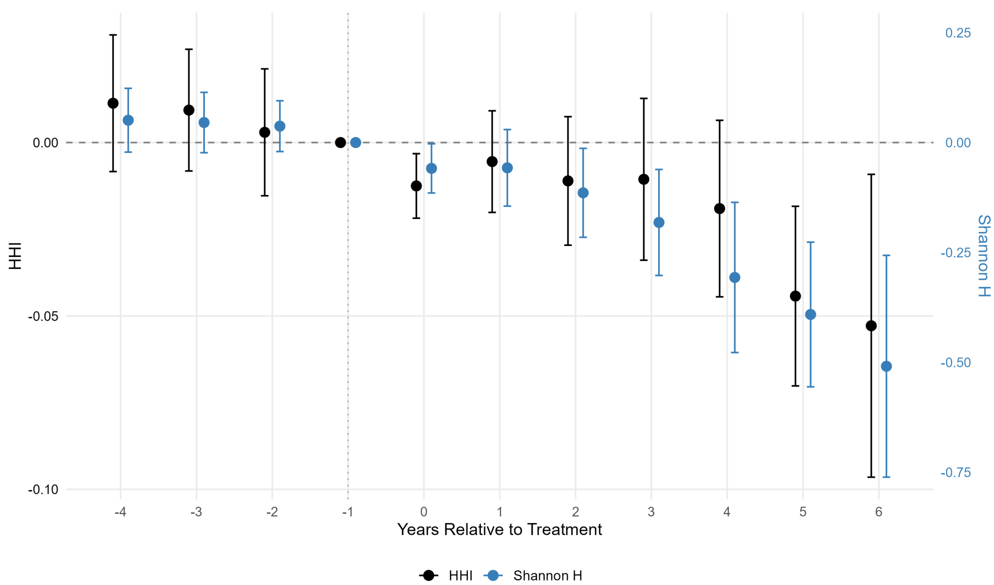
3.2.2 Market Concentration

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 3.2.4, 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 are outcome specifications following 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.

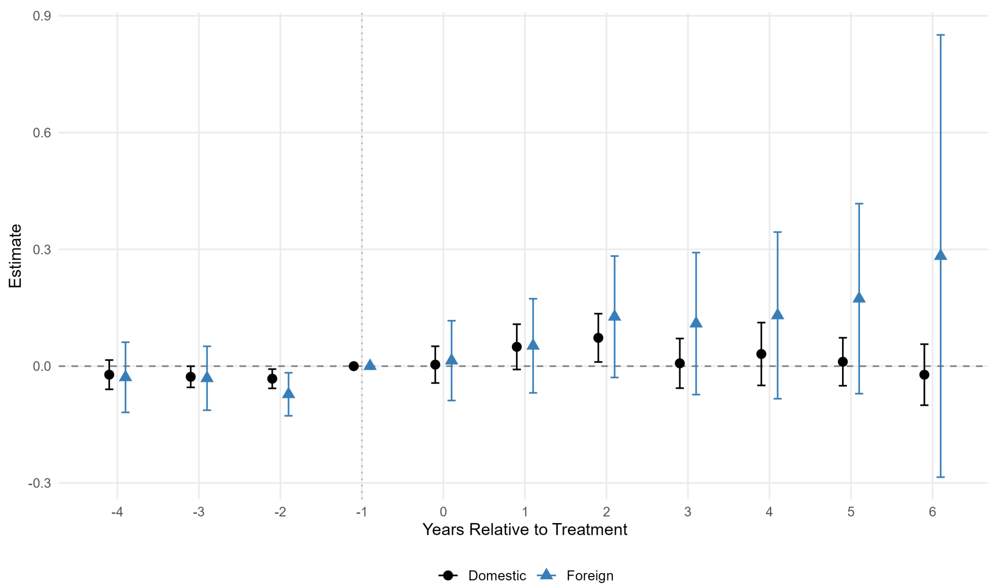
Figure 4: Reinsurance Market Concentration: HHI and Shannon Entropy



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 τ 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. Together with the average-cession-size results, this count decomposes the ceding response into an extensive margin (number of relationships) and an intensive margin (size per relationship): $\log(\text{ceded}) = \log(\#\,\text{reinsurers}) + \log(\text{avg. ceded per reinsurer})$. 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 τ 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 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

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

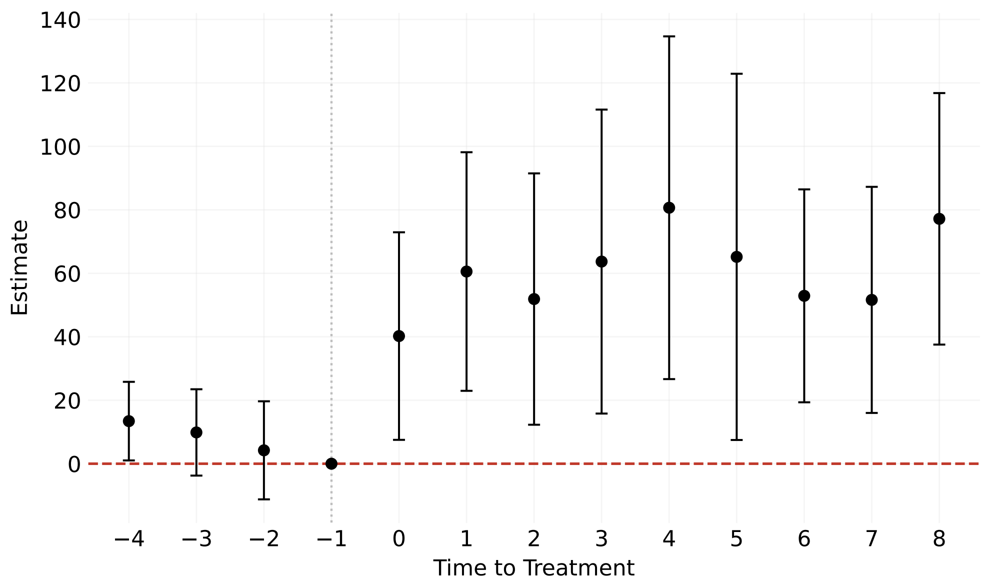
where R_{ijt} is the recoverable owed by reinsurer j, $p(r_j^*)$ is the one-year impairment probability associated with the reinsurer's financial-strength rating, and λ_{ijt} is the uncollateralized share of the recoverable under the rules applicable to the cedent. The baseline assumes a 50% recovery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

Figure 6 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 6: 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 τ 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 bands are 90% confidence intervals. The sample is unaffiliated third-party property-casualty reinsurance, a national book broader than the homeowners cessions in Figures 2–**??**. 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 sensitivity analyses.

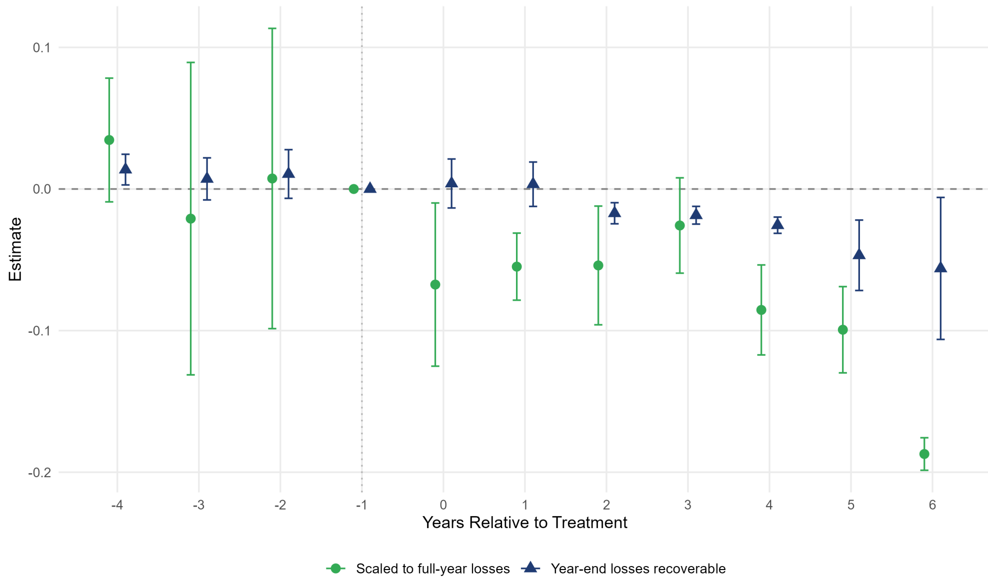
3.2.4 A Proxy for Reinsurance Price Declines

In this section, we study the effect of the collateral reform on a proxy for reinsurance prices. Ideally, we would measure the price of reinsurance as the premium paid per dollar of expected loss transferred to the reinsurer. However, we do not observe expected losses, so we proxy for them with realized losses: a contract that pays back more in losses per dollar of premium ceded is cheaper. Moreover, we face an additional data limitation: our contract-level (insurer-by-reinsurer) data record premiums ceded for the entire calendar year, but not the full-year losses ceded under each contract. They record only the losses recoverable from the reinsurer at year-end. The former is an annual flow as would be found on a cash-flow statement; the latter is an end-of-year stock item from the balance sheet.

Using year-end recoverables in place of full-year ceded losses is innocuous as long as the share of a year's ceded losses still outstanding at year-end is stable across reinsurers and over time. Because our event study compares each insurer to itself relative to the year before treatment, it differences out any reinsurer that is persistently slow or fast to settle, along with any common time trend. It would not handle a change in settlement speed that coincides with treatment, for example if foreign reinsurers paid faster or slower after the reform. We return to this concern below.

For each foreign reinsurer we run the same two-stage specification (equations 5–6), with losses recoverable as the outcome and premiums ceded as the exposure, so the coefficient measures dollars recoverable per dollar ceded. As before, we flip the sign of these coefficients, so that a negative value means cheaper reinsurance (more losses recovered per dollar of premium ceded). The baseline series (in blue) in Figure 7 shows no pre-trend and a steady decline in price after a state adopts the reform. By the fifth year after adoption, the losses insurers recover from foreign reinsurers rise by about \$0.05 per dollar of premium ceded. At baseline they recover about \$0.09 per dollar ceded, so a dollar of recovered loss cost about \$11 in premium before the reform and about \$7 after. Hence the price of foreign reinsurance fell by roughly a third. This is consistent with the model-implied price decrease we observe in Section 5 below.

Figure 7: Reinsurance Prices, Foreign Cessions



Notes: Slope event study for non-NAIC (foreign) homeowners multiple-peril cessions, estimated from equations (5) and (6) with state-of-domicile and year fixed effects, insurer DPW weights, and standard errors clustered by insurer. The exposure variable is premium ceded. In the baseline series (blue triangles) the outcome is year-end losses recoverable; in the second series (green circles) the outcome rescales year-end recoverables to full-year ceded losses using the insurer-level ratio of full-year ceded incurred losses to year-end recoverables (see text). Coefficients are sign-flipped so that negative values denote a cheaper price (more loss returned per dollar of premium ceded); the reference period is $\tau = -1$.

To check whether year-end recoverables are a good proxy for full-year losses, we use the correct, annual flow losses paid on ceded reinsurance, but observed at the insurer level, not at the reinsurer-insurer contract level. We rescale each counterparty's year-end recoverables to a full-year-loss basis using the insurer's ratio of full-year losses to year-end recoverables, and re-estimate. This assumes that the ratio between an insurer's total year-end losses recoverable and its annual losses paid applies equally across all its reinsurance contracts. The rescaled series (in green) in Figure 7 tracks the baseline closely, and the post-treatment decline is if anything a little larger. So the proxy is not driving our result; correcting for it only strengthens it.

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 λ_t with a county-by-quarter fixed effect λ_{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 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^{wind}_{ict} and r^{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-included 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 8.

