Antitrust in the Age of Intellectual Property: Licensing, Merger Approval, and Product Prices

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Abstract

How do firms respond to antitrust oversight when competition increasingly centers on intellectual property? We study whether firms strategically deploy IP licensing to navigate merger review and how such licensing shapes market outcomes. IP licensing rises sharply in reportable mergers, particularly in horizontal transactions, concentrated industries, and settings with heightened enforcement risk. Using granular product-level data, we show that licensing-facilitated mergers are followed by significantly higher prices, especially where competitive overlap is greatest. Licensing also increases completion rates and accelerates closing for challenged deals. Consistent with a strategic mechanism, firms license lower-value patents under less restrictive terms, limiting economic concessions. Together, the evidence suggests that IP licensing can undermine the effectiveness of antitrust regulation, allowing firms to secure approval while preserving the very market power the regulation seeks to constrain.

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1. Introduction

Regulation seeks to preserve competition, yet firms can strategically structure transactions to offset, dilute, or circumvent its intended effects. This tension is particularly acute in mergers and acquisitions, where firms have substantial discretion over deal design, while antitrust authorities evaluate anticipated competitive effects under time pressure and informational asymmetries. In particular, firms can shape both the economic substance and the contractual form of a deal in ways that influence regulatory assessment. These strategic considerations have become even more consequential as competition increasingly centers on innovation, intangible capital, and intellectual property (e.g., De Loecker, Eeckhout, and Unger (2020)). As a result, the allocation and deployment of technology assets, through licensing, divestitures, and other contractual arrangements, have moved to the forefront of mergers and antitrust policy.

In this paper, we study how firms strategically deploy intellectual property (IP) licensing under antitrust oversight and assess the real economic consequences of such actions. We find that licensing activity rises sharply when transactions are subject to antitrust oversight, particularly in horizontal transactions, highly concentrated industries, and settings characterized by elevated enforcement risk. The increase is concentrated in lower-quality patents, suggesting that licensing commitments are structured to satisfy regulatory demands at minimal competitive cost to the acquirer. Using granular product-level pricing data, we further show that mergers accompanied by licensing commitments are followed by significantly higher prices. Taken together, the evidence indicates that IP licensing operates as a strategic tool through which firms dilute regulatory competitive pressures.

A large literature studies the competitive effects of mergers and the role of antitrust enforcement, beginning with Williamson (1968) and Farrell and Shapiro (1990). A central and unresolved question concerns the effectiveness of antitrust interventions, such as divestitures, designed to mitigate competitive harm (e.g., Kwoka (2014)). By contrast, despite the growing centrality of intellectual property in modern competition, there is little empirical evidence on the use and effectiveness of IP-based interventions. We address this gap by providing causal evidence on how firms deploy IP licensing and how such actions shape subsequent market outcomes.

We aim to answer four related questions. First, does intellectual property licensing increase when a merger is subject to antitrust scrutiny? In particular, is licensing more likely in reportable transactions and in deals characterized by heightened enforcement risk, such as large horizontal mergers in concentrated industries or transactions that attract regulatory and public attention? Second, does licensing effectively facilitate regulatory approval? Specifically, are challenged mergers that include licensing commitments more likely to complete and to close more quickly? Third, conditional on completion, how does licensing affect competitive outcomes? Do mergers accompanied by licensing commitments lead to higher product prices, particularly in overlapping and highly concentrated markets where antitrust concerns are strongest? Fourth, what types of intellectual property are licensed, and under what contractual terms? Are the patents licensed in scrutinized transactions lower in value or quality, and do royalty rates, duration, and exclusivity suggest minimal competitive transfer?

To answer these questions, we assemble new transaction-level and product-level datasets that link intellectual property (IP) licensing contracts to merger transactions and to post-merger product prices. We obtain detailed licensing contracts from RoyaltyStat's *Royalty Rate* and *Licensing Agreement* databases, augmented with exhibits filed with the United States Securities and Exchange Commission (SEC), documents obtained through Freedom of Information Act (FOIA) unredactions, and inter-firm cooperation records from the Securities Data Company (SDC). We retrieve information about merger transactions from the SDC Mergers and Acquisitions database. We obtain detailed retail product prices from NielsenIQ's Retail Scanner Data, which provide weekly transaction-level information on prices and quantities sold at the Universal Product Code (UPC)–store level across U.S. grocery, drugstore, mass merchandise, and other retail channels.

We provide four sets of empirical results that jointly characterize the strategic use of IP licensing in merger transactions and evaluate its economic consequences. First, we examine whether IP licensing increases in response to antitrust exposure. Exploiting variation in the Hart–Scott–Rodino (HSR) reporting threshold, we compare licensing behavior following reportable and non-reportable transactions. We find that firms announcing reportable deals are 4.1 percentage points more likely to enter into licensing agreements in the following year, representing a 50% increase relative to the unconditional licensing probability of 8%. The effect is statistically significant at the 1% level and remains robust after including firm fixed effects, industry-by-year fixed effects, and quantic polynomials of firm size. In contrast, firms announcing only non-reportable deals exhibit no significant change in licensing behavior. This pattern suggests that licensing is not a generic byproduct of acquisition activity or the scale of a firm's operations, but is concentrated precisely when transactions trigger antitrust scrutiny.

In dynamic lead–lag analyses, we find no differential pre-trends in licensing behavior prior to reportable deal announcements, alleviating concerns about reverse causality or persistent firm characteristics. Licensing increases sharply between deal announcement and completion, and not before announcement or after deal closure, indicating that firms adjust IP deployment specifically during periods of regulatory uncertainty. A difference-in-differences analysis around the 2001 HSR threshold increase (e.g., Wollmann (2019)) further confirms a causal interpretation: transactions that became exempt from mandatory review after the reform exhibit a significant decline in licensing activity relative to unaffected deals.

Second, we exploit cross-sectional variation in antitrust exposure and merger outcomes to investigate whether licensing is concentrated in high antitrust risk settings and whether it matters for regulatory outcomes. The effects are driven by the largest reportable deals. Among reportable transactions exceeding $1 billion, the likelihood of licensing increases by additional 3.2 percentage points relative to smaller reportable transactions, implying a 40% increase relative to the unconditional licensing rate of 8%. We find no comparable licensing response among smaller reportable or non-reportable transactions.

We find similar amplified effects for horizontal mergers and in highly concentrated industries, where traditional competitive concerns are front and center (e.g., Eckbo (1983, 1985); Stillman (1983); Eckbo and Wier (1985); Fee and Thomas (2004); Shahrur (2005); Fathollahi, Harford, and Klasa (2022); Liebersohn (2024)). Specifically, horizontal deals are 1.7 percentage points more likely to be accompanied by licensing relative to non-horizontal reportable deals, and deals in industries with above-median concentration exhibit licensing rates that are 2.2 percentage points higher than those in less concentrated markets. Licensing responses are also stronger when scrutiny is elevated: industries with recent peer challenges and deals receiving high media or social media attention exhibit incremental increases of 6, 2.1, and 2.2 percentage points, respectively. Across specifications, the magnitudes represent economically meaningful increases relative to the 8% baseline rate.

We next examine whether licensing affects regulatory outcomes using deal-level regressions that include acquirer controls, deal characteristics, industry-by-year fixed effects, and firm fixed effects. Among transactions formally challenged by antitrust authorities, licensing materially increases the probability of completion. The interaction between *Licensing* and *Challenged* implies an increase in completion probability of approximately 9.7 percentage points relative to challenged deals without licensing. Licensing also accelerates resolution: conditional on completion, challenged deals accompanied by licensing close roughly 43% faster relative to challenged deals without licensing. In economic terms, this corresponds to a reduction in review time of approximately 227 days relative to the average 532-day review period for challenged reportable transactions. These results indicate that licensing affects outcomes precisely when regulatory constraints bind.

Taken together, these findings show that licensing both scales with the underlying determinants of enforcement risk and materially affects regulatory outcomes when scrutiny arises. Licensing is concentrated in large horizontal mergers, in concentrated markets, and in transactions attracting elevated attention, and it increases completion probabilities and reduces review duration precisely in challenged deals. This alignment suggests that the deployment of intellectual property is closely intertwined with the merger review process. In innovation-driven markets, firms appear to adjust the allocation of intangible assets along the margins that regulators monitor most closely, highlighting how antitrust scrutiny shapes transaction design. The fact that licensing has no effect in non-challenged transactions further suggests that its value arises from regulatory interaction rather than standalone transactional efficiency.

In the third set of analyses, we turn to the central question of real economic effects and, critically, to distinguishing between competing interpretations of licensing. Licensing could reflect benign technology sharing or efficiency-enhancing collaboration, in which case mergers accompanied by licensing should exhibit muted or neutral price effects. Alternatively, licensing may function primarily as a strategic tool that facilitates approval while leaving underlying competitive forces intact or even enabling greater market power. These competing views generate sharply different predictions for post-merger pricing.

To distinguish between these views, we examine post-merger product prices. Using NielsenIQ's granular Universal Product Code (UPC)–level pricing data covering 3.5 million product–state observations across 497 completed deals, we examine price changes over the period from two years prior to the deal announcement to two years after the deal completion. We find that reportable mergers accompanied by IP licensing are followed by a 1.1% increase in prices relative to comparable reportable deals without licensing. Given the sample's average post-merger price change of 1.89%, this represents a 58% amplification in price growth. The estimates are statistically significant at the 1% level and robust to product-by-state fixed effects. In dynamic event-time analyses, we find no differential pre-trends and a sharp divergence following deal completion, indicating that licensing facilitates mergers that subsequently exercise greater pricing power.

Moreover, the pricing effects are economically concentrated in precisely those settings where antitrust concerns are strongest. Post-merger price increases are larger for products that overlap directly with target offerings, for transactions in highly concentrated industries, and when enforcement salience is elevated due to recent peer challenges or intense media scrutiny. These heterogeneous effects align with market-power interpretations in the merger literature (e.g., Ashenfelter and Hosken (2010); Fathollahi, Harford, and Klasa (2022); Kepler, Naiker, and Stewart (2023)). Importantly, licensing does not appear to neutralize competitive harm in these settings; instead, price effects are amplified when competitive overlap is greatest.

The pricing results suggest that licensing does not fully neutralize competitive harm and may instead facilitate mergers that subsequently exercise greater market power. This pattern naturally shifts attention to the nature of the licensing itself. The strategic view predicts selective concessions: firms license patents that satisfy review requirements while preserving economically important intellectual assets. In contrast, an efficiency-based interpretation would predict licensing of high-value technologies that meaningfully sustain competition. To distinguish between these mechanisms, in the final set of analyses, we examine the economic value and contractual structure of licensed patents.

Using stock-market-based patent valuation measures (Kogan et al. (2017)), we show that patents licensed in reportable transactions are economically and statistically lower in value than patents licensed outside the merger context. The estimated coefficients imply that patents licensed in reportable deals are approximately 40–42% lower in market value relative to other licensed patents. These effects are statistically significant at conventional levels and remain robust to the inclusion of firm fixed effects, indicating that they are not driven by persistent differences in firms' innovation portfolios. By contrast, licensing associated with non-reportable transactions exhibits no systematic difference in patent value. The magnitudes indicate that acquirers facing antitrust scrutiny strategically license economically marginal patents, preserving their highest-value intellectual assets while satisfying regulatory demands.

Political connections further shape licensing strategies. Politically connected acquirers, measured using lobbying expenditures and political contributions (Cooper, Gulen, and Ovtchinnikov (2010); Hong and Kostovetsky (2012); Fidrmuc, Roosenboom, and Zhang (2018)), license patents of significantly lower economic value in reportable transactions relative to non-connected firms. The estimates imply that, conditional on being in a reportable deal, politically connected firms license patents that are an additional 29–52% lower in value relative to otherwise similar non-connected acquirers. Combining the baseline and interaction effects, patents licensed by connected firms in reportable transactions are approximately 60–76% lower in market value than patents licensed outside the merger context.

These estimates suggest that political capital operates as a substitute for substantive intellectual property concessions. Firms with stronger regulatory access appear able to satisfy antitrust scrutiny while relinquishing even less economically meaningful technology, retaining their highest-value assets while still securing approval.

Finally, we examine contract-level licensing terms. Licensing agreements associated with reportable transactions feature economically lower royalty rates, shorter durations, and weaker exclusivity provisions relative to licenses unrelated to M&A activity. Lower royalty rates limit the immediate transfer of rents; shorter contract horizons restrict longer-term competitive exposure; and weaker exclusivity provisions imply less restrictive commitments. Across all contractual dimensions, the structure of licensing agreements points toward narrower and less economically consequential concessions. Consistent with the evidence on patent value, firms under antitrust scrutiny appear to calibrate both the assets they license and the terms under which they license them to preserve core competitive advantages while satisfying regulatory demands.

Overall, across all margins (what is licensed, who licenses it, and how it is licensed), the pattern is consistent: firms calibrate IP concessions to satisfy regulatory scrutiny while preserving competitive advantages. Licensing in this setting appears not as a broad mechanism for technology diffusion, but as a targeted instrument deployed to navigate antitrust review at a minimal economic cost to acquirers and a considerable economic cost to consumer surplus.