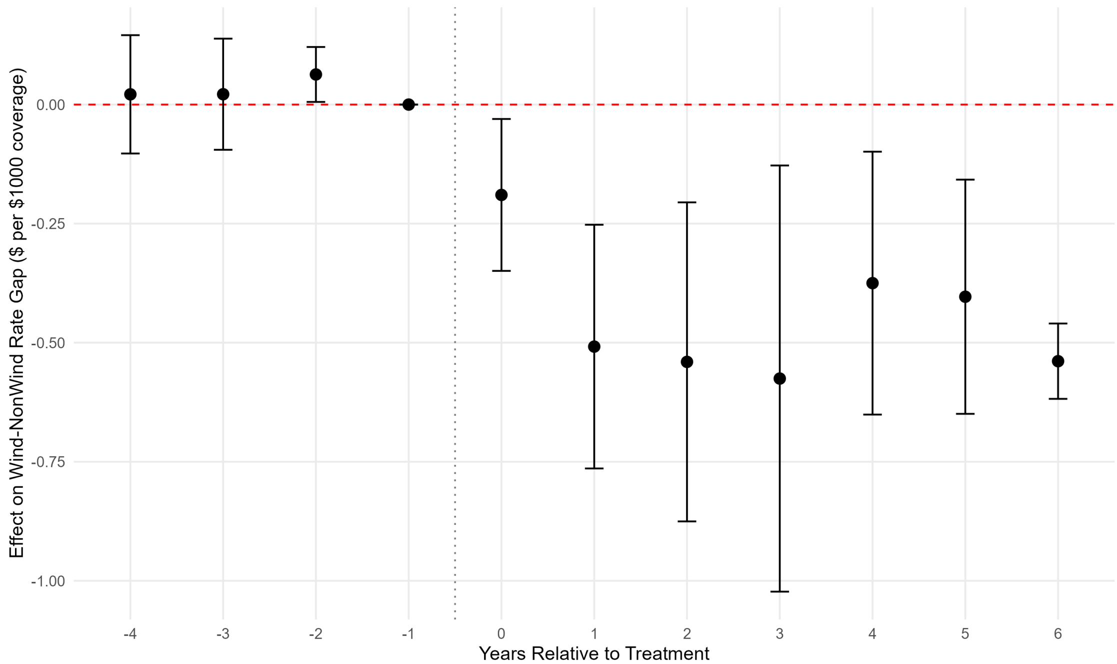
We see a clear drop in the pricing of wind-exposed homeowners insurance policies after treatment, relative to non-wind-exposed policies. The baseline rate per $1,000 of exposure is approximately 4.13 (per Table 2). Hence, a drop of 0.5 equates to a rate decrease of approximately 12%. As a placebo, the non-wind rate level falls by far less, and only at longer horizons (Appendix Figure 20). Moreover, the effect is similar among incumbent firms (Appendix Figure 21), 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 22).

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

Figure 8: Policy Pricing: Wind-Included vs. Wind-Excluded Coverage

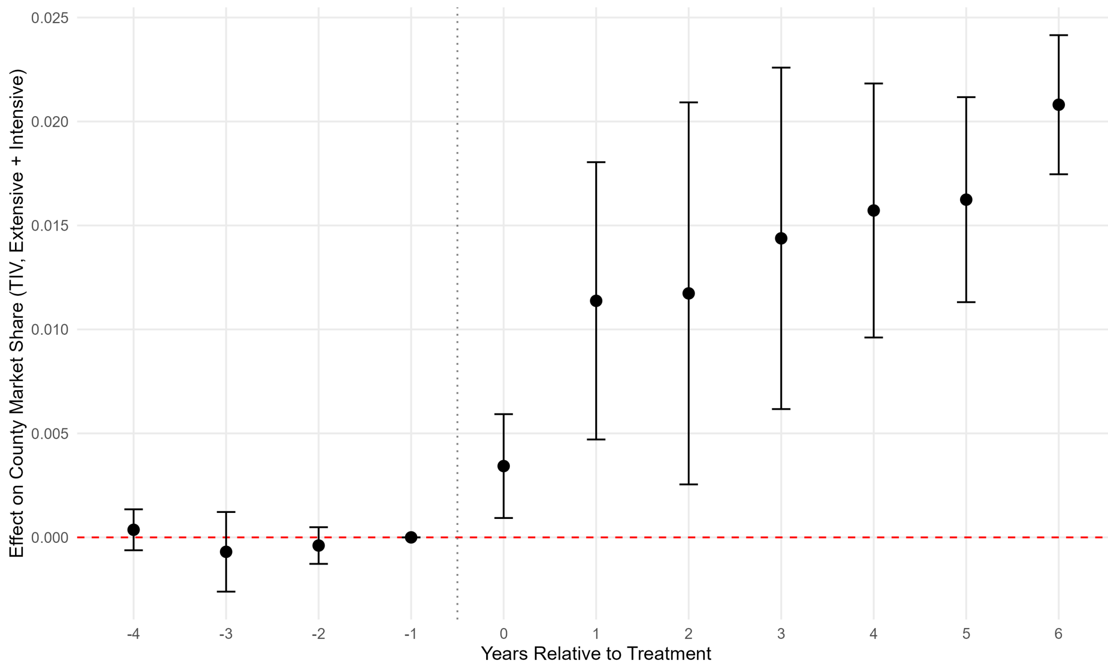


Notes: The figure reports a triple-difference event study estimated on insurer-county-quarter observations from Florida QUASR data. The outcome is the *wind premium gap*-the within-insurer-county-quarter difference in the rate per $1,000 of coverage between wind-inclusive and wind-excluded policies, 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 τ 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.

will rise. The event study in Figure 9 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 9: Insurance Market Share: Total Insured Value



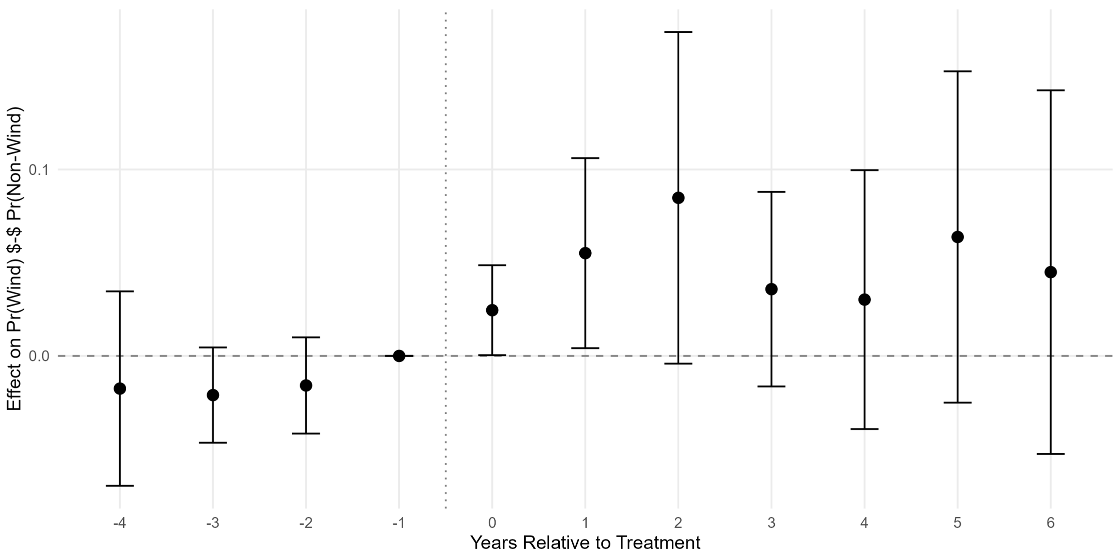
Notes: The figure reports an event study estimated on an insurer-county-quarter panel constructed from Florida QUASR data. For each cell, the outcome is the insurer's share of total insured value (TIV) in the county-quarter, which co-moves with both the extensive margin (number of policies written) and the intensive margin (coverage depth per policy). Event time τ 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 α_i and county-by-quarter fixed effects λ_{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 10 reports the event-study estimates.

Figure 10: 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 τ 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.[7] 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.[8]

[7] 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.

[8] 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 sample, homeowners' insurance policy counts exceed mortgaged owner-occupied units in $1{,}599$ of $2{,}934$ county-quarters (54.5%), 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), ξ_{jmt} is unobserved product quality, and ε_{ijmt} is an i.i.d. type-I extreme value error. The price coefficient α_i is lognormally distributed, with $\bar{\alpha}$ governing the mean disutility of premiums and σ_α governing the dispersion of price sensitivity across the population. A separate normally distributed random coefficient β_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.[9] We normalize the outside option of not purchasing a wind-coverage product so that $u_{i0mt} = \varepsilon_{i0mt}$.[10]

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:

$$\ln(mc_{jmt}) = \gamma_0 + \gamma_r \, \text{post785}_{ft} + \gamma_c \, \mathbb{1}[\text{Citizens}]_{jt} + \bar{\mathbf{w}}'_{fmt}\boldsymbol{\gamma}_m + \omega_{jmt}. \tag{15}$$

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 γ_r is the parameter of central interest, summarizing how reinsurance-market reform passes through to insurers' marginal costs. The Citizens indicator absorbs the fact that Florida's government-sponsored residual-market insurer operates under a distinct cost structure. $\mathbf{w}_{fmt}$ includes log exposure per policy, which proxies for the scale

[9] 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).

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

of insured structures, county-by-year fixed effects, which control for the riskiness level of the geographic market, and firm fixed effects. ω_{jmt} is the cost residual.

Explicitly modeling the marginal cost function, rather than backing it out solely from the pricing FOCs, is essential for our identification strategy: the reduced-form cost shifter post785_{ft} delivers exogenous variation in marginal costs that, combined with demand-side instruments, allows us to separately identify movements in marginal costs from movements in markups.

5.2 Estimation and Identification

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:

$$\mathbb{E}\Big[\xi_{jmt}(\boldsymbol{\theta}) \cdot \mathbf{z}^d_{jmt}\Big] = \mathbf{0}, \qquad \mathbb{E}\big[\omega_{jmt}(\boldsymbol{\theta}) \cdot \mathbf{z}^s_{jmt}\big] = \mathbf{0}. \tag{16}$$

We estimate the model on 192,018 product-market observations spanning 2,931 county-quarter markets and 137 insurers active in the Florida wind-coverage segment between 2010 and 2020.

Demand Instruments. Premiums are endogenous because insurers observe preference shocks ξ_{jmt} when setting prices, inducing correlation between p_{jmt} and ξ_{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 contemporaneous 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.

Supply Instruments. The challenge in identifying the cost function is to find an excluded instrument that shifts reinsurance exposure without directly shifting homeowners' claims costs. We use a granular Bartik shift-share constructed from the pre-reform composition of each cedent's reinsurance portfolio interacted with the staggered state-level adoption of the reinsurance collateral policy change,

$$Z_{ft}^{\text{Bartik}} = \sum_{s} \underbrace{\text{share}_{fs}^{\text{pre}}}_{\substack{\text{firm } f\text{'s pre-period share} \\ \text{of ceded premiums to reinsurers in state } s}} \cdot \underbrace{\text{post785}_{st}}_{\substack{\text{state } s \text{ adopted} \\ \text{the policy change by time } t}} . \tag{17}$$

Pre-period shares are computed over 2009–2012 using NAIC Schedule F Part 3, which describes an insurer's reinsurance agreements. The exogeneity of this Bartik instrument comes from the granularity of firm-level reinsurer portfolios (Borusyak et al., 2022): different primary insurers are exposed to different combinations of reinsurer-domicile states for reasons unrelated to Florida-specific cost shocks. We include rival sums of the Bartik instrument as additional supply moments, as well as BLP and differentiation instruments built from log exposure per policy, to discipline the remaining cost parameters.

Identification. 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, σ_α, is identified by second-moment variation: substitution 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, σ_0, is identified by variation in the total inside share relative to what the mean utility alone would predict.

The effect of the reinsurance collateral policy change on marginal costs, γ_r, is identified by the correlation between the Bartik shift-share and insurers' recovered marginal costs, net of the demand-side markups implied by equation (14). Intuitively, conditional on the implied markup, the Bartik-induced variation in reinsurance exposure traces out movements in mc_{jmt}, and the average slope pins down γ_r. The remaining cost parameters $(\gamma_0, \gamma_e, \gamma_c, \boldsymbol{\gamma}_\tau, \boldsymbol{\gamma}_m)$ are identified by standard variation in log exposure per policy, the Citizens indicator, year fixed effects, and Mundlak group means, all of which are treated as predetermined.

5.3 Estimation Results

Table 3 reports four specifications: a demand-only benchmark (column 1), a joint specification with a price random coefficient only (column 2), our preferred joint specification (column 3), and an extension that adds insurer-type random coefficients (column 4). The four columns yield economically similar magnitudes, both in estimates and derived statistics. Therefore, we will focus our discussion on column (3).

Consumer Preferences and Elasticities. Panel (a) of Table 3 reports consumer preference estimates. The mean price coefficient $\bar{\alpha}$ is -0.45 (in \$1,000s), confirming that consumers dislike increases in the annual premiums. Price sensitivity is highly dispersed — the lognormal scale parameter $\sigma_\alpha = 1.25$ generates a long right tail of price-sensitive households who substitute toward the uninsured outside option. The dispersion of the inside-good random coefficient is small and statistically indistinguishable from zero ($\sigma_0 = 0.02$, s.e. 1.09), indicating that the firm, county, and year-quarter Mundlak controls absorb essentially all of the heterogeneity in inside-share valuation that would otherwise load onto σ_0.

We report two additional statistics to characterize the substitution patterns implied by these estimates. First, at the intensive margin, we calculate own-price elasticities at the status quo. The intensive margin adjustments that we observe are changes in the insurance's coverage limit. The mean and median are -2.32 and -2.02, respectively, with an interquartile range of -2.53 to -1.64. These magnitudes lie within the range of recent estimates for Florida homeowners (Oh et al. 2026 report -1 to -3; Ge 2022 report similar magnitudes).

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 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 64.0%, implying an arc elasticity of take-up with respect to price of -0.53; a symmetric 10% price increase reduces take-up to 57.2%, implying an elasticity of -0.58. In additive terms, a \$100 reduction in annual premiums raises take-up by roughly 1.9 percentage points, while a \$500 reduction raises it by 10.0 points. These magni-

Table 3: Joint BLP Estimates: Demand and Cost Function

	(1)	(2)	(3)	(4)
(a). Demand: Linear Parameters and Random Coefficients				
Mean price coefficient ($-\bar{\alpha}$)	0.468^{***}	0.576^{***}	0.447^{***}	0.543^{***}
	(0.006)	(0.005)	(0.005)	(0.005)
Dispersion of price RC (σ_{price})	1.338^{***}	1.201^{***}	1.246^{***}	1.276^{***}
	(0.027)	(0.004)	(0.004)	(0.005)
Dispersion of constant RC (σ_{const})	0.008	-	0.019	0.031
	(29.834)		(1.094)	(1.544)
Dispersion of Citizens RC (σ_{citz})	0.000	-	-	0.091
	(121.859)			(0.603)
Dispersion of FL-specialist RC (σ_{fl})	0.000	-	-	0.000
	(13.117)			(3.493)
(b). Cost Function ($\log mc_{jmt} = \mathbf{w}'_{jmt}\boldsymbol{\gamma} + \omega_{jmt}$)				
Constant (γ_0)	-	-6.178^{***}	-9.706^{***}	-7.014^{***}
		(0.252)	(0.434)	(0.326)
log(exposure/policy) (γ_e)	-	0.758^{***}	0.747^{***}	0.760^{***}
		(0.025)	(0.034)	(0.027)
Citizens indicator (γ_c)	-	0.326^{***}	0.393^{***}	0.333^{***}
		(0.010)	(0.012)	(0.010)
Post Model Law 785 (γ_r)	-	-0.088^{***}	-0.104^{***}	-0.080^{***}
		(0.003)	(0.004)	(0.003)
(c). Derived statistics (means across products × markets)				
Mean own-price elasticity	−2.344	−2.627	−2.321	−2.534
Mean marginal cost ($1,000s)	1.802	1.501	1.399	1.473
Mean observed price ($1,000s)	2.296	2.296	2.296	2.296
Mean Lerner index $(p - mc)/p$	0.332	0.304	0.337	0.312
Reinsurance / marginal cost (aggregate)	0.610	0.735	0.788	0.748
Mean own-cost pass-through ($\partial p/\partial mc$)	-	1.261	1.387	1.280
Mundlak controls (firm, county, county×year)	-	Yes	Yes	Yes
Year fixed effects	-	Yes	Yes	Yes

Notes: Standard errors stacked below estimates in parentheses; significance: $^{***}p < 0.01$, $^{**}p < 0.05$, $^{*}p < 0.10$. Prices are in $1,000s of annual premium. The price random coefficient is log-normal: $\alpha_i = -(\bar{\alpha} + \exp(\sigma_{\text{price}}\nu_i^{\alpha}))$ with $\nu_i^{\alpha} \sim N(0,1)$. All other random coefficients are normally distributed: $\beta_{ki} = \beta_k + \sigma_k \nu_{ki}$ with $\nu_{ki} \sim N(0,1)$. Demand moments use Berry et al. (1995), differentiation (Gandhi and Houde, 2019), and recentered other-state loss instruments; supply moments additionally include a Bartik shift-share instrument built from pre-period reinsurer-domicile portfolio shares interacted with staggered state-level Model Law 785 adoption. Column (1) is demand-only: marginal cost is the FOC residual $p - \mu$ (no cost block estimated), so panel (b) is omitted. In panel (c), marginal cost is the structural MC $\exp(\mathbf{w}'\gamma + \omega)$ in columns (2)–(4) and the FOC residual $p - \mu$ in column (1). The Lerner index is computed at observed prices as $L = (p - mc)/p = \mu/p$, averaged across products. The reinsurance / marginal cost row reports the aggregate ratio where the numerator is non-affiliated reinsurance premium ceded per policy, computed over observations with $mc > 0$ and non-missing reinsurance data. When we calculate pass-through $\partial p_j/\partial mc_j$, the inverse Bertrand-FOC Jacobian $\Upsilon = (I - dM/dp)^{-1}$ is restricted to the private-product block and re-inverted with Citizens prices held fixed; the diagonal of the resulting matrix is the own-cost rate. Sample: wind-only Florida homeowners, 2010Q1–2020Q4; $N = 152,507$ product-market observations across 2,935 county-quarter markets and 131 firms.

tudes imply a sizable extensive margin: combining the 9.9% marginal-cost reduction with the 1.39 own-cost pass-through rate (discussed below) gives an implied 13.7% premium reduction, which corresponds to roughly a 5 percentage-point increase in take-up in partial equilibrium, before any Bertrand repricing or entry response.

Cost Functions and Markup. Panel (b) of Table 3 reports the cost function estimates. The log exposure elasticity $\hat{\gamma}_e = 0.75$, indicating that a 10% increase in the insured value per policy raises marginal cost by roughly 7.5%, close to but below the unitary elasticity that would arise if claims scaled linearly with insured values. The Citizens indicator $\hat{\gamma}_c = 0.39$ confirms that Florida's residual-market insurer operates with a higher cost stack than private carriers, consistent with its statutory mandate to cover risks that the private market declines. 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), which averages $1,399.

The reinsurance-reform coefficient is the parameter of central interest: $\hat{\gamma}_r = -0.104$, 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 $\exp(-0.104) - 1 \approx 9.9\%$. The estimate is robust across various model specifications, which yield a range of -7.7% to -9.9%.

Comparing the mean estimated marginal cost of $1,399 against a mean observed premium of $2,296, we derive an average Lerner index of 0.34.[11] This is consistent with a highly concentrated market with meaningful product differentiation across wind carriers. This implied markup closely aligns with the direct business underwriting margin for Florida homeowners reported in NAIC Schedule data: Pooling all FL homeowners' insurers from 2010–2020, the average direct-basis combined ratio is 0.73, implying an underwriting margin of 0.27; the same series outside the 2017–2018 Hurricane Irma window averages 0.41.

The non-affiliated reinsurance premium ceded per policy averages 0.79 of marginal cost in our preferred specification, with values across specifications spanning 0.61 to 0.79. Multiplying by the estimated cost-to-price ratio $(1 - L) = 0.66$ rescales this to the more familiar reinsurance-share-of-direct-premium basis, giving $0.79 \times 0.66 = 0.52$ for column 3 and a cross-spec range of roughly $(0.42, 0.52)$. 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 lie at

[11] 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.

the upper end of the $[0.30, 0.50]$ band typically cited for Florida homeowners in the catastrophe-reinsurance literature (**?**Born and Klimaszewski-Blettner, 2013a). 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 ($\partial p/\partial mc$) is 1.39 (median 1.22, inter-quartile range $[1.18, 1.54]$). Pass-through above one is a known feature of differentiated products. In Bertrand competition with random-coefficient logit demand, when the curvature of demand makes inverse-elasticity markups rise with cost, equilibrium prices overshoot uniform cost shocks (Weyl and Fabinger, 2013). This implies that the -9.9% marginal-cost reduction from the reform translates into approximately a -13.7% reduction in equilibrium premiums on treated products under a fixed market structure, closely matching the 12% intensive-margin estimate from the reduced-form analysis in Section 4.

To position our pass-through estimates within the recent homeowners-insurance literature, we benchmark it against three external estimates. The literature reports several distinct objects — a disaster-risk gradient $\partial p_{HO}/\partial L^e$, a shock-DiD treatment effect on annual premiums, and a reinsurance-cost elasticity η_{reins}. We convert our estimated pass-through ρ into each paper's units using simple chain-rule and accounting identities, and find that our estimate reproduces all three benchmarks within tight tolerances.

First, we compare with the disaster-risk gradient in Keys and Mulder (2024). They report that, in the cross-section of mortgage escrow data, a one-standard-deviation increase in expected catastrophe loss L^e raises annual homeowners premiums by approximately \$450. We map ρ to this object by the chain rule, $\partial p_{HO}/\partial L^e = \rho \times \partial mc/\partial L^e$. The cost-side scaling factor $\hat{\beta} \equiv \partial mc/\partial L^e$ is estimated on a separate NAIC insurer-state-year panel, regressing marginal cost on a long-run loss-ratio proxy with insurer and year fixed effects; we obtain $\hat{\beta} \approx 0.94$. Our state-level L^e has a cross-sectional standard deviation of 0.105, but Keys and Mulder identify their gradient off zip-level dispersion, which is substantially larger because most variation in catastrophe exposure is within states (e.g., coastal vs. inland Florida) rather than across them. Following standard rule-of-thumb adjustments for between-zip vs. between-state variance, we inflate the state-level standard deviation by a factor of three, yielding an implied gradient $\rho \times \hat{\beta} \times 3 \cdot \mathrm{SD}(L^e) \approx \$550$ — within 20% of the published \$450.[12]

Second, we compare with the reinsurance shock-DiD in Oh et al. (2024). They estimate that the discrete post-2022 reinsurance-pricing shock raised annual premiums by approximately \$425 in top-decile catastrophe-exposed zip codes. To map ρ to this object, we treat their event as a proportional shock to the rate-on-line p_R (the per-unit reinsurance price) of the consensus industry magnitude, $+30\%$, applied to a baseline reinsurance cost per policy of $C_R \approx \$1{,}275$ in a top-decile zip. Under our cost decomposition, this maps one-for-one into a marginal-cost shock of $\Delta mc \approx \$383$, which $\rho = 1.39$ translates into a predicted premium change of approximately \$531 — bracketing the \$425 estimate from above. Lower shock assumptions ($+20\%$ rate-on-line) deliver \$354, so the treatment effect lies inside the range implied by ρ at plausible shock magnitudes.

Third, we compare with the reinsurance-cost elasticity in Born and Klimaszewski-Blettner (2013b); Niehaus (2002). They report $\eta_{\text{reins}} \equiv (\partial p_{HO}/p_{HO})/(\partial C_R/C_R)$, the percent response of homeowners premiums to a percent change in the per-policy reinsurance cost C_R, with values in the range 0.2–0.5. By construction $\eta_{\text{reins}} \approx \rho \cdot s$, where $s \equiv C_R/p_{HO}$ is the reinsurance share of the premium dollar. Our sample-wide s is 0.21, giving $\eta_{\text{reins}} \approx 0.15$; at the FL-coastal share $s \approx 0.30$ this rises to ≈ 0.20, at the lower edge of the literature band. The mild gap is consistent with Florida's binding rate-regulation friction relative to the cross-state samples on which those papers were estimated.

6 Welfare Effect of the Collateral Reform

We now use the estimated model to quantify the welfare effects of the credit-for-reinsurance reform. Holding fixed the set of operating insurers and the residual-market regulator, we re-simulate Bertrand pricing equilibria, sequentially turning on each mechanism through which the reform affects consumer prices: lower marginal costs, strategic markup adjustment, and entry of new product offerings. The decomposition isolates the extent to which the consumer welfare gain from the reform reflects mechanical cost pass-through, oligopolistic best-response in markups, and expanded consumer choice. We measure consumer welfare as the Marshallian gain from price reductions, weighted by post-reform purchase shares, summed across products and households.[13]

[12] Replacing the $3\times$ multiplier with FEMA NRI zip-level AAL data would tighten this comparison; we leave this for ongoing work.