This paper contributes to several strands of literature. First, it contributes to research on mergers and antitrust enforcement. Prior work documents that mergers, particularly horizontal deals, can lead to higher prices and less competition (e.g., Kim and Singal (1993); Ashenfelter and Hosken (2010); Fathollahi, Harford, and Klasa (2022); Bonaimé and Wang (2024)). Several studies examine how firms circumvent regulatory oversight by fragmenting deals to avoid reporting thresholds (Wollmann (2019); Kepler, Naiker, and Stewart (2023)) or by exploiting accounting discretion to mask competitive harm (Kepler, McClure, and Stewart (2025)). We advance this literature by showing that firms strategically deploy IP licensing to facilitate approval and accelerate deal completion in innovation-driven markets.

Second, the paper contributes to theliterature on market power and intangible assets in modern economies (De Loecker, Eeckhout, and Unger (2020)). While prior research emphasizes the role of M&A in relaxing financial constraints on R&D (Phillips and Zhdanov (2013)), realizing technological synergies (Bena and Li (2014)), or preempting nascent competition through "killer acquisitions" (Cunningham, Ederer, and Ma (2021)), we introduce a regulatory dimension to the deployment of intangible assets. We demonstrate that intellectual property functions not only as a productive input but as a strategic instrument for navigating the market for corporate control.

Third, the paper contributes to the literature on corporate political strategy under regulation. Prior studies document the role of political connections in corporate enforcement intensity and merger regulation (e.g., Correia (2014); Lambert (2019); Fidrmuc, Roosenboom, and Zhang (2018); Mehta, Srinivasan, and Zhao (2020)). We extend this literature by showing that political capital influences not just the likelihood of enforcement, but the magnitude of firms' economic concessions through the strategic licensing of lower-value IP.

More broadly, the paper highlights an underexplored dimension of regulatory design: when enforcement relies on negotiated commitments involving complex intangible assets, firms may retain substantial discretion over the economic substance of those commitments. Understanding this strategic interaction between firms and regulators is central to evaluating the efficacy of competition policy in innovation-driven markets.

2. Institutional Background on Antitrust Process and IP Licensing

2.1 Antitrust regulation and pre-merger notification program

The Hart-Scott-Rodino Antitrust Improvements Act of 1976 (HSR Act) established the federal premerger notification program, under which the parties to certain mergers or acquisitions are required to notify the DOJ and the FTC before consummating the proposed transaction. This pre-merger notification program plays a pivotal role in mitigating antitrust risks by allowing regulatory agencies to review potentially anticompetitive transactions prior to their consummation. Without such provisions, the enforcement agencies have historically found it costly and often infeasible to fully restore competition once a merger has been completed.

The HSR Act initially mandated filings for all transactions exceeding a $15 million threshold. However, in an effort to reduce the regulatory burden on smaller, less economically significant transactions, a major amendment was passed, effective February 1, 2001. This amendment raised this value to $50 million under the so-called size-of-transaction test. [1] Since this regulatory change substantially altered the scope of mandatory review, our empirical analysis focuses on acquisitions announced from February 2001 onward, thereby avoiding inconsistencies associated with earlier filing requirements.

Figure 1 illustrates the statutory timeline of the HSR premerger notification process. Once both the acquirer and target submit complete filings, the HSR Act triggers an initial waiting period of 30 days (or 15 days for cash tender offers and bankruptcies). During this period, regulatory

[1] Since 2004, the HSR reporting threshold has been adjusted annually based on changes in gross national product and reached $126.4 million in 2025 (https://www.ftc.gov/enforcement/competition-matters/2025/02/new-hsr-thresholds-filing-fees-2025). For transactions exceeding the size-of-transaction threshold but below a higher cutoff (e.g., $200 million in 2001), pre-merger antitrust review is required only if the sales or assets of the acquirer and target are above specified levels (e.g., $10 million for one party and $100 million for the other in 2001), a condition known as the size-of-person test. Our baseline findings are robust after excluding transactions with values between the size-of-transaction threshold and the higher cutoff that may be exempt from review under the size-of-person test.

authorities assess the transaction and determine whether further action is warranted. Outcomes generally fall into three main categories. First, if the waiting period expires without intervention, the transaction may proceed as planned. Second, firms may request early termination. If granted by both the DOJ and FTC, early termination shortens the review period and accelerates deal completion. Third, in transactions raising more substantial anti-competitive concerns, regulators may issue a "second request," requiring merging parties to provide extensive additional information and documentation. The issuance of a second request stops the initial clock, resets the waiting period upon resubmission, and typically extends regulatory scrutiny.

If regulators conclude that a proposed transaction may violate antitrust laws, they can formally challenge the deal. Such challenges typically result in three possible outcomes: (i) blocking the transaction entirely, (ii) imposing a consent decree that requires parties to take remedial action, or (iii) negotiating a settlement to mitigate anticompetitive risks.

In summary, the HSR Act significantly reduces the likelihood of post-merger antitrust challenges by front-loading regulatory oversight into the pre-completion phase. However, this added scrutiny introduces profound uncertainty for participating firms. Acquirers facing a second request, for instance, incur substantial legal and compliance costs and often experience significant delays in deal completion. Moreover, regulatory challenges often lead to breakup fees or other financial burdens that can distort strategic motives and threaten anticipated synergies. These considerations underscore why firms may seek strategies, such as IP licensing, to alleviate antitrust concerns during high-stakes M&A transactions.

2.2 IP licensing, M&As, and antitrust regulation

The strategic management of intellectual properties has become a central theme in corporate value creation, with its role extending well beyond the direct commercialization of a firm's own innovations.[2] While the literature has documented IP licensing as an important mechanism to monetize R&D, facilitate technology diffusion, and navigate market entry (e.g., Gallini (1984); Teece (1986)), it has taken on an increasingly important role in shaping the outcomes of major corporate transactions, particularly in mergers and acquisitions subject to antitrust scrutiny.

Antitrust authorities are primarily concerned with M&A deals that may substantially lessen competition. In IP-intensive sectors such as technology and pharmaceuticals, regulators often worry that a merger could allow the combined entity to leverage its control over critical intellectual assets (e.g., patents, software platforms, or proprietary content) to foreclose rivals, stifle innovation, or raise barriers to entry. These concerns are especially salient when the merged firm could restrict access to essential technologies or inputs that competitors require to remain viable.

To address these risks, antitrust authorities frequently require remedies as a condition for deal approval. These remedies vary in form. Some require the divestiture of assets or business units, while others rely on commitments that constrain how firms deploy or share key assets. While effective at reducing market share, firms often view divestiture as a costly solution due to the potential loss of synergies and the information asymmetry inherent in selling complex assets. Consequently, IP licensing has gained prominence as a flexible instrument for addressing regulators' competitive concerns, particularly in IP-intensive industries. By committing to license IPs to competitors, often on fair, reasonable, and non-discriminatory (FRAND) terms, an acquirer can credibly alleviate regulators' concerns that the merger will lead to foreclosure or exclusionary conduct. Such commitments signal that the transaction is motivated by efficiency gains rather than the accumulation of anticompetitive market power. In addition, compared to divesting critical

[2] In addition to internal exploitation, firms can monetize their intellectual properties through licensing or sale. While both facilitate the dissemination of innovation, licensing allows the firm to retain ownership of the underlying intellectual assets.

assets, IP licensing allows the acquirer to retain ownership of the valuable IP while directly addressing the specific competitive harm identified by regulators, thereby increasing the likelihood of regulatory approval.

A prominent contemporary example illustrating the strategic use of IP licensing in antitrust mitigation is Microsoft's $68.7 billion acquisition of Activision Blizzard. Announced in January 2022, the transaction immediately attracted intense antitrust scrutiny from regulators worldwide, including the FTC in the U.S., the Competition and Markets Authority (CMA) in the UK, and the European Commission. A central concern was that Microsoft could leverage its Xbox ecosystem by making Activision's flagship contents – most notably the *Call of Duty* franchise – exclusive to its own platform, thereby distorting competition in both the console and nascent cloud gaming markets. In response, Microsoft proactively addressed these concerns by offering legally binding IP licensing contracts to its key competitors, such as Nintendo and Sony, guaranteeing continued access to Activision's content across rival platforms. These licensing commitments directly targeted regulators' competitive concerns and played a critical role in securing regulatory approval. The Microsoft-Activision case illustrates how IP licensing has evolved from a simple monetization tool into a core strategical instrument in M&A transactions, enabling acquirers to navigate heightened antitrust enforcement while preserving deal synergies in large-scale, IP-centric mergers.

3. Data, Sample Construction, and Descriptive Statistics

3.1 IP licensing agreements

To analyze the strategic deployment of intellectual property, we construct a comprehensive sample of IP licensing agreements from two primary sources: (1) the RoyaltyStat database and (2) the SDC Joint Venture database. The construction of the licensing sample proceeds in two steps.

First, we extract licensing agreements from RoyaltyStat, one of the few companies that systematically collects information on IP licensing transactions. RoyaltyStat complies its data from material disclosures mandated by the U.S. Securities and Exchange Commission (SEC), whereby licensing agreements are often included as exhibits in standard scheduled filings or one-off filings. When these contracts contain commercially sensitive information, firms may request confidential treatment (CTR), which allows for the redaction of royalty rates and other proprietary terms. CTR protection typically lasts for 10 years from the filing date or until the contract expires. After CTR expiration, uncensored versions of these filings can be obtained through requests under the Freedom of Information Act (FOIA). Reflecting this regulatory process, RoyaltyStat organizes its licensing data into two distinct datasets.

The first dataset, the Royalty Rate Database, includes licensing agreements in which royalty rates are disclosed. These rates may become publicly available at the time of the filing if not redacted or subsequently obtained by RoyaltyStat through FOIA requests. In addition to royalty rates, this dataset provides detailed information on licensors and licensees, and effective dates. The second dataset, the Licensing Agreement Database, includes licensing agreements in which CTR protection has not yet expired, and royalty rates therefore remain unavailable. For these agreements, RoyaltyStat extracts information such as the identities of the licensor and licensee and the effective year of the agreement. To further refine RoyaltyStat's raw data, we leverage the recent advances in large language models (LLMs) and employ GPT-4o-mini to extract additional contract details, such as precise effective dates, from SEC EDGAR filings that include licensing exhibits. In both datasets, the underlying patents covered by the license are available for a subset of agreements, allowing for detailed valuation analysis.

Next, to capture licensing agreements that are voluntarily disclosed but not subject to SEC reporting requirements, we supplement RoyaltyStat with data from SDC Joint Venture database. In addition to SEC filings, SDC collects data on inter-firm cooperation agreements, including licensing deals, from various international counterparts, trade publications, news wires and other public sources. To identify relevant agreements and extract key details, such as licensors, licensees, and effective dates, we employ GPT-4o-mini to perform textual analysis of the deal synopses provided by SDC.

In our analyses, a licensing contract is between two entities: a licensor and a licensee. The filing firm must be publicly traded, typically the licensor or the licensee, while the counterparty can either be a commercial entity (e.g., a private or publicly listed firm) or a non-commercial entity (e.g., a non-profit hospital, university, or individual inventor). For each licensing agreement, we link the licensor to the Compustat-CRSP-Merged database using either the firm's Central Index Key (CIK) or a fuzzy name-matching algorithm. This procedure yields a sample of 23,740 licensing agreements spanning the period from 2001 to 2022, involving 3,530 unique public licensors. On average, licensing firms rank among the largest 10% of public firms by market value (with a median ranking in the largest 30%), highlighting their economic importance in technological innovation and competitive dynamics.

3.2 M&A sample, summary statistics, and variable construction

Our initial M&A sample consists of all completed and withdrawn U.S. mergers & acquisitions announced between February 1, 2001, and December 31, 2022, as reported in the SDC Mergers and Acquisition database. Consistent with the literature (e.g., Moeller, Schlingemann, and Stulz (2005)), we retain only transactions with deal values exceeding $1 million, as smaller transactions are unlikely to trigger antitrust scrutiny. We focus on transactions in which the acquirer is a public firm or a subsidiary of a public firm, [3] while the target can be a public firm, a subsidiary of a public firm, or a private firm. We further restrict our sample to transactions where the acquirer's pre-deal ownership stake in the target is below 50% and post-deal ownership is at least 50%. We exclude deals involving repurchases of the acquirer's own shares. To ensure the availability of firm-level financial data, we retain only those deals for which the acquirer has non-missing financial variables in the Compustat-CRSP-Merged database. [4] We exclude deals involving targets in certain sectors that are subject to industry-specific merger regulations and are unrelated to our analysis. Specifically, we exclude deals where the target is a financial firm (SIC codes 6000-6999) or a regulated utility (SIC codes 4900-4999). We also exclude acquisitions of hotels and motels (SIC code 7011), which are exempt from pre-merger review under FTC guidelines (FTC (2008)). Finally, we retain only transactions for which the U.S. Census Bureau's four-firm concentration ratio is available.

The sample selection procedure is summarized in Table 1, Panel A, and yields a final sample of 11,890 transactions. Among these, 5,890 are classified as reportable deals (i.e., transactions exceeding the applicable HSR pre-merger notification threshold at the time of announcement), and 6,000 are classified as non-reportable deals. Table 1, Panel B presents the summary statistics of key deal-level characteristics and reveals significant differences between reportable and non-reportable deals. Reportable deals are substantially larger, with an average transaction value of $1,352.77 million compared to $22.40 million for non-reportable deals. They also exhibit lower completion rates (96% versus 99%), consistent with greater regulatory scrutiny

[3] Under the Hart-Scott-Rodino (HSR) Act, the "acquiring entity" is defined based on the ultimate parent entity. Accordingly, when a subsidiary engages in a merger or acquisition, its ultimate parent company may also be subject to antitrust scrutiny, even if the parent firm is not directly involved in the transaction.
[4] We use the financial information of the acquirer's ultimate parent firm when the acquirer is a subsidiary of a publicly traded company.

faced by larger transactions. On average, reportable deals take 75 days to complete, compared to 26 days for non-reportable deals, highlighting the procedural delays associated with the pre-merger review. With respect to target attributes, reportable deals are less likely to involve private targets (38% versus 62%) but more likely to involve subsidiary targets (37% versus 33%). Reportable deals also display a modestly higher incidence of stock payment (12.36% versus 10.35%), reflecting differences in deal size and financing needs. Notably, 2% of reportable transactions are formally challenged by antitrust authorities, whereas non-reportable deals, by construction, do not face regulatory intervention.

4. Strategic IP Licensing in Merger & Acquisitions

4.1 Baseline model and results

In this section, we formally test whether acquiring firms strategically deploy IP licensing as a response to antitrust risk. As discussed in Section 2, M&A transactions that exceed the HSR pre-merger notification threshold are subject to mandatory review and potential enforcement actions, while those falling below it are generally exempt. We leverage this institutional feature to isolate the effect of antitrust exposure on firms' IP licensing behavior, distinguishing the specific pressure of regulatory oversight from the general economic determinants of M&A activity.

To estimate the relationship between antitrust risk and IP licensing, we employ a firm-year panel regression specified as follows:

$$License_{i,t+1} = \beta_0 + \beta_1 Reportable_{i,t} + \beta_2 NonReportable_{i,t} + X_{i,t}\gamma + \Psi_j + \theta_t + \epsilon_{i,t+1} \quad (1)$$

where the dependent variable, $License_{i,t+1}$ is an indicator equal to one if firm i enters into one or more IP licensing agreements in year t+1.[5] The key independent variable, $Reportable_{i,t}$, is an indicator equal to one if firm i announces at least one reportable M&A deal (i.e., exceeding the annual HSR threshold) in year t. Crucially, to differentiate regulatory motives from general acquisition-related synergies, we include $NonReportable_{i,t}$, an indicator for firm i announcing only deals that fall below the HSR threshold in year t. The omitted category consists of firm-years in which firm i does not announce any M&A transactions. To absorb time-invariant industry characteristics and aggregate macroeconomic shocks, we include industry fixed effects Ψ_j and year fixed effects θ_t. The vector of control variables, $X_{i,t}$, accounts for various time-varying firm characteristics that may be correlated with both M&A activity and IP licensing, including market-to-book ratio, leverage, return on assets (ROA), R&D ratio, and an indicator for high industry concentration. To account for the potential scaling effects of firm size on licensing capacity, we further control for quintic polynomials of firm size. All continuous variables are winsorized at the 1st and 99th percentiles to mitigate the influence of outliers. Summary statistics of these variables are presented in Table 1, Panel C. Standard errors are clustered at the firm-level to account for within-firm serial correlations among the residuals.

Table 3 presents the baseline regression results. In column (1), we find a positive and statistically significant relationship between reportable deal activity and subsequent licensing. The coefficient of $Reportable$ is 0.041, which is economically substantial. It implies that firms announcing reportable deals are 4.1 percentage points more likely to engage in IP licensing in the following year. Given the unconditional licensing probability of 8% in our sample, this

[5] We report OLS estimates because linear probability models yield more precise estimates of marginal effects when high-dimensional fixed effects are included (Angrist and Pischke, 2008). As a robustness check, we also estimate logit specifications and obtain qualitatively similar results.

corresponds to a 50% increase in licensing propensity relative to the baseline. In contrast, the coefficient of $NonReportable$ is small and insignificant. This null result provides a critical counterfactual: firms undertaking acquisitions that do not trigger mandatory antitrust review do not exhibit elevated licensing activity. This divergence suggests that the licensing is not a generic by-product of acquisition activity or the scale of a firm's operations, but is specific to the regulatory scrutiny faced by the acquirer. We assess the robustness of this result using tighter identification specifications in columns (2) and (3). Column (2) incorporates industry-by-year fixed effects to control for time-varying industry conditions, such as technological waves or consolidation trends, that might simultaneously drive M&A and licensing. Column (3) further includes firm fixed effects to account for unobservable, time-invariant firm characteristics such as persistent differences in business models or IP strategies. Across all specifications, the coefficient of $Reportable$ remains positive and statistically significant, while the coefficient of $NonReportable$ remains small and insignificant. Finally, column (4) restricts the sample to completed deals only and yields quantitatively similar results.

Taken together, these findings provide strong support for the acquirers' strategic deployment of IP licensing. The fact that licensing scales specifically with HSR-reportable status and is absent in non-reportable deals indicates that firms adjust the contractual form of their intellectual property holdings in response to the threat of antitrust scrutiny.

4.2 Alternative explanations and the timing of licensing

A central challenge to our identification strategy is distinguishing strategic regulatory responses from the mechanical correlates of firm scale and complexity. While the baseline result in Section 4.1 is consistent with the hypothesis that acquirers strategically deploy IP licensing to mitigate antitrust risk, it could also be driven by confounding factors unrelated to antitrust considerations. For example, reportable deals inherently involve larger firms with broader IP portfolios and more complex operations. If these firms are simply more prone to licensing irrespective of the transaction, our results might capture a scale effect rather than a regulatory response. Alternatively, firms with licensing-intensive business models may be more likely to pursue large acquisitions in the first place. In this section, we address these concerns using two complementary approaches.

4.2.1 Dynamic lead–lag analysis

First, following the approach of Barrot and Sauvagnat (2016) and Babina et al. (2024), we estimate a distributed lead-lag model using firm-level panel data. This dynamic framework allows us to trace the evolution of IP licensing activity around M&A announcements. This approach serves two critical purposes. First, it provides evidence against reverse causality by testing whether firms that undertake reportable acquisitions already exhibit differential licensing behavior prior to deal announcements. Second, it sheds light on the timing of the licensing response relative to the announcement of M&A transactions (i.e., the arrival of regulatory scrutiny).

We specify the dynamic lead-lag model as follows:

$$License_{i,t} = \alpha + \sum_{\tau=-2}^{2} \beta_{\tau} Reportable_{i,t-\tau} + \sum_{\tau=-2}^{2} \lambda_{\tau} NonReportable_{i,t-\tau}$$

$$+X_{i,t}\gamma + \psi_i + \theta_{j,t} + \varepsilon_{i,t} \quad (2)$$

where $Reportable_{t-\tau}$ ($NonReportable_{t-\tau}$) is a dummy variable equal to one if firm i announces at least one reportable (non-reportable) deal in year $t-\tau$. If a firm announces both reportable and non-reportable deals in the same year, $NonReportable$ is set to 0. We include firm fixed effects (ψ_i) to absorb time-invariant firm characteristics and industry-year fixed effects ($\theta_{j,t}$)

to control for time-varying industry-level shocks. Each lead-lag coefficient β_τ (λ_τ) captures the response of IP licensing activity in year t to reportable (non-reportable) M&A deals announced in year $t - \tau$.

Figure 2 reports estimated lead-lag coefficients. The results provide strong evidence against reverse causality or persistent firm characteristics. The solid dots show that the coefficients on the leads ($\tau < 0$) are statistically indistinguishable from zero, indicating that firms announcing reportable deals do not exhibit systematically differential licensing behavior prior to the announcement. This lack of pre-trends suggests that our baseline results are not driven by persistent firm characteristics or pre-existing licensing trajectories. In contrast, we observe a significant increase in IP licensing deployment following reportable deal announcements ($\tau > 0$). Furthermore, firms announcing only non-reportable deals (represented by the solid squares) show no comparable increase in licensing activity.

To further elucidate the timing of IP licensing relative to the M&A process, we complement the firm-year analysis with a deal-level approach that zooms in on the transaction timeline. Specifically, we estimate the following regression model:

$$License\ (a, b)_{i,d} = \alpha + \beta Reportable_{i,d} + X_{i,d,t}\gamma + \psi_i + \theta_{j,t} + \varepsilon_{i,d} \tag{3}$$

The dependent variable is a dummy equal to one if the acquirer i of deal d enters into one or more licensing agreements as a licensor within the specified event window (a, b) relative to the deal announcement (day *0*) or completion (day *end*), and zero otherwise. Again, the indicator, $Reportable_{i,d}$, captures whether the deal exceeds the HSR notification threshold. Figure 3 plots the estimated coefficients from across these alternative event windows. Consistent with the dynamic firm-level evidence in Figure 2, we find that the increase in IP licensing associated with reportable deals is concentrated precisely within the period of regulatory uncertainty – between deal announcement and completion. There is no differential licensing behavior between reportable and non-reportable deal acquirers prior to deal announcement (when antitrust risk has not yet materialized) nor after deal completion (when regulatory uncertainty is resolved). This temporal alignment reinforces the interpretation that licensing is deployed reactively to navigate the review process.

4.2.2 Exogenous variation from the 2001 HSR amendment

In our second approach, we exploit the 2001 amendment to the Hart-Scott-Rodino Act as a quasi-natural experiment. Effective February 1, 2001, the amendment raised the reporting threshold for M&A deals from $15 million to $50 million. This regulatory change effectively exempted a specific subset of transactions, those above $15 million but below $50 million, from mandatory pre-merger review, thereby removing the threat of antitrust scrutiny for these deals while leaving deal economics unchanged. We therefore conduct a difference-in-differences (DiD) analysis using firm-year observations from 1998 to 2003.

We defined the treated firms as those announcing deals with transaction values in the newly exempted range ($15 million to $50 million) and the control firms as those announcing deals that remain below the threshold (less than $15 million). We exclude years in which firms announce deals exceeding $50 million to ensure a clean comparison. The omitted category consists of firm-years in which firms do not announce any M&A transactions.

Table A1 in Internet Appendix presents the results. The key independent variable is the interaction term $Treat \times Post$, where $Post$ is an indicator for years 2001 and later. The negative and statistically significant coefficient on the interaction term suggests a decline in IP licensing following the threshold increase for firms whose deals became exempted from mandatory review. In contrast, firms announcing control deals exhibit no such change in licensing behavior. The evidence that the removal of regulatory oversight leads to a decline in licensing provides causal support for our mechanism: when the regulatory constraint is relaxed, the strategic incentive to license dissipates.

Taken together, the dynamic panel evidence, the granular deal-level timing analysis, and the DiD results collectively strengthen our causal interpretation. These findings suggest that the observed increase in IP licensing is not a generic feature of large-scale M&A or a proxy for firm complexity. Rather, it appears to be a targeted strategic tool deployed specifically when regulatory scrutiny is salient.

4.3 Cross-sectional tests and the role of antitrust risk

To further link the relationship between M&A activity and IP licensing to antitrust considerations, we explore cross-sectional variations in regulatory exposure and enforcement intensity. If IP licensing is indeed a strategic response to enforcement pressure, it should not be uniformly distributed across all reportable deals. Instead, it should be concentrated in transactions where the threat of intervention is most acute. We examine whether the baseline effect documented in Section 4.1 varies along three conceptually distinct dimensions of antitrust risk: review intensity, competitive overlap, and enforcement salience.

The first dimension captures the expected intensity of the regulatory review. While all reportable deals trigger a filing, the probability of facing a resource-intensive investigation varies significantly with deal size. As shown in Appendix Table A2, the likelihood that a deal receives a second request rises substantially for transactions above $1 billion. Motived by this nonlinearity, we partition reportable deals into two groups based on deal size. *Reportable large* indicates that a firm announces a reportable deal valued above $1 billion, while *Reportable small* captures reportable deals above the HSR threshold but below $1 billion. Table 3 column (1) presents the results. The coefficient of *Reportable large* is large and statistically significant, indicating that the increase in IP licensing is primarily driven by the largest reportable transactions. By contrast, the effect of *Reportable small* deals is significantly weaker. The likelihood of licensing increases by an additional 3.2 percentage points (=0.040-0.008) relative to smaller reportable transactions. This suggests that although these transactions may still be subject to review, the anticipated regulatory burden is lower. This gradient is consistent with antitrust review being both more salient and more resource-intensive for high-value transactions. Importantly, firms undertaking non-reportable deals show no significant increase in IP licensing, reinforcing the view that strategic licensing is tied to regulatory exposure rather than deal activity per se.

The second dimension captures the extent to which a transaction is likely to raise substantive competitive concerns. We first distinguish *Reportable horizontal* and *Reportable non-horizontal* deals, with analogous categories for non-reportable transactions. Horizontal deals , which involve overlapping business segments, are the canonical target of antitrust enforcement due to their potential to increase market power and unilateral effects. As shown in Table 3, column (2), we find that firms engaging in reportable horizontal deals exhibit 1.7 percentage points (=0.025-0.008) higher likelihood of subsequent IP licensing relative to those engaging in reportable non-horizontal transactions.[6] This pattern suggests that acquirers proactively address regulators' concern in transactions that are most likely to attract scrutiny. We complement this analysis by incorporating the industry market structure. Using the U.S. Census Bureau's top-four-firm concentration ratio, we classify industries as *high concentration* when the top-four market share exceeds the annual median. In column (3), we find that *Reportable high concentration* deals display a significantly larger post-deal licensing response by 2.2 percentage points (=0.024-0.002)

[6] One-sided test of *Reportable horizontal* and *Reportable non-horizontal* yields a p-value of 0.055, indicating significance at the 10% level.

higher than *Reportable low concentration* deals. Together, these results suggest that firms strategically deploy IP licensing where traditional competitive concerns are front and center (with competitive overlap and industry concentration), and are likely to trigger regulatory red flags.