[13] We adopt the Marshallian welfare measure as the headline because it makes only assumptions disciplined by the data we observe: equilibrium shares at observed prices and price changes implied by the structural pricing first-order condition. An alternative measure that aggregates the (income-deflated) inclusive value of the consumer choice set

Throughout, dollar figures are reported per household per year, since premiums in our estimation sample are measured as average annual premium per policy in force; aggregates use Florida's roughly 4.9 million owner-occupied households.

6.1 Consumer Gains from the Collateral Reform—Three Mechanisms

We construct four counterfactual equilibria, indexed (0)–(3). *Counterfactual (0)* represents the world without the reform: insurers face their pre-reform marginal cost—reflecting the higher reinsurance loading that prevailed when foreign reinsurers were required to post full collateral—reform-induced entrant products are absent from the market, and incumbent insurers set markups in Bertrand equilibrium against this restricted product space. *Counterfactual (1)* introduces only the mechanical cost pass-through of the reform: marginal cost falls to its post-reform level, but markups remain fixed at their no-reform values and the choice set continues to exclude entrants. This counterfactual measures how much of the reform's price reduction would survive in a partial-equilibrium world in which firms did not respond strategically. *Counterfactual (2)* additionally allows incumbents to re-optimize markups against the lower marginal cost, still without entry. The difference between (1) and (2) isolates the extent to which industry-wide cost reductions induce competitive markup compression beyond mechanical pass-through. *Counterfactual (3)* is the observed equilibrium: post-reform marginal cost, full Bertrand pricing, and the realized post-reform product space including 86 entrant offerings. The difference between (2) and (3) isolates the reform's variety channel. Table 4 reports the per-channel welfare gain across the four counterfactuals.

Cost pass-through (0)→(1). The reform's first effect is to lower insurers' marginal costs by reducing the collateral that foreign reinsurers must post against ceded business. Among insurers that respond to the reform, our estimates imply that marginal cost falls by approximately \$132 per policy per year following adoption (Table 5). At fixed markups, this cost reduction passes through one-for-one into consumer prices, generating an average welfare gain of \$63 per household per year, or approximately \$308 million per year across Florida homeowners. This channel

books additional surplus from the existence of products consumers do not actually purchase, even when those products have negligible market share, because every consumer is assumed to draw an idiosyncratic taste shock over every available product. This *love-of-variety* surplus is not pinned down by observed shares and prices; it depends on parametric assumptions about consumer tastes for rarely-chosen products and scales with the inverse of the demand-side price coefficient. The two measures coincide closely for any counterfactual that holds the choice set fixed, so the choice between them matters only for the variety channel discussed below. We report Marshallian estimates throughout the main text and provide the inclusive-value upper bound in Appendix **??** as a robustness comparison.

Table 4: Welfare decomposition of the reinsurance collateral reform: three channels, Florida homeowners insurance market.

Channel	\$/HH/yr	FL annual	Share of total	Mechanism
(0)→(1): Cost pass-through	\$63	\$308 M	34%	Lower marginal cost flows directly into prices
(1)→(2): Markup adjustment	\$58	\$285 M	31%	Industry-wide cost shock compresses markups in oligopoly
(2)→(3): Variety (entry)	\$65	\$319 M	35%	Reform-induced products expand the consumer choice set
Total (0)→(3)	\$186	\$912 M	100%	

Notes: Per-household figures are weighted means across the 67 Florida counties; FL annual aggregates use the panel-average count of 4.9 million owner-occupied households. The cost-pass-through and markup-adjustment shares of the cost+markup subtotal follow Table 5: 51.9% mechanical, 48.1% markup. Welfare is measured as Marshallian consumer surplus.

maps most directly to the reform's regulatory intent: the policy was designed to reduce the cost of foreign reinsurance, and any positive cost pass-through flows to consumers.

Markup adjustment (1)→(2). The second effect is the additional price reduction generated by oligopolistic best-response. When marginal cost falls industry-wide rather than for a single firm, every insurer reprices simultaneously and each best-responds to its rivals' lower prices. In a differentiated-products oligopoly, this generates an equilibrium pass-through above one. We measure this directly: among treated products, the per-policy price effect of removing the reform is \$252 per household per year against a cost effect of \$132, an implied pass-through ratio of $252/132 \approx 1.91$ (Table 5). Aggregated over the share distribution, allowing markups to re-equilibrate contributes an additional \$58 per household per year of welfare, or \$285 million annually—roughly the same magnitude as the mechanical channel. A static cost-plus pricing model that ignored equilibrium repricing would therefore understate the consumer welfare gain by approximately one-half. Lower reinsurance costs reduce marginal costs directly and also discipline markups; the latter channel is roughly as large as the former.

Variety (2)→(3). The third effect captures consumer welfare gains from expanded product offerings. We identify 86 firm-county product offerings as reform-induced entrants—firms whose first appearance in a Florida county post-dates their domicile state's adoption of Model Law 785—and the difference between counterfactual (2) and the observed equilibrium (3) measures the welfare consumers would lose if these products were absent. The variety channel contributes $65 per household per year ($319 million annually). The three channels are therefore roughly equal in magnitude: each contributes between 31 and 35 percent of the total reform welfare gain of $186 per household per year, or $912 million annually for Florida homeowners.

Table 5: Equilibrium prices, marginal costs, and markups under reform vs. counterfactual scenarios ($/HH/yr; S3-converged subsample).

Sample	N	p^{S1}	p^{int}	p^{S3}	mc^{post}	mc^{pre}	μ^{S1}	μ^{S3}	$\Delta\mu$
All products, S3-loaded	152,507	1,721	1,826	1,922	1,501	1,607	220	316	+96
Treated	121,839	1,676	1,808	1,928	1,442	1,574	234	354	+120
Treated, high-risk	68,089	2,036	2,210	2,353	1,902	2,076	134	277	+143
Treated, low-risk	53,750	1,221	1,299	1,390	860	939	360	451	+91
Untreated	30,668	1,899	1,899	1,899	1,734	1,734	165	165	+0

Notes: Levels are mean dollars per household per year, computed on the subsample of products for which the S3 (pre-reform) Bertrand simulation converged. p^{S1} is the post-reform Bertrand equilibrium price; mc^{post} is observed marginal cost; $\mu^{S1} = p^{S1} - mc^{post}$. The interim scenario holds markups fixed at μ^{S1} but resets marginal cost to its pre-reform level, $p^{int} = mc^{pre} + \mu^{S1}$. p^{S3} is the full pre-reform Bertrand equilibrium with mc^{pre} and re-optimized markups; $\mu^{S3} = p^{S3} - mc^{pre}$. "Treated" rows have post_785 $= 1$. "High-risk" ("low-risk") counties are above (below) the median of mean expected annual loss. The mechanical price effect of removing the reform is $p^{int} - p^{S1} = mc^{pre} - mc^{post}$; the strategic effect is $p^{S3} - p^{int} = \Delta\mu$. Citizens is pinned at observed prices in all simulations.

Robustness on the variety channel. Our identification of post-785 entrants is necessarily generous: the firm-county appearance criterion does not distinguish reform-induced entry from portfolio expansion that would have occurred regardless. In particular, 90 percent of identified entrant rows belong to FL-domiciled firms, whose treatment status reflects FL's 2009 adoption (which precedes our 2010 sample window) rather than reform-induced cross-state entry. We therefore re-simulate the pre-reform Bertrand equilibrium under two tighter attributions of the entrant set. The first restricts entrants to non-FL-domiciled firms, retaining 10 percent of entrant rows and 23 percent of entrant firms (20 of 86)—these are firms whose entry timing is plausibly tied to their own home-state's adoption. The second restricts further to staggered firms, those whose home-state adoption falls strictly within the 2010–2020 sample window, retaining only 3.1 percent of entrant rows and 11 firms—these are the firms with the cleanest reform-induced-entry signal. The cost-pass-through and markup-adjustment channels are unaffected by either restriction because they are computed without altering the choice set; the welfare consequence is on the variety channel alone. Under the non-FL restriction, the Marshallian variety channel is \$77 per household per year (96 percent of the headline \$80 estimate); under the staggered-only restriction, it is \$77 per household per year (97 percent of headline).[14] The variety channel is therefore demonstrably robust to entrant attribution: even when we drop 97 percent of identified entrant rows and retain only firms whose home-state adoption is squarely within the sample window, the welfare estimate barely moves. The 11 staggered non-FL firms, whose entry is most plausibly attributable to the reform itself, account for nearly all of the variety welfare.

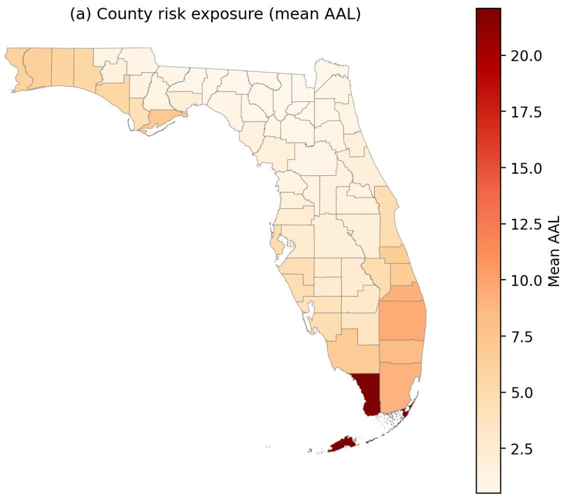
6.2 Spatial Incidence of the Collateral Reform

We next examine how the welfare gain is distributed across Florida's counties. Figure 11 maps four quantities side by side: county risk exposure in panel 11a, the aggregate welfare gain from all three channels combined in panel 11b, the welfare gain from the cost and markup channels alone in panel 11c, and the residual welfare gain from the variety channel in panel 11d (the difference between panels 11b and 11c). The figure makes clear that the spatial incidence of the reform is shaped by two opposing forces.

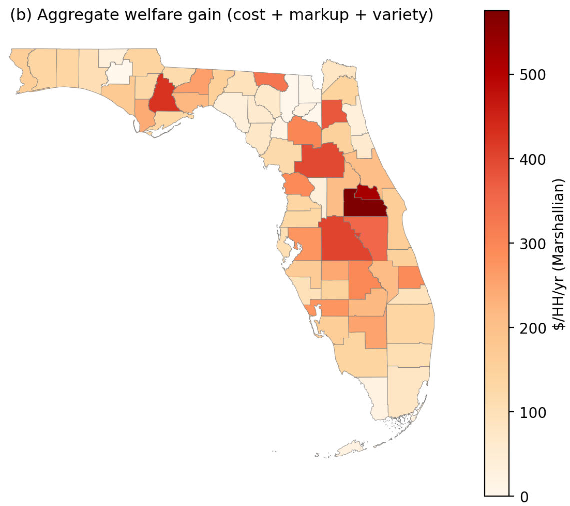
Cost and markup channels favor coastal high-risk counties. Both channels of the cost component scale with reinsurance exposure, though to different degrees. The cost pass-through channel scales sharply with risk: high-risk counties gain \$72 per household per year against \$37 in low-risk counties, a roughly two-to-one ratio that mirrors the cross-sectional variance in ceded reinsurance share. The markup adjustment channel is nearly flat across the risk gradient (\$60 per household per year in high-risk counties against \$53 in low-risk counties)—markup amplification is a generic property of the oligopolistic equilibrium rather than a function of where reinsurance binds. Together, the two cost-side channels deliver \$131 per household per year in high-risk counties and \$90 per household per year in low-risk counties, and the cross-county correlation between welfare gain and mean expected annual loss is positive (Pearson $r = +0.33$, household-quarter weighted). Panel 11c of Figure 11 shows the resulting incidence: the cost+markup gain is concentrated along the coastal panhandle, southeast Florida, and the southwest Gulf coast—precisely the regions of highest hurricane exposure where reinsurance utilization is largest.

[14] All three estimates use the same incumbent-only Marshallian formula, market-exclusion criteria, and household-quarter weights described in the notes to Figure 11. They differ only in the size of the entrant set removed in the pre-reform Bertrand simulation. The headline removes 20,519 entrant rows from 86 firms; the non-FL restriction removes 2,121 rows from 20 firms; the staggered-only restriction removes 629 rows from 11 firms.