The third dimension captures enforcement salience. Heightened regulatory salience and elevated scrutiny may alter firms' beliefs about the likelihood of intervention and the expected severity of remedies, thereby increasing the value of proactive licensing strategies. To test this channel, we partition reportable deals based on whether industry peers faced formal DOJ or FTC challenges in the year proceeding the focal deal, with analogous categories for non-reportable transactions. Column (4) of Table 3 shows that firms announcing reportable deals in industries with recent peer challenges (*Reportable peer challenge*) exhibits a significantly stronger licensing response with a 6 percentage points (=0.069-0.009) higher than *Reportable non-peer challenge* deals. This finding is consistent with visible scrutiny actions increasing perceived antitrust risk and encouraging precautionary behavior among rivals. We further proxy for enforcement salience using measures of external monitoring and public visibility. Specifically, we use media coverage and social media coverage as indicators of heightened scrutiny and reputational pressure.[7] We find that firms announcing reportable deals with above-median media coverage (*Reportable high media*) or social media attention (*Reportable high social media*) display stronger increases in IP licensing activity by 2.1, and 2.2 percentage points higher in licensing probability, respectively. These patterns suggest that public visibility amplifies regulatory pressure and strengthens firms' incentives to mitigate competitive concerns before they escalate.

[7] We measure a firm's media coverage as the number of news articles mentioning the firm in a given year, using the RavenPack database. A firm's social media coverage is measured as annual number of social media messages mentioning the firm (i.e., attention). We thank Cookson, Lu, Mullins, and Niessner (2024) for generously sharing the firm attention data. Because firm attention coverage begins in 2012, the regression sample in column (6) therefore contains fewer observations.

Overall, these cross-sectional results reinforce the interpretation that IP licensing is a strategic instrument calibrated to antitrust risk. Licensing activity is most pronounced precisely in setting where regulatory scrutiny is expected to be most intense, i.e., large deals, horizontal transactions, in concentrated industries, and environments characterized by heightened enforcement salience. These results complement the baseline findings by showing that firms' licensing responses are closely aligned with the underlying economic and institutional determinants of antitrust exposure.

4.4 The effect of licensing on M&A outcomes

Having established that firms systematically deploy IP licensing under the threat of antitrust enforcement, we next evaluate this strategy matters for regulatory outcomes. Does the strategic deployment of IP licensing ultimately benefit acquirers by facilitating deal completion and reducing regulatory frictions?

From a conceptual perspective, IP licensing presents a tradeoff for acquiring firms. On the one hand, licensing entails the dilution of monopoly rents and the sharing of valuable intellectual assets with potential rivals. On the other hand, it functions as a potent remedy: by committing to share technology, acquirers can credibly alleviate regulators' concerns about foreclosure or exclusionary conduct. If effective, the licensing strategy should lower the regulatory hurdle for approval and accelerate the timeline for complex transactions.

To test whether licensing facilitates regulatory approval and deal completion, we estimate the following deal-level regression model:

$$Outcome_{i,d} = \alpha + \beta_1 Challenged_{i,d} \times License(0, end)_{i,d} + \beta_2 License(0, end)_{i,d} + \beta_3 Challenge_{i,d} + X_{i,d}\gamma + \Psi_i + \theta_{j,t} + \epsilon_{i,d} \quad (4)$$

where $Outcome_{i,d}$ represents two alternative measures of deal success. First, $Completion$ is a binary indicator equal to one if the M&A transaction is completed and zero if it is withdrawn. Second, $Duration$ is the natural logarithm transformation of the number of days between deal announcement and completion. The key independent variable, $License(0, end)_{i,d}$ is an indicator equal to one if the acquirer enters into one or more IP licensing agreements during the transaction window (from announcement to completion or withdrawal).

To isolate the effect of licensing when it matters most, we construct an indicator, we construct an indicator, $Challenged_{i,d}$, equal to one if the transaction is formally challenged by the DOJ or FTC. We interact this variable with licensing activity, $License(0, end)_{i,d}$. The coefficient on the interaction term, $Challenged \times License(0, end)$, captures the differential effect of licensing specifically for transactions that face formal regulatory opposition. $X_{i,d}$ includes deal- and acquirer-level controls, such as payment method, target type, and firm characteristics described in Section 3.2. We also control for industry-year and firm fixed effects to absorb time-varying industry-specific shocks and unobservable acquirer heterogeneity.

Table 4 presents the results. We first examine the probability of deal completion in column (1). The estimated coefficient of the interaction term is positive and statistically significant, indicating that licensing materially increases the probability of completion for transactions facing regulatory headwinds. In terms of economic magnitude, the estimates imply that for challenged deals, the deployment of IP licensing increases the completion probability of approximately 9.7 percentage points (=0.126-0.029) relative to challenged deals without licensing. This suggests that licensing effectively rescues transactions that might otherwise be blocked or abandoned.

Column (2) of Table 4 examine deal duration, conditional on completion. That is, in this analysis, we exclude withdrawn deals. The interaction term enters with a negative and statistically significant coefficient of -1.818, suggesting that licensing dramatically accelerates the resolution of challenged transactions. The economic significance of this effect is substantial. Conditional on completion, challenged deals accompanied by licensing close roughly 42.6% (=($e^{-1.818 + 1.263} - 1) \times 100$) faster relative to challenged deals without licensing. In terms of time saved, this corresponds to a reduction of approximately 227 (=$75.07 \times e^{1.958}$) days relative to the average 532-day review period for challenged transactions. The fact that licensing has little or even negative effect in non-challenged transactions (reducing completion probability and increasing review duration) further suggests that its value arises from regulatory interaction rather than standalone transactional efficiency.

Taken together, these findings underscore the effectiveness of IP licensing as a tool for mitigating antitrust risks in high-stakes M&A transactions. By signaling a willingness to share market rents and reduce competitive harm, acquiring firms can materially alter regulatory outcomes, increasing completion probabilities and neutralizing the costly delays associated with antitrust intervention. The evidence reinforces the interpretation that licensing is not a generic feature of M&A activity, but a targeted mechanism deployed to navigate the most severe regulatory constraints.

5. Product Market Competition Following Mergers with IP Licensing

In this section, we examine the real economic consequences of acquisitions facilitated by IP licensing. A large literature documents that horizontal mergers can reduce product market competition, leading to higher prices and increased rents for merging firms and their rivals (e.g.,Kim and Singal (1993); Ashenfelter and Hosken (2010); Fathollahi, Harford, and Klasa (2022); Bonaimé and Wang (2024)). These effects typically arise when mergers allow firms to exercise greater market power by limiting output, raising prices, or softening competitive interactions.

However, mergers accompanied by IP licensing introduce a critical ambiguity due to competing interpretations of licensing. On the one hand, licensing may function as a pro-competitive mechanism: it reflects benign technology sharing or efficiency-enhancing collaboration, thereby mergers accompanied by licensing could exhibit muted or neutral price effects. On the other hand, if licensing primarily serves as a strategic tool to facilitate regulatory approval without fundamentally altering market structure, such mergers may still leave underlying competitive forces intact and enable acquirers greater market power

These competing views generate sharply different predictions for post-merger pricing, therefore leaving the product-market consequences of licensing-facilitated mergers an empirical question. In the remainder of this section, we investigate this question by examining product-level price changes following reportable mergers with and without IP licensing.

5.1 Effects on product market competition: Product prices

We first examine how acquisitions facilitated by IP licensing affect product prices of acquirers and their industry rivals for overlapping products. If IP licensing effectively preserves competitive conditions, we would expect post-merger price increases to be muted or neutralized. Conversely, if licensing is used strategically to circumvent regulatory scrutiny, we would expect product prices to rise following licensing-facilitated mergers.

We conduct this analysis using granular product-level pricing data from the NielsenIQ Retail Scanner Database, which provides comprehensive weekly information on product quantities sold and prices at the retail-store level from 2006 to 2023. Each product is uniquely identified by its Universal Product Code (UPC). Following Meier et al. (2025), we map products to their owning firms using a 2014 snapshot of UPC-to-company mappings.[8] We link product owners to acquirers and industry rivals through fuzzy name matching. To focus on the relevant competitive set, we focus on common products defined as those falling within the same product module shared by an acquirer and its industry rivals.

For each product, we construct weekly state-level prices by aggregating store-level observations.[9] Specifically, for a given product and state, the weekly price is calculated as total sales revenue across all stores in that state-week divided by the total number of units sold. We measure price change by comparing the average price over the two years following deal completion to the average price over the two years prior to the deal announcement.[10] The percentage point change in price is defined as the difference between these two averages normalized by the pre-announcement price, multiplied by 100. This procedure yields a sample of approximately 3.5 million product-state observations across 497 completed deals, 30,355 distinct products, and 51 states.

To estimate the effect of licensing on post-merger price changes, we use the following regression model:

$$\Delta\, Product\ Price_{i,d,p,s} = \alpha + \beta_1 Reportable_{i,d} \times License(0,end)_{i,d} + \beta_2 Reportable_{i,d} + \beta_3 License(0,end)_{i,d} + X_{i,d}\gamma + \Psi_p + \theta_t + \epsilon_{i,d,p,s} \tag{5}$$

[8] We obtain the 2014 snapshot from an open-source dataset originally hosted under the Product Open Data (POD) project, the predecessor to barcode.io (https://github.com/EventideSystems/brocade_sources).
[9] The product price-week panel is unbalanced for two reasons. First, if a store reports no sales of a product in a given week, the corresponding product price-week observation is missing from the Retail Scanner database. Second, for products offered for free through promotions, the database records the price as one cent and we therefore exclude such observations from our analysis.
[10] To account for the unbalanced nature of the product price data and to more accurately measure price changes associated with M&A transactions, we require products to have non-missing prices for at least 12 weeks in both the pre-announcement and post-completion periods.

where $\Delta\, Product\ Price_{i,d,p,s}$ is the percentage price change of common product p sold by acquirer i (or its rivals) in state s.[11] $Reportable_{i,d}$ is an indicator equal to one if the value of deal d exceeds the applicable HSR notification threshold, and $License(0, end)_{i,d}$ indicates if the acquirer licenses IP during the transaction window of deal d. The vector $X_{i,d}$ contains deal-level characteristics. We also include product fixed effects Ψ_p and year fixed effects θ_{t}.

Table 5 presents the regression results. In columns (1), the coefficient of the interaction term $Reportable \times License$ is positive and statistically significant. The estimate implies that reportable mergers accompanied by IP licensing are followed by a 1.1% increase in prices relative to comparable reportable deals without licensing. This magnitude is economically meaningful: given the sample's average post-merger price change of 1.89%, the estimated effect represents a 58% amplification in price growth. The result remains robust when we include state fixed effects in column (2) and product-by-state fixed effects in column (3), indicating that it is not driven by local demand shocks or time-invariant product-market differences.

To trace the dynamics of these price changes, we augment the long-difference specification with an event-time analysis. Following Liebersohn (2024), we adopt the regression-based approach advocated by Wooldridge (2023, 2025) to estimate a pooled difference-in-differences specification that allows treatment effects to vary by event time and cohort as follows:

$$Price_{d,p,s,t} = \alpha + \sum_{\tau=-8}^{8} \beta_\tau Reportable_d \times License_d \times \mathbf{1}\{t = \tau\} + X\gamma$$

$$+\psi_{cohort,t} + \theta_{p,s} + \varepsilon_{d,p,s,t} \qquad (6)$$

[11] Industry rivals are defined as firms operating in the same four-digit SIC industry as the acquirer.

Here, $Price_{d,p,s,t}$ is the natural logarithm of the average price of common product p sold by the acquirer i (or its rivals) in state s in quarter t. The event-time indicators $\mathbf{1}\{t = \tau\}$ span from eight quarters prior to the deal announcement to eight quarters after deal completion. The specification includes additional interaction terms and controls (X), cohort-by-quarter fixed effects ($\psi_{cohort,t}$), where cohorts are defined by deal announcement year, and product-state fixed effects ($\theta_{p,s}$).

Figure 4 plots the coefficient estimates of β_τ. We find no evidence of differential pre-trends: prices of common products are stable and statistically indistinguishable between reportable deals with and without licensing in the quarters leading up to the acquisition. In contrast, prices diverge significantly following deal completion. This post-completion spike suggests that licensing facilitates the consummation of mergers that subsequently allow the merged firm to exercise greater pricing power.

Taken together, the price evidence suggests that licensing does not fully neutralize competitive harm. Instead, it indicates that licensing functions as a strategic facilitator: it enables the approval of transactions that subsequently lead to higher product prices, effectively shifting the cost of regulatory compliance onto consumers.