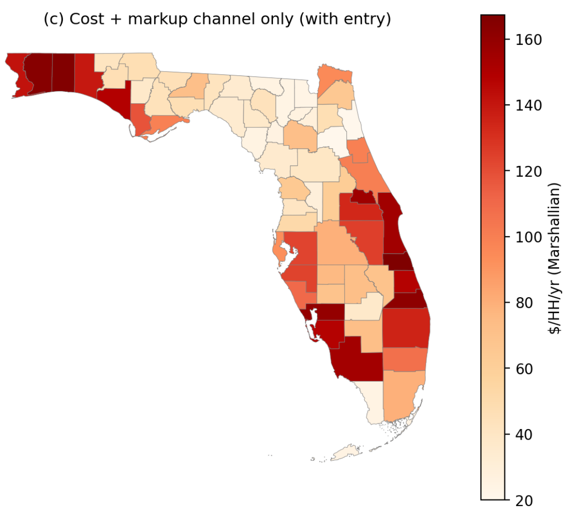
Figure 11: Spatial incidence of the credit-for-reinsurance reform on Florida homeowners.



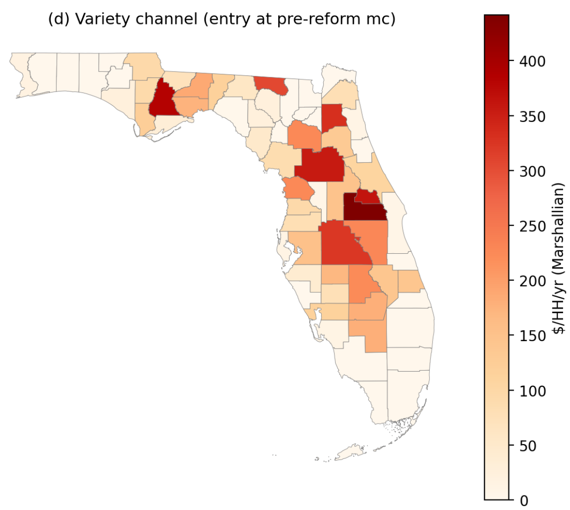
(a) County risk exposure (mean AAL).



(b) Aggregate welfare gain (cost + markup + variety).



(c) Cost + markup channel only.



(d) Variety channel: panel (b) minus panel (c).

Notes: Welfare is reported in dollars per household per year under the Marshallian measure. Per-row contributions are computed on incumbent (non-entrant) products only; the variety channel is read off as $\sum_{j \notin \text{entrants}} s_j^{S_1}(p_j^{S_4} - p_j^{S_3})$, the price reduction incumbent insurers offer when reform-induced entrants compete in the choice set. Markets in which the pre-reform Bertrand simulation reached an off-equilibrium price (max $|\Delta p| > \$10{,}000$/HH/yr) or failed to converge are excluded. An alternative ordering of the chain decomposition—adding cost+markup first at the no-entry choice set, then layering entry on top—is reported in Appendix **??**; the total welfare gain is identical but the per-channel split shifts toward the variety channel because incumbent markups are wider in the no-entry world.

The variety channel favors inland low-risk counties. The spatial pattern of the variety channel is the inverse, as panel 11d of Figure 11 makes visually apparent. Welfare gains from expanded product offerings are largest in inland low-AAL counties and smallest—in some cases negative—in coastal high-AAL counties. Sorted by AAL quintile, the Marshallian variety gain is \$168 per household per year in the lowest quintile, falls to \$30 in the second-highest quintile, and turns negative at $-\$17$ in the highest quintile (cross-county Pearson $r = -0.61$ versus mean AAL). Five counties account for 69 percent of the Florida variety dollar—Orange, Polk, Marion, Seminole, and Hillsborough, predominantly mid-AAL counties along the central-Florida I-4 corridor. The mechanism is not that entry is mechanically concentrated in inland areas: entrant-row shares are 13.4 percent in high-risk counties versus 15.0 percent in low-risk counties, and the cross-county correlation between entry intensity and AAL is essentially zero. Rather, removing entrants matters more in thinner choice sets where the marginal product carries more weight, and coastal incumbents have already absorbed most of the cost shock through markup adjustment, leaving the remaining product set highly substitutable at the share-weighted Marshallian margin.

Net effect: a flatter incidence than either channel predicts. Because the cost-and-markup channels favor coastal high-risk counties while the variety channel favors inland low-risk counties, the two are partial substitutes across space. The total welfare gain therefore exhibits a much flatter spatial incidence than either component would predict in isolation: cross-county Pearson r falls from $+0.33$ for the cost+markup channel (panel 11c) to -0.52 for the full reform gain (panel 11b). The reform's distributional incidence is therefore not a simple coastal-versus-inland story: the cost-side benefits accrue to high-exposure coastal homeowners, while the variety-side benefits accrue to consumers in central Florida markets where the post-reform choice set is thicker and competitive discipline is greatest.

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.

Reduced collateral also left cedents with more unsecured counterparty exposure, although cedents shifted toward rated reinsurers. The reform therefore lowered the cost of transferring catastrophe risk while shifting some protection from collateral toward reinsurer solvency.

We find that the reform cut foreign reinsurance prices roughly in half, 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 12% 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 5%, with three-quarters of that reduction reaching homeowners as lower premiums. The aggregate consumer-surplus gain is roughly $200 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 incrementally 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.

References

AIR Worldwide, "The Coastline at Risk: 2013 Update to the Estimated Insured Value of U.S. Coastal Properties," AIR Worldwide Corporation 2013. Now Verisk Extreme Event Solutions.

Anand, Vaibhav, Douglas Bujakowski, and Kyeonghee Kim, "Regulatory Reform and Resilience: The Effect of Certified Reinsurer Programs on U.S. Insurance Markets," 2026. Working paper.

Arrow, Kenneth J. and Robert C. Lind, "Uncertainty and the Evaluation of Public Investment Decisions," *American Economic Review*, 1970, *60* (3), 364–378.

Berry, Steven, James Levinsohn, and Ariel Pakes, "Automobile Prices in Market Equilibrium," *Econometrica*, 1995, *63* (4), 841–890.

_ , _ , and _ , "Voluntary Export Restraints on Automobiles: Evaluating a Trade Policy," *American Economic Review*, 1999, *89* (3), 400–430.

Berry, Steven T. and Philip A. Haile, "Identification in Differentiated Products Markets Using Market Level Data," *Econometrica*, 2014, *82* (5), 1749–1797.

Bloomberg News, "Hedge Fund Money Is Reshaping a 180-Year-Old Insurance Model," Bloomberg April 2026. April 12, 2026.

Boomhower, Judson, Meredith Fowlie, Jacob Gellman, and Andrew Plantinga, "How Are Insurance Markets Adapting to Climate Change? Risk Classification and Pricing in the Market for Homeowners Insurance," NBER Working Paper 32625, National Bureau of Economic Research 2024.

_ , Patricia Born, Benjamin Collier, and Tobias Huber, "Pricing Climate Risk: Hurricane Models and Home Insurance Over the Last Two Decades," *Available at SSRN 6170366*, 2026.

Born, Patricia H. and Barbara Klimaszewski-Blettner, "Should I Stay or Should I Go? The Impact of Natural Disasters and Regulation on U.S. Property Insurers' Supply Decisions," *Journal of Risk and Insurance*, March 2013, *80* (1), 1–36.

_ and _ , "Should I Stay or Should I Go? The Impact of Natural Disasters and Regulation on U.S. Property Insurers' Supply Decisions," *Journal of Risk and Insurance*, 2013, *80* (1), 1–36.

Borusyak, Kirill and Peter Hull, "Nonrandom Exposure to Exogenous Shocks," *Econometrica*, 2023, *91* (6), 2155–2185.

_ , _ , and Xavier Jaravel, "Quasi-Experimental Shift-Share Research Designs," *Review of Economic Studies*, 2022, *89* (1), 181–213.

__ **, Xavier Jaravel, and Jann Spiess**, "Revisiting Event-Study Designs: Robust and Efficient Estimation," *Review of Economic Studies*, 2024, *91* (6), 3253–3285.

Carrillo, Gabriel, Daniel Telljohann, and Charles Nyce, "The 30th Anniversary of Hurricane Andrew: Evolution of the Florida Homeowners Insurance Market," *Risk Management and Insurance Review*, 2022, *25* (3), 239–270.

CAS Risk-Based Capital Dependencies and Calibration Working Party, "Report 3: Solvency II Standard Formula and NAIC Risk-Based Capital (RBC)," CAS E-Forum, Fall 2012, Volume 2, Casualty Actuarial Society 2012.

Chamberlain, Gary, "Panel Data," *Handbook of Econometrics*, 1984, *2*, 1247–1318.

Cole, Cassandra R, Kathleen A McCullough, and Lawrence S Powell, "Collateralization of international reinsurance liabilities in the US insurance industry," *Insurance markets and companies: analyses and actuarial computations*, 2010, (1, Iss. 2), 30–36.

Cummins, J. David, "Should the Government Provide Insurance for Catastrophes?," *Federal Reserve Bank of St. Louis Review*, 2006, *88* (4), 337–379.

__ , "The Bermuda Insurance Market: An Economic Analysis," Report, Bermuda Insurance Development Council, Hamilton, Bermuda May 2008.

__ **and Mary A. Weiss**, "Convergence of Insurance and Financial Markets: Hybrid and Securitized Risk-Transfer Solutions," *Journal of Risk and Insurance*, 2009, *76* (3), 493–545.

de Chaisemartin, Clément and Xavier D'Haultfœuille, "Two-Way Fixed Effects Estimators with Heterogeneous Treatment Effects," *American Economic Review*, September 2020, *110* (9), 2964–96.

Federal Insurance Office, U.S. Department of the Treasury, "The Breadth and Scope of the Global Reinsurance Market and the Critical Role Such Market Plays in Supporting Insurance in the United States," Report, U.S. Department of the Treasury, Washington, DC December 2014. Completed pursuant to Title V of the Dodd-Frank Wall Street Reform and Consumer Protection Act.

Florida Office of Insurance Regulation, "Florida Residential Property Market Share Report," Quarterly QUASR data, compiled by Citizens Property Insurance Corporation 2024.

Froot, Kenneth A., "The Market for Catastrophe Risk: A Clinical Examination," *Journal of Financial Economics*, 2001, *60* (2-3), 529–571.

__ , "Risk Management, Capital Budgeting, and Capital Structure Policy for Insurers and Reinsurers," *Journal of Risk and Insurance*, 2007, *74* (2), 273–299.

__ **and Jeremy C. Stein**, "Risk Management, Capital Budgeting, and Capital Structure Policy for Financial Institutions: An Integrated Approach," *Journal of Financial Economics*, 1998, *47* (1), 55–82.

— **and Paul G. J. O'Connell**, "On the Pricing of Intermediated Risks: Theory and Application to Catastrophe Reinsurance," *Journal of Banking & Finance*, 2008, *32* (1), 69–85.

Gallagher Re, "Florida Market Watch: Year End Statutory 2025 Results," Report April 2026. Reports reinsurance utilization as ceded written premium over gross written premium (CWP/GWP) for Florida personal-property-focused insurers.

Gandhi, Amit and Jean-François Houde, "Measuring Substitution Patterns in Differentiated-Products Industries," *NBER Working Paper No. 26375*, 2019.

Gardner, John, "Two-Stage Differences-in-Differences," Technical Report, arXiv 2022. arXiv:2207.05943.

Ge, Shan, "How Do Financial Constraints Affect Product Pricing? Evidence from Weather and Life Insurance Premiums," *Journal of Finance*, 2022, *77* (1), 449–503.

Goodman-Bacon, Andrew, "Difference-in-differences with variation in treatment timing," *Journal of Econometrics*, 2021, *225* (2), 254–277. Themed Issue: Treatment Effect 1.

Grace, Martin F., Robert W. Klein, and Zhiyong Liu, "Increased Hurricane Risk and Insurance Market Responses," *Journal of Insurance Regulation*, 2007, *26* (2), 3–32.

Gron, Anne, "Capacity Constraints and Cycles in Property-Casualty Insurance Markets," *RAND Journal of Economics*, 1994, *25* (1), 110–127.

Jia, Ruo, Jieyu Lin, Michael R. Powers, and Hanyang Wang, "Catastrophe Risk Sharing among Individuals, Private Insurance, and Government," *Journal of Risk and Insurance*, 2025, *92* (2), 263–311.