5.2 Heterogeneous effect of IP licensing on product prices

To further examine the competitive mechanisms underlying licensing-facilitated mergers, we conduct a set of cross-sectional analyses. These tests are designed to identify the conditions under which IP licensing is mostly likely to affect product-market outcomes. Specifically, we assess whether the pricing effects documented in Section 5.1 vary systematically with antitrust concerns. If licensing serves as a genuine restorative remedy, it should dampen price increases in high-concern transactions. Conversely, if licensing is deployed strategically to clear regulatory hurdles without sacrificing core market power, we would expect price effects to be amplified where antitrust concerns are strongest.

First, we consider product-level overlap. We examine whether common products sold by the acquirer and its industry rivals also overlap with the target's product portfolio. In these overlapping products, the merger eliminates a direct competitor, creating the strongest incentive for unilateral price increases. If IP licensing primarily serves as a facilitating device that allows such mergers to clear regulatory review, rather than a mechanism that preserves competition, we would expect post-merger price increases to be concentrated in these overlapped products.

Second, we examine industry structure. Prior research shows that mergers occurring in more concentrated industries are more likely to generate significant changes in competitive behavior (e.g., Ashenfelter, Hosken, and Weinberg (2014); Grullon, Larkin, and Michaely (2019)). We therefore test whether the pricing power associated with licensing-facilitated deals is stronger in industries characterized by greater ex-ante market concentration.

Third, we consider the salience of antitrust enforcement. Elevated enforcement salience may alter firms' expectations regarding the likelihood of regulatory intervention and the severity of potential remedies, increasing the perceived value of IP licensing as a precautionary strategy. If licensing is deployed strategically in response to elevated enforcement risk, rather than reflecting benign technology diffusion, its pricing effect should be amplified when the acquirer anticipates tougher scrutiny.

To empirically test these hypotheses, we estimate cross-sectional regressions that extend Equation (5) by interacting $Reportable_{i,d} \times License(0, end)_{i,d}$ with variables capturing product overlap, industry concentration, recent peer challenges, or media scrutiny intensity. The results are presented in Table 6. We define *Overlapped product* as an indicator for common products that belong to the same product module as the target's products. *High industry concentration* equals one if the acquirer operates in an industry where the top-four firm market share exceeds the sample median. To capture enforcement pressure, *Peer challenge* indicates whether any merger involving the acquirer's industry peers was formally challenged by the DOJ or FTC in the year preceding the focal deal. Finally, *High media coverage* and *High social media coverage* proxy for public scrutiny, defined as indicators equal to one if the acquirer's media or social media attention is above the sample median.

Across all specifications in Table 6, the coefficients of the triple interaction terms are positive and statistically significant. This indicates that post-merger price increases are not uniform; rather, they are concentrated in precisely those settings where antitrust concerns are strongest. We find that price effects are significantly stronger for overlapped products, for deals in highly concentrated industries, and when enforcement salience is elevated due to recent peer challenges or heightened media scrutiny.

These findings suggest that licensing does not neutralize competitive harm in the most economically consequential settings. Instead, it facilitates the completion of mergers that allow firms to exercise greater pricing power where it matters most. Collectively, the heterogeneity results reinforce the interpretation that IP licensing operates as a strategic tool to navigate antitrust scrutiny rather than as a mechanism that broadly preserves competition. Licensing-facilitated mergers generate the largest price effects precisely when transactions are most likely to raise competitive concerns, underscoring the nuanced role of IP licensing in shaping post-merger outcomes.

6. More Nuances on IP Licensing

The granularity of our licensing data allows us to look beyond the extensive margin (the decision to license) and investigate the specific mechanisms through which acquiring firms strategically deploy IP licensing during M&A transactions. While the previous sections document that IP licensing increases in response to antitrust risk and facilitates the completion of reportable transactions, important questions remain regarding how firms structure these licensing agreements. Do these licenses represent a genuine transfer of competitive capability, or are they structured to minimize the erosion of post-merger rents?

In this section, we examine two key dimensions of licensing strategy: the economic value of the underlying assets and the contractual terms that govern their use. Together, these analyses shed light on the fine-grained strategies firms employ when navigating antitrust scrutiny.

6.1 Value of licensed patents

An acquiring firm faces a critical choice regarding which assets to share. The strategic view predicts selective concessions: firms license patents that satisfy review requirements while preserving economically important intellectual assets. In contrast, an efficiency-based interpretation would predict licensing of high-value technologies that meaningfully sustain competition. To distinguish between these mechanisms, in this section, we examine the economic value of licensed patents.

To test this intuition, we use licensing agreements in RoyaltyStat datasets that explicitly identify the underlying licensed patents and estimate the following regression model at the license-patent level:

$$PatValue_{i,l,p} = \alpha + \beta_1 Reportable_{i,l} + \beta_2 NonReportable_{i,l} + X_i \gamma + \Gamma_i + \theta_t + \epsilon_{i,l,p} \quad (6)$$

The dependent variable, $PatValue_{i,l,p}$, represents measures of patent quality and value. Following Kogan et al. (2017), we employ stock-market-based metrics to quantify the value and importance of patent.[12] For each licensed patent p associated with license agreement l, we distinguish whether the license is related to a reportable acquisition ($Reportable_{i,l}$), a non-reportable acquisition ($NonReportable_{i,l}$), or is unrelated to any M&A deal. The structure allows us to include firm fixed effects Γ_i to control for time-invariant firm characteristics that may be correlated with both licensing activity and patent value. We also include year fixed effects θ_t and a vector of licensor characteristics X_i.[13]

Table 7, Panel A presents the results. Columns (1) and (2) use the natural logarithm of KPSS nominal and real patent values (measured in millions of dollars), respectively. The coefficient of $Reportable$ is negative and statistically significant. The magnitude is economically substantial: patents licensed in reportable deals are approximately 40–42% lower in market value compared to patents licensed outside the context of M&A transactions.[14]

By contrast, licensing activity associated with non-reportable deals exhibit no systematic differences in patent valuation relative to licensing agreements unrelated to M&A activity. This divergence suggests that licensing decisions in smaller transactions are driven by standard business considerations, whereas licensing in reportable transactions is distorted by regulatory incentives. These findings support the strategic view: acquiring firms in reportable transactions appear to

[12] Unlike forward citations (Hall, Jaffe, and Trajtenberg (2001, 2005)), which measure scientific impact and suffer from severe truncation bias in recent years, the KPSS measure reflects the ex-ante private economic value of innovation embedded in market prices, making it more suitable for quantifying the financial value of licensed assets. In untabulated robustness tests, we repeat our analysis using citation counts and find that, while the coefficient of reportable is statistically insignificant, the interaction between reportable and political connections (Panel B) remains statistically significant and consistent with our findings using the KPSS measure.

[13] We employ year fixed effects rather than industry-by-year fixed effects due to limited variation in political connections within industry-years in the cross-sectional analysis presented in Table 7, Panel B.

[14] The coefficient of $Reportable$ in column (1) is -0.552, which corresponds to a 42.4% ($= e^{-0.552} - 1$) lower value. The coefficient of $Reportable$ in column (2) is -0.506, which corresponds to a 39.7% ($= e^{-0.506} - 1$) lower value.

selectively license economically marginal patents to comply with regulatory demands while preserving their most valuable intellectual assets.

6.2 Heterogeneity in IP licensing: Relationship between regulators and acquirers

We next explore the heterogeneity in acquirers' IP licensing strategies by examining how firms' political relationships with regulatory agencies shape their licensing behavior. Prior studies have documented that politically connected firms enjoy a range of advantages, such as greater access to government contracts (Tahoun (2014); Goldman et al. (2013)) and reduced regulatory enforcement (Correia (2014); Lambert (2019)). In the context of mergers and acquisitions, Fidrmuc, Roosenboom, and Zhang (2018) show that lobbying activity prior to deal announcements can mitigate antitrust scrutiny and is associated with more favorable regulatory outcomes. Building on these findings, we hypothesize that acquirers with stronger political connections are better positioned to license lower-value patents while retaining more valuable intellectual assets when navigating antitrust review.

To test this conjecture, we use two measures to capture a firm's political connections: lobbying expenditures and political contributions. Data on lobbying activity are obtained from the LobbyView database (Kim (2018)), which includes a comprehensive record of lobbying reports filed under the Lobbying Disclosure Act of 1995. Data on political action committee (PAC) contributions are obtained from Federal Election Commission, following prior studies such as Cooper, Gulen, and Ovtchinnikov (2010) and Hong and Kostovetsky (2012). Using these data, we estimate modified versions of Equation (6) by interacting political connection measures with indicators for licensing activity associated with reportable and non-reportable transactions.

Table 7, Panel B presents the results. Political connections further shape licensing strategies. In columns (1) and (2) (columns (3) and (4)), *Political-connected* is an indicator equal to one if the licensor reports positive lobbying expenditures (political contribution) in the three years prior to the licensing effective year. In three out the four specifications, the estimated coefficient of the interaction term *Reportable* × *Political-connected* is negative and statistically significant, suggesting that politically connected acquirers license patents of significantly lower economic value in reportable transactions relative to non-connected firms. The estimates imply that, conditional on being in a reportable deal, politically connected firms license patents that are an additional 29–52% lower in value relative to otherwise similar non-connected acquirers.[15] Combining the baseline and interaction effects, patents licensed by connected firms in reportable transactions are approximately 60–76% lower in market value than patents licensed outside the merger context.

These estimates suggest that political capital operates as a substitute for substantive intellectual property concessions. Politically connected firms appear able to leverage their regulatory access to satisfy antitrust concerns while relinquishing even less economically meaningful technology, retaining their highest-value assets while still securing approval. Overall, these findings highlight the interaction between political capital and corporate strategy in shaping how firms deploy IP licensing to mitigate antitrust risk.

6.3 License contract terms

In addition to selecting which patents to license, acquirers may strategically structure the terms agreements, such as royalty rates, duration, and exclusivity, to limit the competitive efficacy of the license. These dimensions provide additional margins through which acquirers can calibrate

[15] The coefficient of *Reportable* × *Political-connected* in column (1) is -0.762, which corresponds to a 51.6% ($= e^{-0.762} - 1$) lower value. The coefficient of the interaction term in column (3) is -0.338, which corresponds to a 28.7% ($= e^{-0.338} - 1$) lower value.

their concessions, signaling cooperation to regulators while preserving core competitive advantages.

To examine how licensing terms vary with antitrust exposure, we use detailed contract information from RoyaltyStat's Royalty Rate Database. This database provides granular data on licensing terms but is available only for a subset of agreements with disclosed royalty rates and contract features. As a result, our analysis in this section focuses on a sample of 351 licenses with non-missing royalty rates and other key contract terms.

We examine three primary contract features. First, *Royalty Rate* is measured as the percentage of revenue paid by the licensee to the licensor. Second, *License Duration* is measured as the number of years the licensing agreement remains in effect. Third, *Exclusivity* is a binary indicator equal to one if the license grants exclusive rights to the licensee and zero otherwise.

Table 8 presents the results. In column (1), we find that licensing agreements associated with reportable transactions feature significantly lower royalty rates than licenses unrelated to M&A activity. This evidence is consistent with acquirers' incentives to limit the flow of rents transferred through licensing. In column (2), the coefficient of *Reportable* is negative (though marginally insignificant) for licensing duration, suggesting that acquirers shorten the effective horizon of licensing agreements in reportable deals. Shorter durations restrict long-term competitive exposure while maintaining near-term regulatory compliance. In column (3), the coefficient of *Reportable* is negative but statistically insignificant for exclusivity, potentially implying less restrictive commitments.

Overall, the evidence in this section highlights the nuanced ways in which acquiring firms structure licensing contracts under regulatory pressure. Consistent with the evidence on patent value, firms under antitrust scrutiny appear to calibrate both the assets they license and the terms under which they license them to preserve core competitive advantages while satisfying regulatory demands. These findings underscore the flexibility firms possess in deploying IP licensing as a strategic tool to navigate antitrust scrutiny.

7. Conclusion

Motivated by heightened enforcement and increased scrutiny of large, IP-intensive M&A transactions, we examine whether acquiring firms strategically deploy IP licensing to navigate antitrust review and whether such licensing meaningfully preserves competition. Using comprehensive data that links IP licensing agreements to U.S. M&A transactions, we provide evidence that firms announcing transactions subject to mandatory pre-merger notification (i.e., reportable-deals) significantly increase IP licensing activity in the period following deal announcement. This response is not a generic feature of acquisition activity; it is absent for non-reportable deals and is concentrated precisely in settings where antitrust risk is most salient.

We demonstrate that this strategy is economically meaningful for M&A deal outcomes. Among transactions that face formal antitrust challenges, IP licensing increases the probability of deal completion and accelerates the resolution of regulatory review. However, licensing does not appear to fully neutralize the competitive consequences of reportable deals. Using granular product-level price data, we show that mergers accompanied by IP licensing are associated with higher post-merger product prices and these price effects are strongest precisely when market-power concerns are most acute. Consistent with a strategic mechanism, we further show that firms selectively license lower-value patents and structure licensing contracts with lower royalty rates and shorter durations to minimize the economic cost of compliance. Politically-connected acquirers are able to make even smaller IP concessions, highlighting the role of regulatory access in shaping corporate strategy.

Our findings contribute to the literature on mergers, innovation, and antitrust by identifying IP licensing as a flexible but imperfect regulatory accommodation. While licensing facilitates deal approval and reduces procedural frictions, it may also enable the consummation of mergers that ultimately weaken competition. These results raise the possibility that commitments to license IP may function primarily as remedies that satisfy formal regulatory requirements without fully constraining the merged firms' ability to exercise market power when firms retain discretion over *which* assets to license and *how* to structure the contracts. Overall, our findings suggest that licensing-based commitments warrant careful scrutiny, evaluated not only by their form but also on their scope and economic relevance. More broadly, our findings highlight the need for continued empirical investigations of how remedy design shapes real economic outcomes in increasingly innovation-driven markets.

Reference

Ashenfelter, Orley and Daniel Hosken, 2010, The Effect of Mergers on Consumer Prices: Evidence from Five Mergers on the Enforcement Margin, *The Journal of Law and Economics* 53, 417–466.

Ashenfelter, Orley, Daniel Hosken, and Matthew Weinberg, 2014, Did Robert Bork Understate the Competitive Impact of Mergers? Evidence from Consummated Mergers, *The Journal of Law and Economics* 57, S67–S100.

Babina, Tania, Anastassia Fedyk, Alex He, and James Hodson, 2024, Artificial Intelligence, Firm Growth, and Product Innovation, *Journal of Financial Economics* 151, 103745.

Barrot, Jean-Noël and Julien Sauvagnat, 2016, Input Specificity and the Propagation of Idiosyncratic Shocks in Production Networks, *The Quarterly Journal of Economics* 131, 1543–1592.

Bena, Jan, and Kai Li, 2014, Corporate Innovations and Mergers and Acquisitions, *The Journal of Finance* 69, 1923–1960.

Bonaimé, Alice and Ye (Emma) Wang, 2024, Mergers, Product Prices, and Innovation: Evidence from the Pharmaceutical Industry, *The Journal of Finance* 79, 2195–2236.

Cookson, J. Anthony, Runjing Lu, William Mullins, and Marina Niessner, 2024, The Social Signal, *Journal of Financial Economics* 158, 103870.

Correia, Maria M., 2014, Political Connections and SEC Enforcement, *Journal of Accounting and Economics* 57, 241–262.

Cunningham, Colleen, Florian Ederer, and Song Ma, 2021, Killer Acquisitions, *Journal of Political Economy* 129, 649–702.

De Loecker, Jan, Jan Eeckhout, and Gabriel Unger, 2020, The Rise of Market Power and the Macroeconomic Implications, *The Quarterly Journal of Economics* 135, 561-644.

Eckbo, B. Espen, 1985, Mergers and the Market Concentration Doctrine: Evidence from the Capital Market, *The Journal of Business* 58, 325–349.

Eckbo, B. Espen and Peggy Wier, 1985, Antimerger Policy under the Hart-Scott-Rodino Act: A Reexamination of the Market Power Hypothesis, *The Journal of Law and Economics* 28, 119–149.

Eckbo, B.Espen, 1983, Horizontal Mergers, Collusion, and Stockholder Wealth, *Journal of Financial Economics* 11, 241–273.

Farrell, Joseph and Carl Shapiro, 1990, Horizontal Mergers: An Equilibrium Analysis, *The American Economic Review* 80, 107–126.

Fathollahi, Maryam, Jarrad Harford, and Sandy Klasa, 2022, Anticompetitive Effects of Horizontal Acquisitions: The Impact of Within-Industry Product Similarity, *Journal of Financial Economics* 144, 645–669.

Fee, C. Edward and Shawn Thomas, 2004, Sources of Gains in Horizontal Mergers: Evidence from Customer, Supplier, and Rival Firms, *Journal of Financial Economics* 74, 423–460.

Fidrmuc, Jana P., Peter Roosenboom, and Eden Quxian Zhang, 2018, Antitrust Merger Review Costs and Acquirer Lobbying, *Journal of Corporate Finance* 51, 72–97.

FTC, 2008, To File or Not to File: When You Must File a Premerger Notification Form. https://www.ftc.gov/sites/default/files/attachments/premerger-introductory-guides/guide2.pdf.

Gallini, Nancy T., 1984, Deterrence by Market Sharing: A Strategic Incentive for Licensing, *The American Economic Review* 74, 931–941.

Goldman, Eitan, Jörg Rocholl, and Jongil So, 2013, Politically Connected Boards of Directors and The Allocation of Procurement Contracts, *Review of Finance* 17, 1617–1648.

Grullon, Gustavo, Yelena Larkin, and Roni Michaely, 2019, Are US Industries Becoming More Concentrated?, *Review of Finance* 23, 697–743.

Hall, Bronwyn H, Adam Jaffe, and Manuel Trajtenberg, 2001, Lessons, Insights and Methodological Tools, *NBER Working Paper* 8498, 4.

Hall, Bronwyn H., Adam Jaffe, and Manuel Trajtenberg, 2005, Market Value and Patent Citations, *RAND Journal of Economics* 36, 16–39.

Hong, Harrison and Leonard Kostovetsky, 2012, Red and Blue Investing: Values and Finance, *Journal of Financial Economics* 103, 1–19.

Kepler, John D., Charles G. McClure, and Christopher R. Stewart, 2025, Competition Enforcement and Accounting for Intangible Capital, *The Journal of Finance*.

Kepler, John D., Vic Naiker, and Christopher R. Stewart, 2023, Stealth Acquisitions and Product Market Competition, *The Journal of Finance* 78, 2837–2900.

Kim, E. Han, and Vijay Singal, 1993, Mergers and Market Power: Evidence from the Airline Industry, *The American Economic Review*, 549–569.

Kim, In Song, 2018, Lobbyview: Firm-Level Lobbying & Congressional Bills Database, *Unpublished manuscript, MIT, Cambridge, MA. http://web. mit. edu/insong/www/pdf/lobbyview. pdf Google Scholar Article Location*.

Kogan, Leonid, Dimitris Papanikolaou, Amit Seru, and Noah Stoffman, 2017, Technological Innovation, Resource Allocation, and Growth, *The Quarterly Journal of Economics* 132, 665–712.

Kwoka, John, 2014, *Mergers, Merger Control, and Remedies: A Retrospective Analysis of US Policy* (Mit Press).

Lambert, Thomas, 2019, Lobbying on Regulatory Enforcement Actions: Evidence from US Commercial and Savings Banks, *Management Science* 65, 2545–2572.

Liebersohn, Jack, 2024, How Does Competition Affect Retail Banking? Quasi-experimental Evidence from Bank Mergers, *Journal of Financial Economics* 154, 103797.

Mehta, Mihir N., Suraj Srinivasan, and Wanli Zhao, 2020, The Politics of M&A Antitrust, *Journal of Accounting Research* 58, 5–53.

Meier, Jean-Marie, Henri Servaes, Jiaying Wei, and Steven Chong Xiao, 2025, Do Consumers Care About ESG? Evidence from Barcode-Level Sales Data, *Working Paper*.

Moeller, Sara B., Frederik P. Schlingemann, and René M. Stulz, 2005, Wealth Destruction on a Massive Scale? A Study of Acquiring-Firm Returns in the Recent Merger Wave, *The Journal of Finance* 60, 757–782.

Cooper, Michael J., Huseyin Gulen, and Alexei V. Ovtchinnikov, 2010, Corporate Political Contributions and Stock Returns, *The Journal of Finance* 65, 687–724.

Phillips, Gordon M., and Alexei Zhdanov, 2013, R&D and the Incentives from Merger and Acquisition Activity, *The Review of Financial Studies* 26, 34–78.

Shahrur, Husayn, 2005, Industry Structure and Horizontal Takeovers: Analysis of Wealth Effects on Rivals, Suppliers, and Corporate Customers, *Journal of Financial Economics* 76, 61–98.

Stillman, Robert, 1983, Examining Antitrust Policy towards Horizontal Mergers, *Journal of Financial Economics* 11, 225–240.

Tahoun, Ahmed, 2014, The Role of Stock Ownership by US Members of Congress on the Market for Political Favors, *Journal of Financial Economics* 111, 86–110.

Teece, David J., 1986, Profiting from Technological Innovation: Implications for Integration, Collaboration, Licensing and Public Policy, *Research Policy* 15, 285–305.

Williamson, Oliver E., 1968, Economies as an Antitrust Defense: The Welfare Tradeoffs, *The American Economic Review* 58, 18–36.

Wollmann, Thomas G., 2019, Stealth Consolidation: Evidence from an Amendment to the Hart-Scott-Rodino Act, *American Economic Review: Insights* 1, 77–94.

Wooldridge, Jeffrey M, 2023, Simple Approaches to Nonlinear Difference-in-Differences with Panel Data, *The Econometrics Journal* 26, C31–C66.

Wooldridge, Jeffrey M., 2025, Two-Way Fixed Effects, the Two-Way Mundlak Regression, and Difference-in-Differences Estimators, *Empirical Economics* 69, 2545–2587.

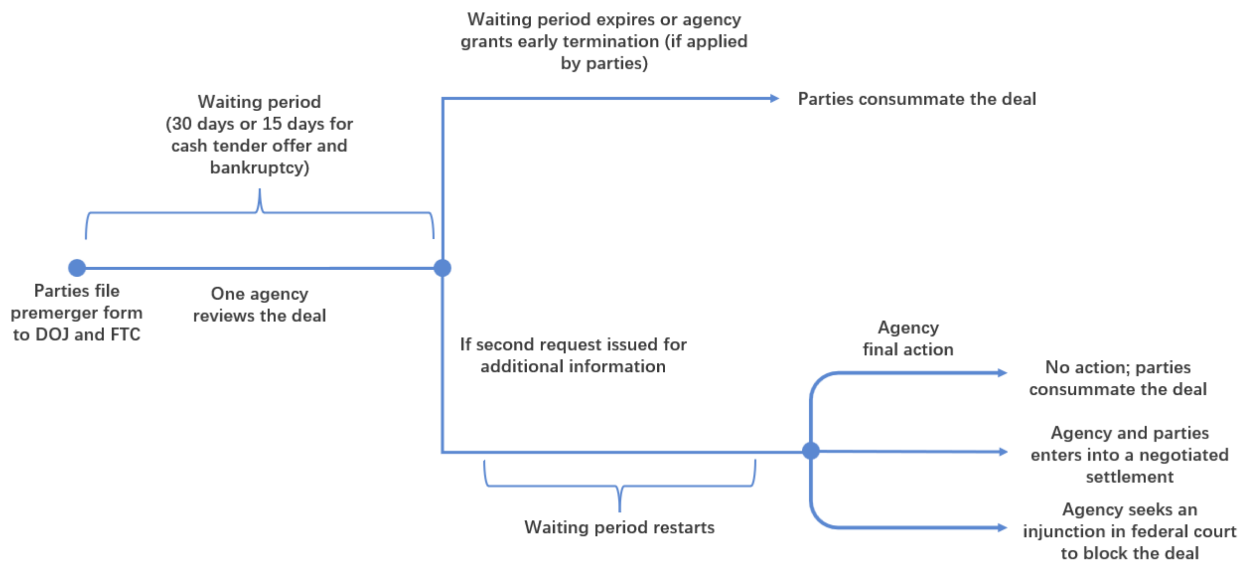

Figure 1. Timeline of HSR premerger notification process

This figure illustrates the statutory timeline of the HSR premerger notification process under the Hart-Scott-Rodino Act. It outlines the initial waiting period, the potential issuance of a Second Request, the extended waiting period, and the final regulatory actions that govern reportable transactions reviewed by the U.S. antitrust agencies.