Keys, Benjamin J. and Philip Mulder, "Property Insurance and Disaster Risk: New Evidence from Mortgage Escrow Data," NBER Working Paper 32579, National Bureau of Economic Research 2024.

Koijen, Ralph S. J. and Motohiro Yogo, "The Cost of Financial Frictions for Life Insurers," *American Economic Review*, 2015, *105* (1), 445–475.

— **and** — , "Shadow Insurance," *Econometrica*, 2016, *84* (3), 1265–1287.

Kousky, Carolyn and Roger Cooke, "Explaining the Failure to Insure Catastrophic Risks," *Geneva Papers on Risk and Insurance—Issues and Practice*, 2012, *37* (2), 206–227.

Lewis, Christopher M. and Kevin C. Murdock, "The Role of Government Contracts in Discretionary Reinsurance Markets for Natural Disasters," *Journal of Risk and Insurance*, 1996, *63* (4), 567–597.

Mundlak, Yair, "On the Pooling of Time Series and Cross Section Data," *Econometrica*, 1978, *46* (1), 69–85.

National Association of Insurance Commissioners, "Credit for Reinsurance Model Law," NAIC Model Laws, Regulations, Guidelines and Other Resources 2019. Model #785; originally adopted 1984, amended November 2011 and June 2019.

Niehaus, Greg, "The Allocation of Catastrophe Risk," *Journal of Banking & Finance*, 2002, *26* (2–3), 585–596.

NOAA, "U.S. Billion-Dollar Weather and Climate Disasters," 2026. National Oceanic and Atmospheric Administration, https://www.ncei.noaa.gov/access/billions/.

Oh, Sangmin, Ishita Sen, and Ana-Maria Tenekedjieva, "Pricing of Climate Risk Insurance: Regulation and Cross-Subsidies," *Journal of Finance*, 2026. Forthcoming.

_ **, Parinitha Sastry, and Ishita Sen**, "Climate Risk, Insurance, and Mortgage Markets," Working Paper, Available via NBER Reporter, 2025 No. 2 2024. Title and venue to be verified; see NBER Reporter overview at https://www.nber.org/reporter/2025number2/housing-climate-risk-and-insurance.

Organisation for Economic Co-operation and Development, "OECD Tax Database: Statutory Corporate Income Tax Rates," https://www.oecd.org/tax/tax-policy/tax-database/ 2024. Accessed October 2025.

Sastry, Parinitha, Ishita Sen, and Ana-Maria Tenekedjieva, "When Insurers Exit: Climate Losses, Fragile Insurers, and Mortgage Markets," 2024. Working paper.

Solomon, Adam, "When Insurance Markets Fail: Catastrophe-Risk Frictions and Public Reinsurance," 2026. Working Paper.

Weyl, E Glen and Michal Fabinger, "Pass-through as an economic tool: Principles of incidence under imperfect competition," *Journal of political economy*, 2013, *121* (3), 528–583.

Winter, Ralph A., "The Dynamics of Competitive Insurance Markets," *Journal of Financial Intermediation*, 1994, *3* (4), 379–415.

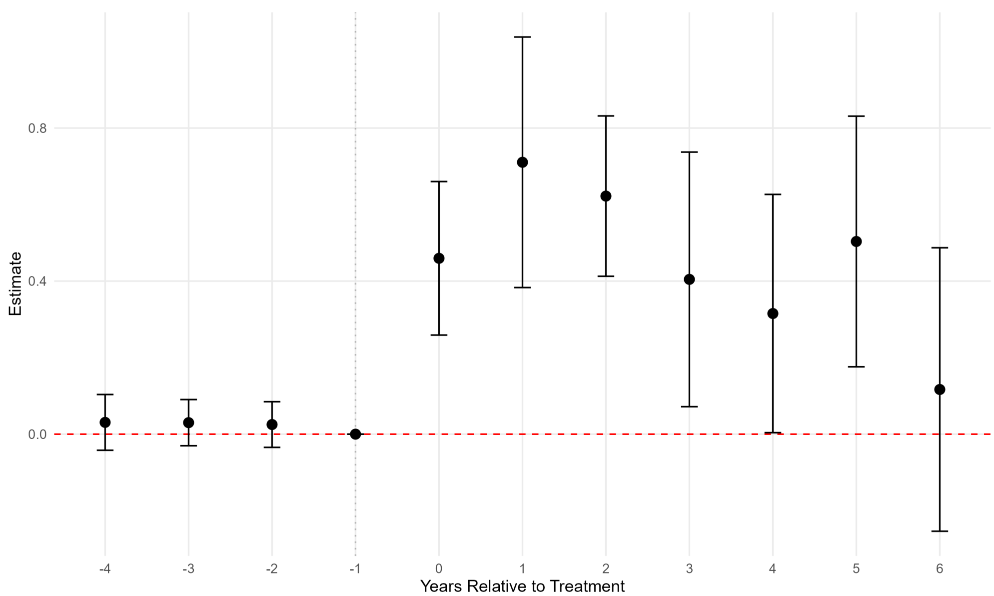
Zanjani, George, "Pricing and Capital Allocation in Catastrophe Insurance," *Journal of Financial Economics*, 2002, *65* (2), 283–305.

A Additional National Reinsurance Market Results

A.1 Domestic Reinsurer Results

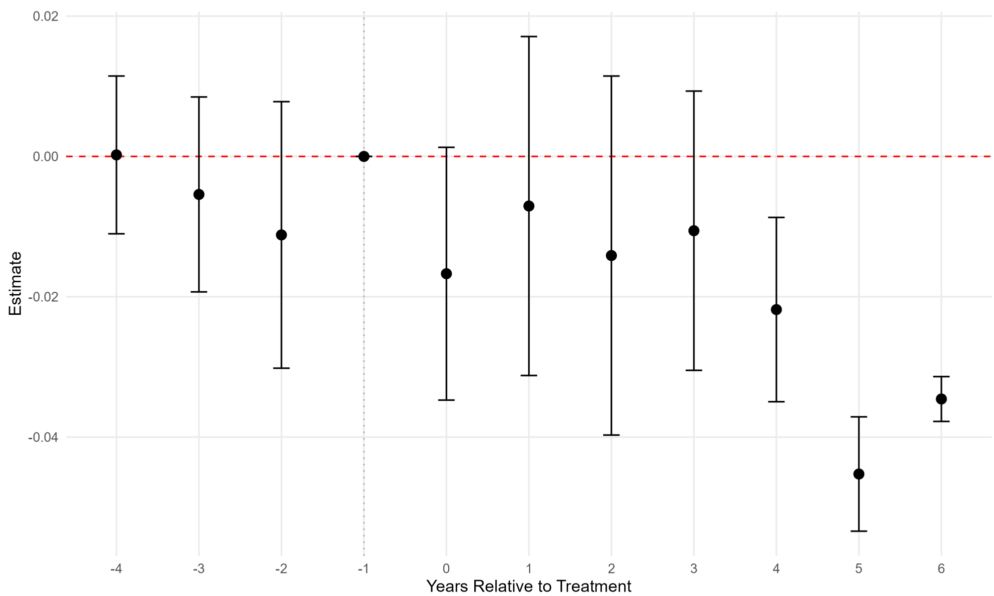
This appendix presents the national reinsurance market results for domestic reinsurers, complementing the foreign reinsurer results in Section 3.2.1.

Figure 12: 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 τ 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 13: Reinsurance Prices: Domestic Reinsurers

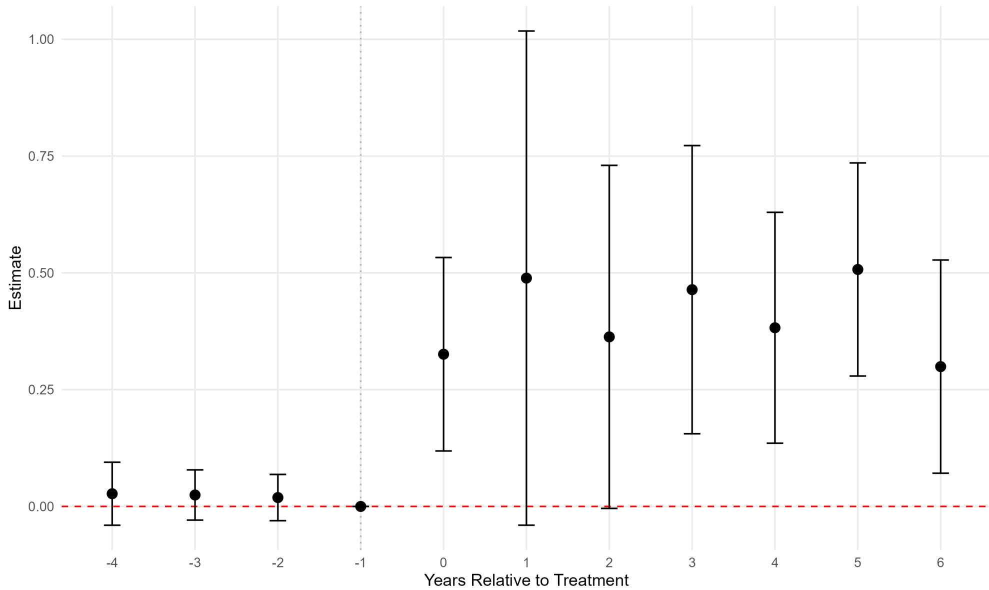


Notes: The figure reports a price slope specification event study estimated on an insurer-by-year balanced panel. The outcome is losses plus loss adjustment expenses (LAE) incurred on reinsurance ceded to domestic reinsurers; the exposure variable is premiums ceded to domestic reinsurers. The sample is restricted to insurer-years with positive ceded premiums to domestic reinsurers. Event time τ 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.

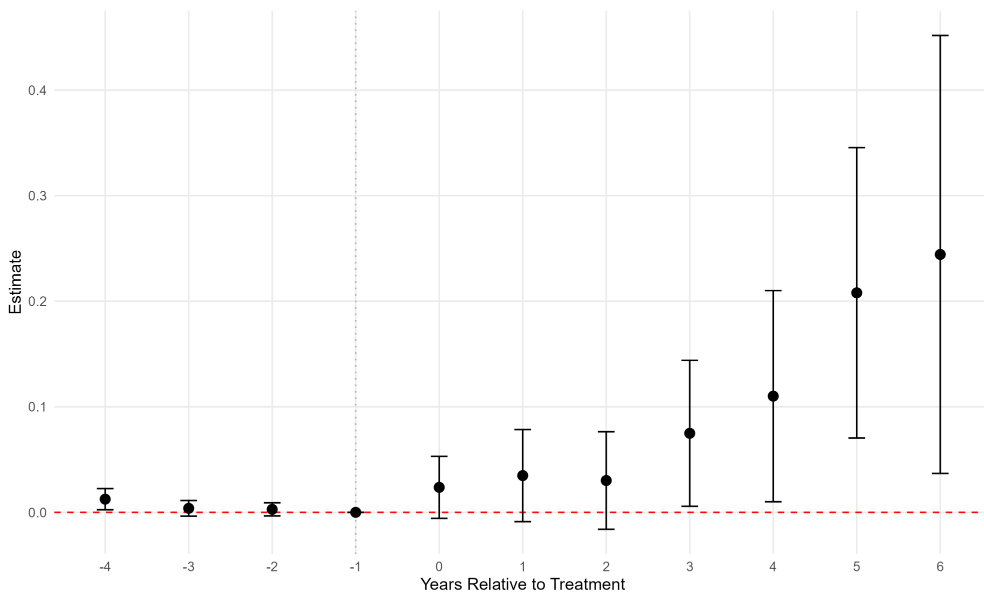
A.2 Affiliated Reinsurer Decomposition

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 14 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 14: Quantity of Reinsurance Ceded: Affiliated vs. Non-Affiliated Reinsurers



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

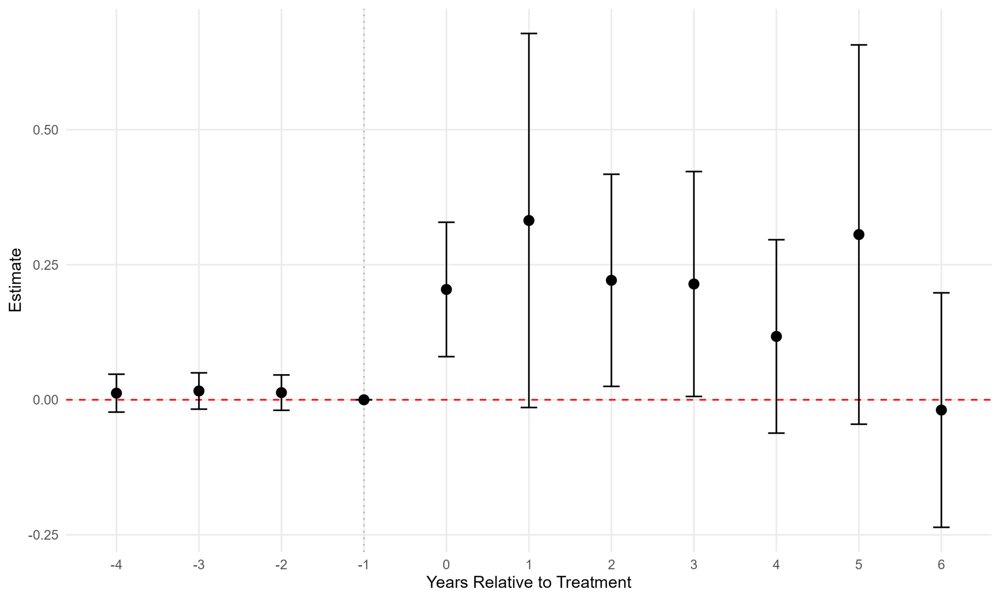


(b) Premiums ceded to non-affiliated reinsurers as a function of DPW

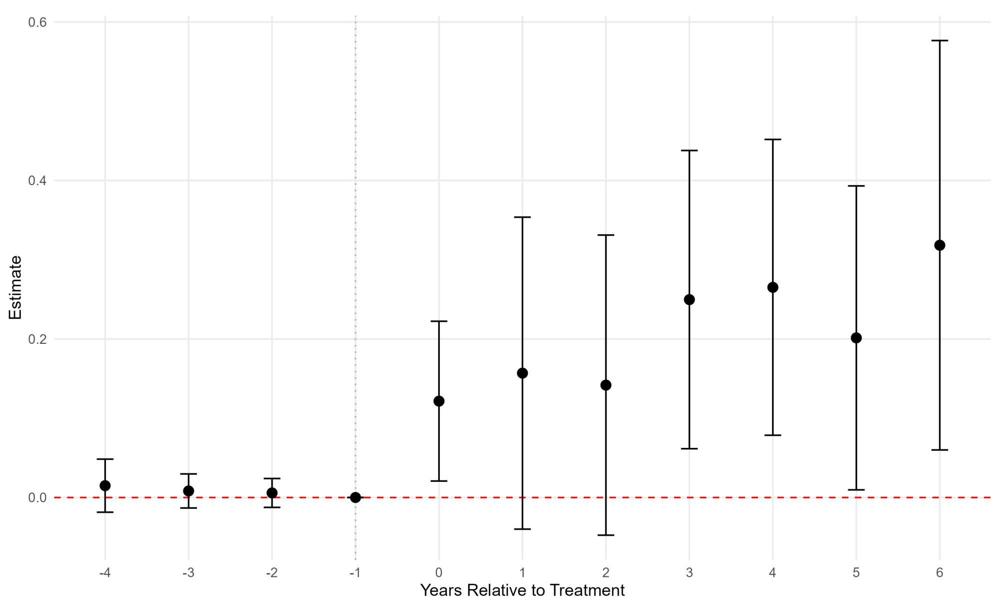
Notes: Each panel reports a slope specification event study estimated on an insurer-by-year balanced panel restricted to insurer DPW > 0. The outcome is dollars of reinsurance ceded to affiliated (panel a) or non-affiliated (panel b) reinsurers; the exposure variable is direct premiums written (DPW). Affiliated reinsurers are entities created primarily to reinsure affiliated companies; non-affiliated reinsurers are traditional commercial reinsurance companies. Event time τ 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 15: 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 τ 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 3.2.3 and reports the full set of estimates and sensitivity analyses.

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, and assign the uncollateralized share λ_{ijt} from the statutory collateral schedule: domestic and U.S.-authorized reinsurers post no collateral ($\lambda = 1$); non-admitted foreign reinsurers posted full collateral before the cedent's domicile state adopted Model Law 785 ($\lambda = 0$) and the post-adoption certified level thereafter, anchored to the observed post-2018 uncollateralized share by rating tier. Expected impairment loss is $\text{EL}_{it} = \text{LGD} \sum_j p(r_j^*)\lambda_{ijt}R_{ijt}$, with $\text{LGD} = 1 - \rho$ 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).

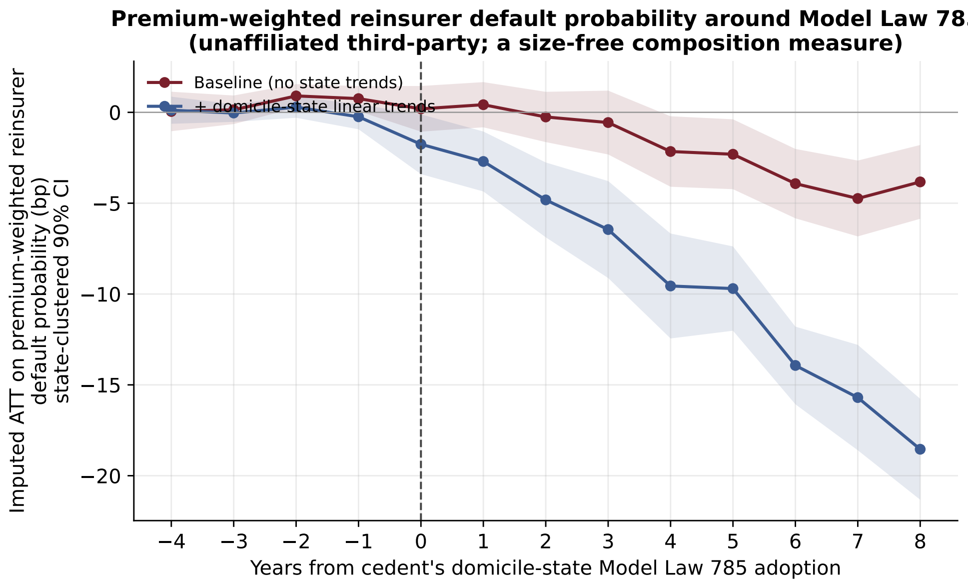
Table 6: Effect of Model Law 785 on cedents' reinsurer impairment exposure

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).

Event studies. Figure 16 reports the event study for the premium-weighted impairment probability, which declines after adoption; correspondingly, the share of premium ceded to unrated reinsurers falls by approximately 1.7 percentage points ($p = 0.02$). Cedents move toward rated counterparties, so the rise in dollar expected losses reflects the larger and less-collateralized exposure rather than a deterioration in average counterparty quality.

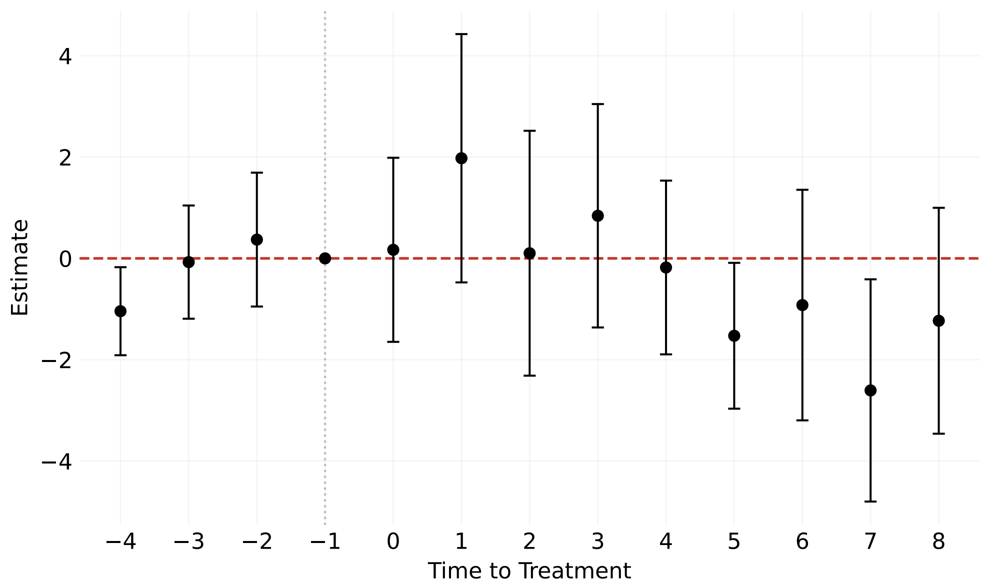
Figure 16: Premium-Weighted Reinsurer Impairment Probability



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. The dollar measure rises in part because treated cedents buy more reinsurance. To isolate whether each dollar of reinsurance became riskier, we scale expected impairment loss by the cedent's total reinsurance premium ceded and re-estimate. The effect is an imprecisely estimated $-\$0.21$ per \$1,000 of premium ceded (s.e. $\$0.86$, $p = 0.81$; pre-reform mean about \$50 per \$1,000), economically and statistically indistinguishable from zero (Figure 17). The rise in dollar expected losses therefore reflects a larger, less-collateralized book rather than a deterioration in expected loss per dollar of reinsurance.

Figure 17: 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 τ 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 6. Standard errors are clustered on cedent domicile state, and bands are 90% 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.

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, complementing the main wind premium results in Section 4.2.1.

Figure 18: Policy Pricing: High vs. Low Risk

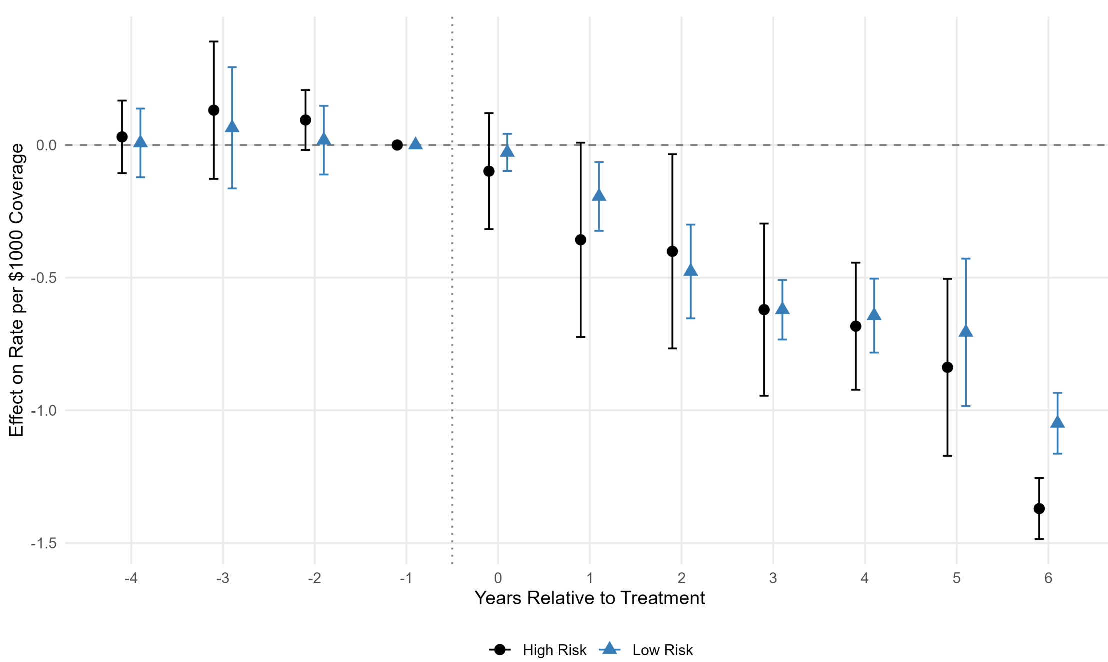
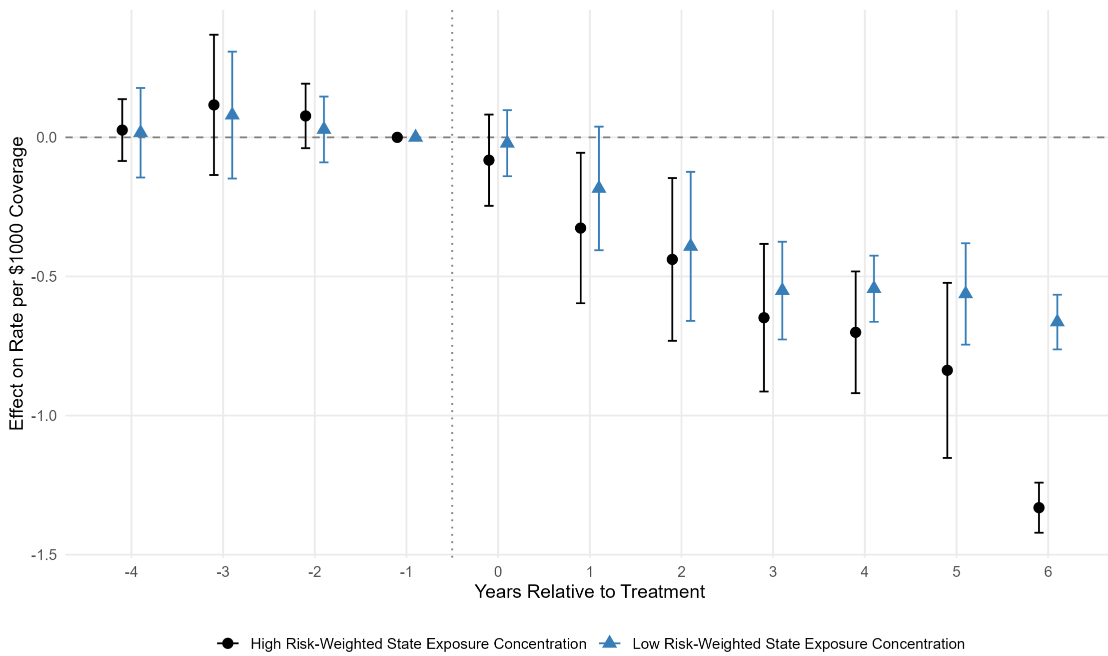