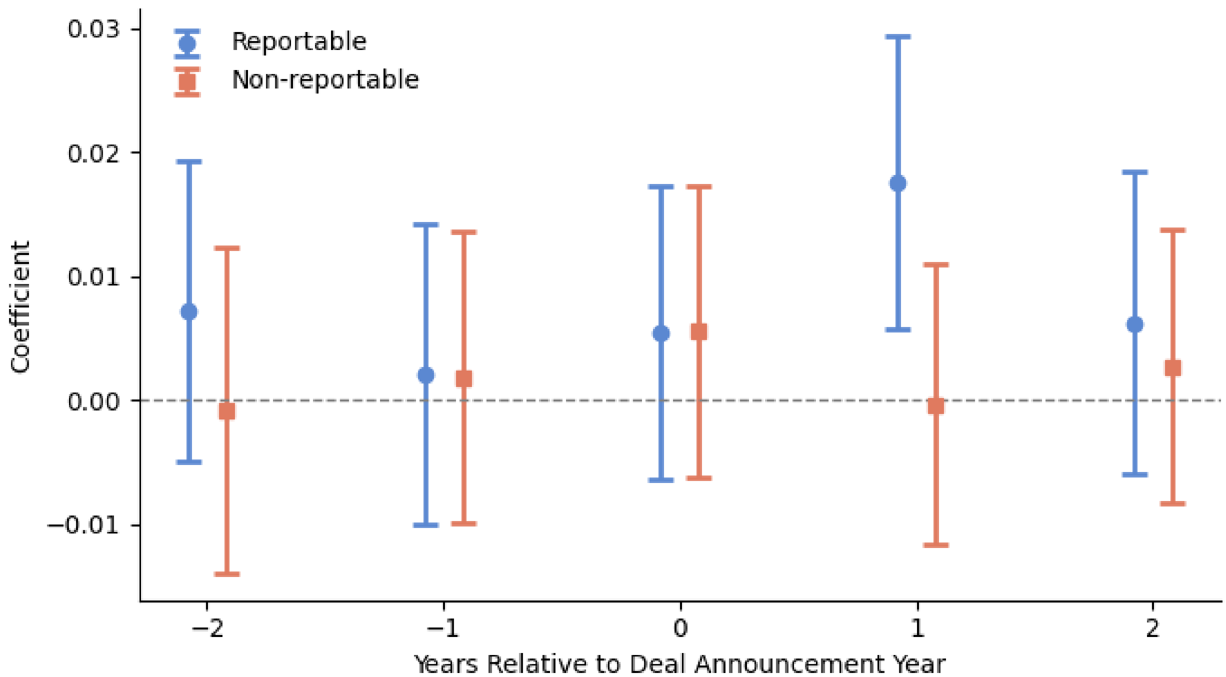

Figure 2. Dynamic relationship between M&A deals and IP licensing

This figure plots the estimated coefficients from a firm-year regression examining the dynamic relationship between M&A activity and IP licensing.

$$License_{i,t} = \alpha + \sum_{\tau=-2}^{2} \beta_\tau Reportable_{t-\tau} = + \sum_{\tau=-2}^{2} \lambda_\tau NonReportable_{t-\tau} + X_{i,t}\gamma + \psi_i + \theta_{j,t} + \varepsilon_{i,t}$$

The unit of observation is a firm-year. The dependent variable, $License_{i,t}$, is a dummy variable equal to one if a firm enters into one or more licensing agreements as a licensor in year *t*, and zero otherwise. $Reportable_{t-\tau}$ is a dummy variable equal to one if the firm announces at least one reportable M&A deal in year $t-\tau$. $NonReportable_{t-\tau}$ is a dummy variable equal to one if the firm announces at least one non-reportable deal in year $t-\tau$. If a firm announces both reportable and non-reportable deals in the same year, *Non-reportable* is set to zero. Control variables defined in Table 2 are included. The specification includes firm fixed effects (ψ_i) and industry-by-year fixed effects ($\theta_{j,t}$). Solid dots (squares) represent the estimated coefficients of *Reportable* (*Non-reportable*), and vertical line segments denote two-sided 95% confidence intervals based on standard errors clustered at the firm level.

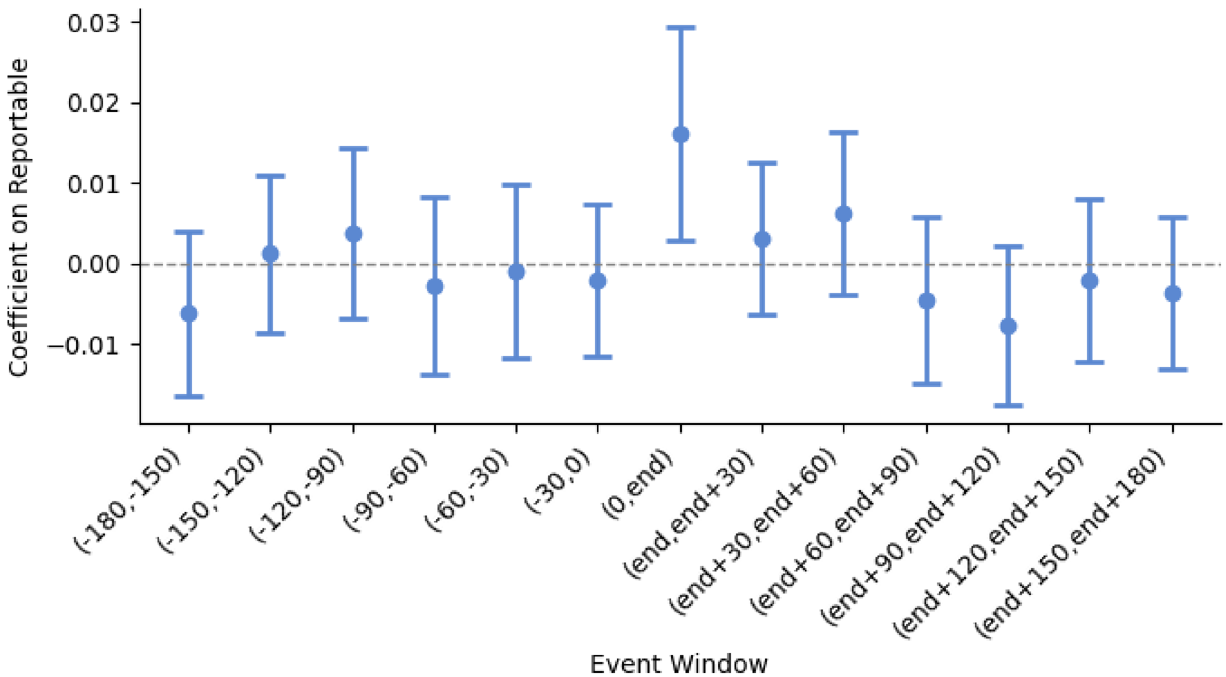

Figure 3. The timing of IP licensing: Deal-level analysis

This figure plots the estimated coefficients from a deal-level regression examining the effect of M&A antitrust risk on the likelihood and timing of IP licensing around M&A transactions.

$$License\ (a,b)_{i,d} = \alpha + \beta Reportable_{i,d} + X_{i,d,t}\gamma + \psi_i + \theta_{j,t} + \varepsilon_{i,d},$$

The unit of observation is a deal. The dependent variable is a dummy variable equal to one if acquirer *i* of deal *d* enters into one or more licensing agreement as a licensor within the specified event window (a,b) relative to the deal announcement or completion date, and zero otherwise. Event windows are shown on the horizontal axis, where day *0* denotes the deal announcement date, *end* denotes the deal completion or withdrawal date, and negative (positive) values denote days before announcement (after completion or withdrawal). *Reportable* is a dummy variable equal to one if the deal value is above the HSR notification threshold, and zero otherwise. Deal- and acquirer-level control variables, $X_{i,d,t}$, include a fifth-order polynomial in firm size (Size through $Size^5$), market-to-book ratio, leverage, ROA, R&D ratio, high industry concentration, private target, % of stock payment, tender, and hostile indicator, as defined in Appendix A. The specification includes firm fixed effects (ψ_i) and industry-by-year fixed effects ($\theta_{j,t}$). Solid dots represent the estimated coefficients of *Reportable* for each event window, and vertical line segments denote two-sided 95% confidence intervals based on standard errors clustered at the firm level.

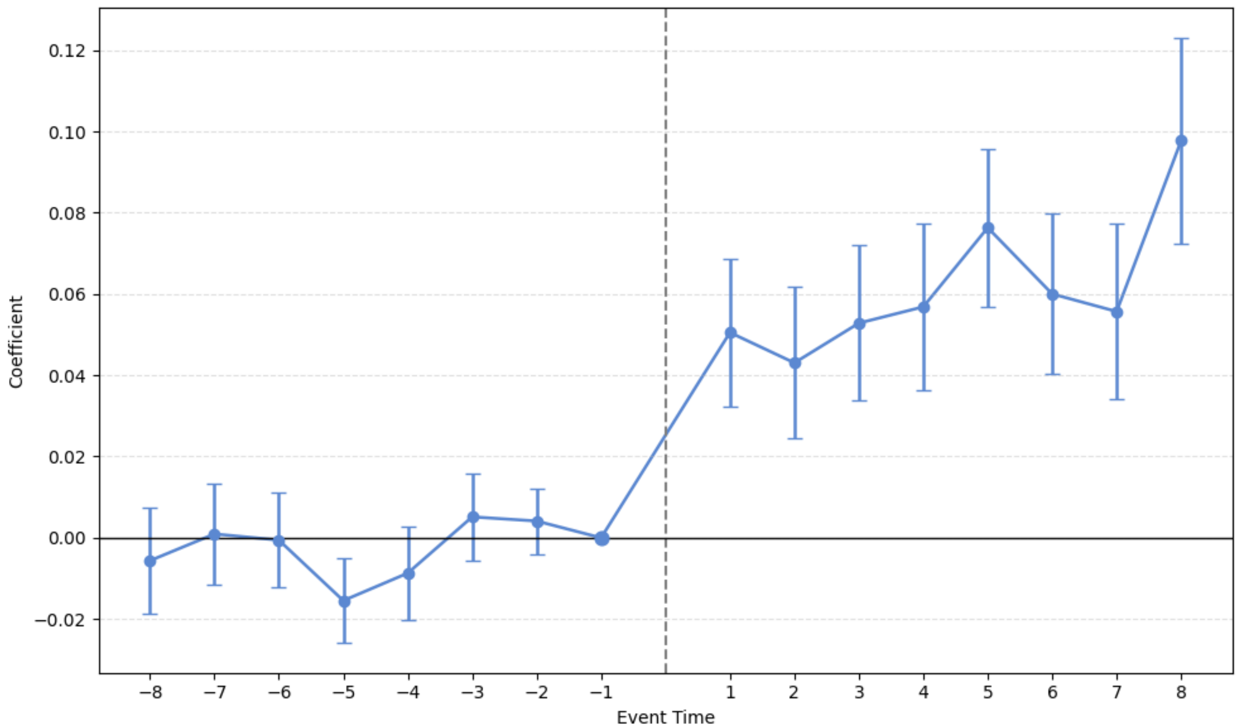

Figure 4. Product prices before and after M&A deals

This figure plots the estimated coefficients from a deal-product-state-quarter regression examining the dynamic of product prices before deal announcement and after deal completion.

$$Price_{d,p,s,t} = \alpha + \sum_{\tau=-8}^{8} \beta_\tau Reportable_d \times License_d \times \mathbf{1}\{t = \tau\} + \boldsymbol{X\gamma} + \psi_{cohort,t} + \theta_{p,s} + \varepsilon_{d,p,s,t}$$

The unit of observation is a deal-product-state-quarter. The dependent variable, *Price*, is the logarithm of the quarterly average product price in a given state. *Reportable* is a dummy variable equal to one if the deal value is above the HSR notification threshold. *License* is a dummy variable equal to one if the firm enters into one or more licensing agreement as a licensor in a given year. The event-time indicators $\mathbf{1}\{t = \tau\}$ measure quarters relative to the deal announcement (pre-period) and deal completion (post-period). $\boldsymbol{X}$ includes all interaction and main-effect terms as well as control variables for high industry concentration and private target status. Cohort-by-quarter fixed effects ($\psi_{cohort,t}$)are included, where cohorts are defined by deal announcement quarter, along with product-by-state fixed effects ($\theta_{p,s}$). Solid dots represent the estimated coefficients of the triple interaction terms, and vertical line segments denote two-sided 95% confidence intervals based on standard errors clustered at the product and state level.

Table 1: Sample selection of M&A deals and summary statistics

This table presents the sample selection procedure for the M&A sample (Panel A), summary statistics for deal characteristics (Panel B), and summary statistics for firm characteristics used in the baseline firm-year analysis (Panel C). Panel A describes the construction of the M&A sample using transactions from the Securities Data Company (SDC) Mergers and Acquisition database announced between February 1, 2001 and December 31, 2022. Panel B presents descriptive statistics for deal-level characteristics, separately for reportable and non-reportable transactions. Reportable deals are defined as transactions with deal values exceeding the applicable HSR premerger notification threshold at the time of announcement; non-reportable deals are defined as those below the threshold. Panel C presents summary statistics for key firm-level variables used in the baseline firm-year panel regressions. All variables are defined in Appendix A.

Panel A: Sample selection	
All U.S. completed and withdrawn deals (> $1 million) value from 2001 to 2022	58,088
Acquirer status: public/subsidiary; target status: public/subsidiary/private	35,768
Pre-deal ownership < 50%; post-deal ownership (intended in withdrawn deals) >= 50%	30,681
Deals not involving repurchase of acquirer's own or its subsidiaries' shares	30,074
Deals involving the acquirer in the CRSP-Compustat Merged database without missing financial variables	17,544
Deals not involving financial or utility targets, and not involving hotel & motel acquirers or targets	13,568
Deals with non-missing industry four-firm ratio from US Census Bureau	11,890

Panel B: Summary statistics of M&A deals

	Reportable deals				Non-reportable deals			
	N	Mean	Median	SD	N	Mean	Median	SD
Deal value ($ million)	5,890	1,352.77	268.00	5,620.24	6,000	22.40	17.00	18.47
Duration	5,890	75.07	48.00	100.26	6,000	26.42	0.00	65.38
Completion	5,890	0.96	1.00	0.21	6,000	0.99	1.00	0.12
Challenge	5,890	0.02	0.00	0.14	6,000	0.00	0.00	0.00
Private target	5,890	0.38	0.00	0.49	6,000	0.62	1.00	0.48
Subsidiary target	5,890	0.37	0.00	0.48	6,000	0.33	0.00	0.47
% of stock payment	5,890	12.36	0.00	28.22	6,000	10.35	0.00	26.25
Tender	5,890	0.06	0.00	0.23	6,000	0.01	0.00	0.09
Hostile	5,890	0.00	0.00	0.07	6,000	0.00	0.00	0.02

Panel C: Summary statistics of firm characteristics

	N	Mean	Median	SD
License	61,640	0.08	0.00	0.27
Reportable	61,640	0.07	0.00	0.25
NonReportable	61,640	0.06	0.00	0.23
Firm size	61,640	5.91	6.05	2.51
Market-to-book ratio	61,640	2.19	1.57	2.90
Leverage	61,640	0.23	0.18	0.29
ROA	61,640	0.04	0.10	0.29
R&D ratio	61,640	5.01	0.01	155.04
High industry concentration	61,640	0.51	1.00	0.50

Table 2: M&A antitrust risk and the likelihood of IP licensing

This table reports firm-year panel regressions examining the effect of M&A antitrust risk on the likelihood of IP licensing. The dependent variable, *License*, is a dummy variable equal to one if a firm enters into one or more licensing agreement as a licensor in a given year, and zero otherwise. *Reportable* (*NonReportable*) is a dummy variable equal to one if the firm announces at least one reportable (non-reportable) M&A deal in the prior year. If a firm announces both reportable and non-reportable deals in the same year, *NonReportable* is set to 0. Columns (1) to (3) include both completed and withdrawn deals, while column (4) restricts the sample to completed deals only. Control variables are defined in Appendix A. All columns include a fifth-order polynomial in firm size to accommodate potential non-linearity. Fixed effects vary by specification: column (1) include year and industry fixed effects; column (2) includes industry-by-year fixed effects; and columns (3) and (4) additionally include firm fixed effects. Standard errors are clustered at the firm-level. t-statistics are reported in parentheses. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively.