Figure 19: Policy Pricing: High vs. Low Expected Loss Counties



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 20 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 21 below, show that if anything the price effects are stronger among incumbents.

Figure 20: Placebo: Wind vs. Non-Wind Rate Levels

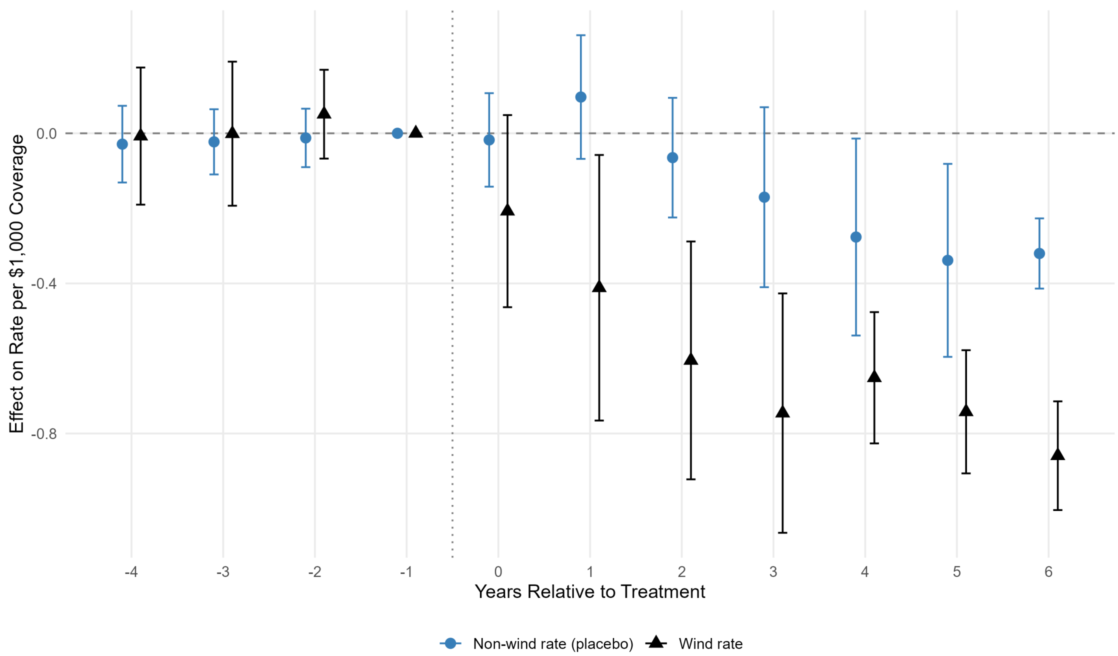
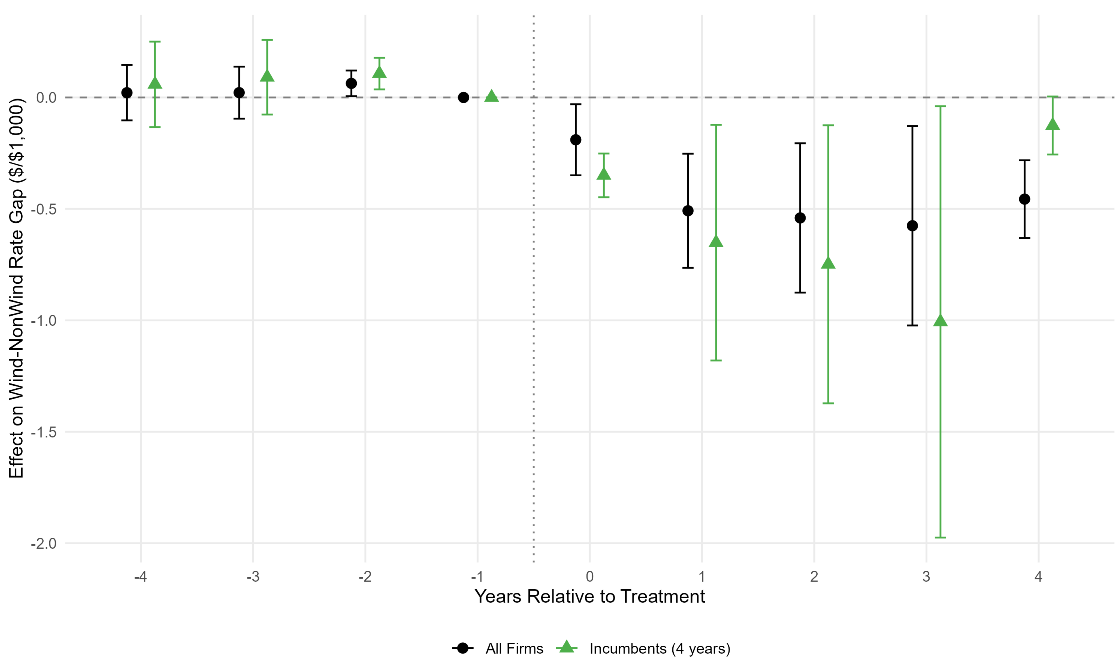


Figure 21: Wind Triple Difference: Incumbents vs. All Firms

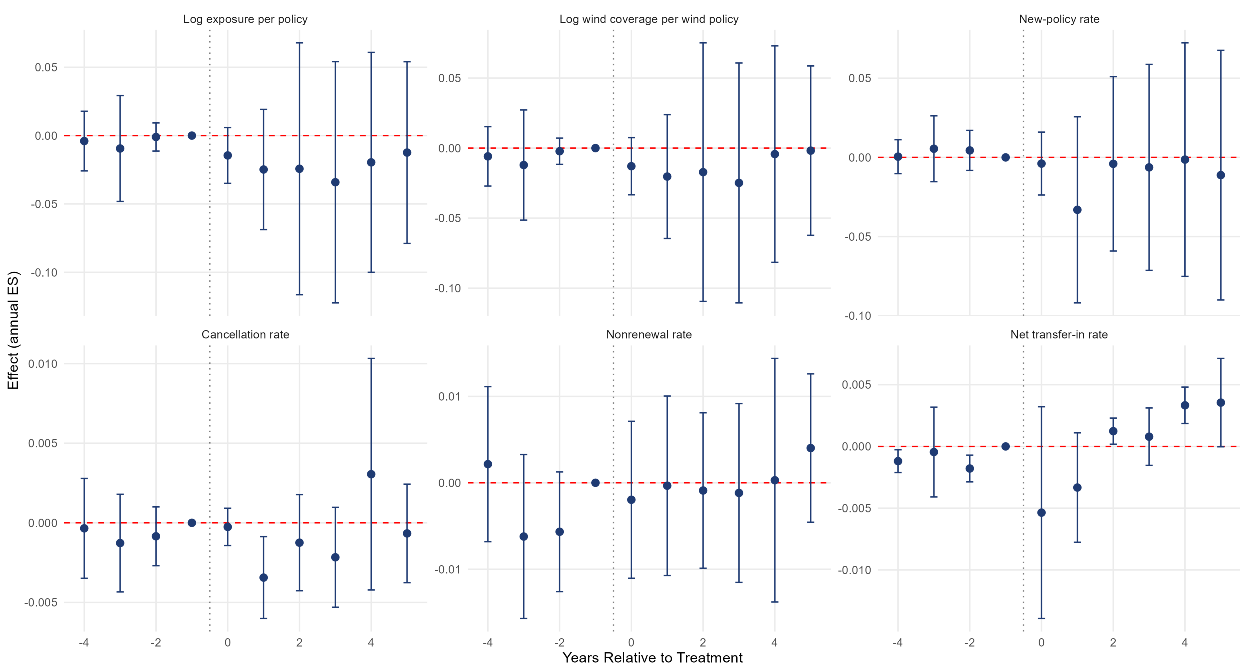


C.4 Unchanged Policy Composition

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 primary Figure 8.

To test this, we study whether the risk (exposure per policy), wind-exposure, or flow of new business change with the reform. Figure 22 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.

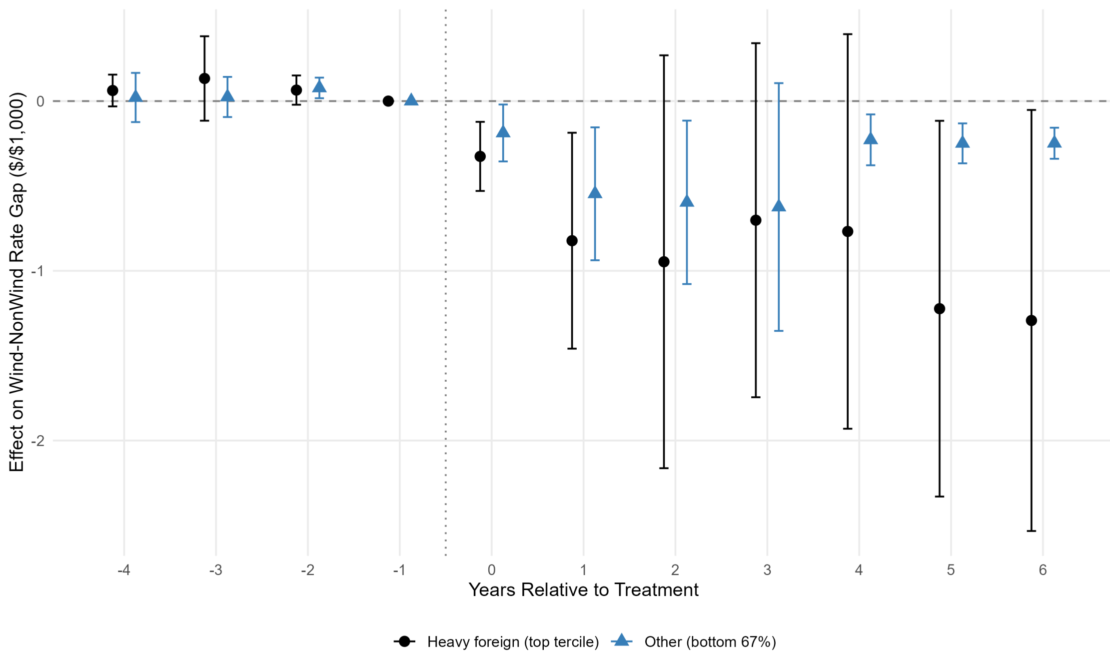
Figure 22: Compositional-Change Robustness: Coverage and Policy-Flow Outcomes



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 23 shows the wind-rate decline concentrates among heavy foreign-reinsurance users, consistent with the collateral reform specifically targeting the cost of foreign reinsurance.

Figure 23: 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