	Dependent variable: License			
Deal sample:		All deals		Completed deals
	(1)	(2)	(3)	(4)
Reportable	0.041***	0.045***	0.014***	0.012**
	(6.96)	(6.91)	(2.69)	(2.24)
NonReportable	0.008	0.008	0.000	0.000
	(1.63)	(1.59)	(0.08)	(0.09)
Market-to-book ratio	0.005***	0.005***	-0.000	-0.001
	(4.50)	(3.99)	(-0.31)	(-0.38)
Leverage	0.012	0.015*	0.024**	0.023*
	(1.52)	(1.72)	(2.03)	(1.96)
ROA	-0.125***	-0.130***	-0.075***	-0.074***
	(-10.51)	(-10.08)	(-4.53)	(-4.44)
R&D ratio	0.001	0.001*	0.002	0.002
	(1.52)	(1.69)	(1.48)	(1.47)
High industry concentration	0.005	-0.003	0.008	0.010
	(1.22)	(-0.52)	(1.11)	(1.36)
High-order firm size polynomials	Yes	Yes	Yes	Yes
Year FE	Yes	No	No	No
Industry FE	Yes	No	No	No
Industry-by-year FE	No	Yes	Yes	Yes
Firm FE	No	No	Yes	Yes
N	61,640	60,511	59,445	59,163
R^2	0.13	0.18	0.39	0.39

Table 3: Cross-sectional tests based on M&A antitrust risk

This table presents cross-sectional tests examining whether the relationship between M&A activity and IP licensing varies with proxies for antitrust exposure. Columns (1) to (3) capture regulatory jurisdiction and expected review intensity. In column (1), *Reportable large* is a dummy variable equal to one if the firm announces at least one reportable deal with transaction value at least $1 billion in the year; *Reportable small* is a dummy variable equal to one if the firm announces reportable deals below $1 billion; and *Non-reportable* equals to one if the firm announces only non-reportable deals. Column (2) distinguishes transactions by competitive overlap. *Reportable horizontal* is a dummy variable equal to one if the firm announces at least one reportable deal in which the acquirer and the target operate in the same four-digit-SIC industry; Reportable *non-horizontal* equals one if the firm announces at least one reportable deal and none are horizontal. *Non-reportable horizontal* equals one if the firm announces non-reportable deals only and at least one of them is horizontal; *Non-reportable non-horizontal* equals one if the firm announces non-reportable deals only and none are horizontal. Column (3) captures industry concentration using the U.S. Census Bureau's top-four-firm concentration ratio: *Reportable high concentration (low concentration)* indicates reportable deals in industries with concentration above (below) the annual median. *Non-reportable high concentration* and *Non-reportable low concentration* are defined using the same concentration cutoff for non-reportable deals. Columns (4) to (6) examine variations in enforcement salience. *Peer challenge* indicates whether an M&A deal involving an industry peer (defined at the four-digit SIC level) was formally challenged by the DOJ or FTC in the year preceding the focal deal. Media and social media coverage are measured as the number of RavenPack news articles and social media messages, respectively, mentioning the acquirer in 12 months prior to the deal announcement. *High (low) coverage* is defined relative to the annual median. All specifications include control variables listed in Table 2, as well as a fifth-order polynomial in firm size. Industry-by-year fixed effects and firm fixed effects are included throughout. Standard errors are clustered at the firm-level. t-statistics are reported in parentheses. F-statistics test the equality of coefficients across the corresponding reportable categories. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively.

Dependent variable: License			
	(1)	(2)	(3)
Reportable large	0.040***		
	(3.14)		
Reportable small	0.008		
	(1.49)		
NonReportable	0.000		
	(0.02)		
Reportable horizontal		0.025***	
		(2.80)	
Reportable non-horizontal		0.008	
		(1.20)	
Non-reportable horizontal		0.003	
		(0.28)	
Non-reportable non-horizontal		-0.001	
		(-0.15)	
Reportable high concentration			0.024***
			(3.22)
Reportable low concentration			0.002
			(0.28)
Non-reportable high concentration			0.010
			(1.21)
Non-reportable low concentration			-0.009
			(-1.40)
F-statistics	5.65**	2.57	4.49**
Controls	Yes	Yes	Yes
High-degree firm size polynomials	Yes	Yes	Yes
Industry-by-year FE	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes
N	59,445	59,445	59,445
R^2	0.39	0.39	0.39

(Continued on next page)

Dependent variable: License			
	(4)	(5)	(6)
Reportable peer challenge	0.069***		
	(3.09)		
Reportable non-peer challenge	0.009*		
	(1.74)		
Non-reportable peer challenge	0.004		
	(0.15)		
Non-reportable non-peer challenge	0.000		
	(0.01)		
Reportable high media		0.023***	
		(3.16)	
Reportable low media		0.002	
		(0.35)	
Non-reportable high media		0.012	
		(1.22)	
Non-reportable low media		-0.005	
		(-0.92)	
Reportable high social media			0.028***
			(2.73)
Reportable low social media			0.006
			(0.77)
Non-reportable high social media			0.015
			(1.07)
Non-reportable low social media			-0.000
			(-0.04)
F-statistics	6.88***	4.39**	3.07*
Controls	Yes	Yes	Yes
High-degree firm size polynomials	Yes	Yes	Yes
Industry-by-year FE	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes
N	59,445	59,445	28,130
R^2	0.39	0.39	0.43

Table 4: The effect of IP licensing on deal completion and duration

This table presents deal-level regressions examining the effect of IP licensing on M&A deal completion and deal duration. Columns (1) includes both completed and withdrawn transactions, with the dependent variable *Completion*, a dummy variable equal to one if the deal is completed and zero otherwise. Columns (2) restricts the sample to completed deals, and the dependent variable *Duration* is the natural logarithm of the number of calendar days between deal announcement and deal completion. *License(0, end)* is a dummy variable equal to one if the acquirer enters into one or more IP licensing agreement as a licensor between deal announcement date and deal completion or withdrawal date. *Challenged* indicates whether the M&A deal is formally challenged by the DOJ or FTC. The interaction term captures whether licensing differentially affects outcomes for challenged transactions. Control variables include deal- and acquirer-level characteristics: market-to-book ratio, leverage, ROA, R&D ratio, high industry concentration, private target, percentage of stock payment, tender offer, hostile indicator, as well as a fifth-order polynomial in firm size. All other variables are defined as in Table 2 and Appendix A. All specifications include industry-by-year fixed effects and firm fixed effects. Standard errors are clustered at the firm-level. t-statistics are reported in parentheses. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively.

Dependent variable:	Completion	Duration
	(1)	(2)
Challenged × License(0, end)	0.126**	-1.818***
	(1.97)	(-5.06)
License(0, end)	-0.029*	1.263***
	(-1.83)	(10.37)
Challenged	-0.078*	1.958***
	(-1.76)	(7.83)
Controls	Yes	Yes
High-order firm size polynomials	Yes	Yes
Industry-by-year FE	Yes	Yes
Firm FE	Yes	Yes
N	8,653	8,315
R^2	0.52	0.62

Table 5: Product-level price changes for M&As with IP licensing

This table presents regression results examining product price changes following M&A transactions accompanied by IP licensing. The unit of observation a deal–product–state level, and the sample includes completed deals only. The analysis focuses on common products sold by the acquirer and its industry peers, defined as products within the same product module. The dependent variable, *Δ Product Price*, is the percentage change a product's price within a state, defined as the difference between average price over the two years following deal completion and the average price over the two years prior to the deal announcement, divided by the pre-announcement price and multiplied by 100. *Reportable* is a dummy variable equal to one if the deal value is above the applicable HSR notification threshold, and zero otherwise. *License* is a dummy variable equal to one if the acquirer enters into one or more licensing agreements as a licensor between the deal announcement and completion dates. Fixed effects vary by specification: column (1) includes year and product fixed effects; column (2) adds state fixed effects; and column (3) includes product–state fixed effects. All other variables are defined as in Table 2 and Appendix A. Standard errors are clustered at the product and state level. t-statistics are reported in parentheses. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively.

	Dependent variable: Δ Product price		
	(1)	(2)	(3)
Reportable × License	1.091***	1.094***	1.109***
	(4.11)	(4.12)	(4.27)
Reportable	0.071	0.071	0.120
	(0.93)	(0.94)	(1.61)
License	-0.718***	-0.720***	-0.750***
	(-2.90)	(-2.91)	(-3.11)
High industry concentration	1.426***	1.435***	1.535***
	(6.68)	(6.74)	(7.35)
Private target	-0.147**	-0.148**	-0.164***
	(-2.46)	(-2.47)	(-2.78)
Year FE	Yes	Yes	Yes
Product FE	Yes	Yes	No
State FE	No	Yes	No
Product-state FE	No	No	Yes
N	3,512,086	3,512,086	3,209,041
R^2	0.34	0.34	0.49

Table 6: Cross-sectional tests on product-level price changes and IP licensing

This table presents cross-sectional tests examining heterogeneity in product price changes following IP-licensing-facilitated M&A transactions. The unit of observation is a deal–product–state, and the sample includes completed deals only. The dependent variable, *Δ Product Price*, is defined as in Table 5. In column (1), *Overlapped product* is a dummy variable equal to one if the common product sold by the acquirer and its peers also belong to a product module offered by the target, and zero otherwise. In column (2), *High industry concentration* is a dummy variable equal to one if the acquirer's industry concentration is above the sample median. In column (3), *Peer challenge* indicates whether an M&A deal involving an industry peer was formally challenged by the DOJ or FTC prior to the focal deal announcement. Columns (4) and (5) examine public visibility using indicators for high media coverage and high social media coverage of the acquirer, defined relative to the sample median. Each specification includes the corresponding triple interaction term between *Reportable*, *License*, and the relevant cross-sectional indicator, as well as all associated lower-order interaction and main-effect terms. Control variables in Table 5 are included. All specifications include year fixed effects and product-state fixed effects. Standard errors are clustered at the product-state level. t-statistics are reported in parentheses. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively.

	Dependent variable: Δ Product price				
	(1)	(2)	(3)	(4)	(5)
Reportable × License × Overlapped product	0.610*				
	(1.83)				
Reportable × License × High industry concentration		1.365**			
		(2.41)			
Reportable × License × Peer challenge			0.763***		
			(4.10)		
Reportable × License × High media coverage				1.511*	
				(1.98)	
Reportable × License × High social media coverage					1.208
					(1.33)
Other interaction and independent terms	Yes	Yes	Yes	Yes	Yes
Controls	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes
Product-state FE	Yes	Yes	Yes	Yes	Yes
N	3,209,041	3,209,041	3,209,041	3,209,041	2,006,987
R^2	0.49	0.49	0.49	0.49	0.49

Table 7: Value of M&A-related licensed patents

This table examines how the value of licensed patents varies with M&A activity. The unit of observation sample is a license-patent. In Panel (A), the dependent variable in column (1), *Nominal value*, is the natural logarithm of KPSS patent value measured in millions of nominal dollars, and in column (2), the dependent variable, *Real value*, is the natural logarithm of KPSS patent value deflated to 1982 (million) dollars using the CPI. Both nominal and real patent values are obtained from Kogan et al. (2017). *Reportable* is a dummy equal to one if the license agreement becomes effective during the licensor's reportable transaction window, and zero otherwise. *NonReportable* indicates whether the license agreement is effective during the licensor's non-reportable deal window. Panel B examines the heterogeneity in licensed patent value by acquirers' political connections. In columns (1) and (2), political connections are measured by lobbying activities, where. *Political-connected* equals one if the acquirer reports positive lobbying expenditures in the three years prior to the deal announcement, and zero otherwise. In columns (3) and (4), political connections are measured by whether the acquirer has positive political action committee contributions in the three years prior to the deal announcement. All regressions include licensor-level controls: a fifth-order polynomial in firm size, market-to-book ratio, leverage, ROA, R&D ratio, and high industry concentration. All specifications include year and firm fixed effects. Standard errors are clustered at the licensor firm level. t-statistics are reported in parentheses. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively.

Panel A: The value of M&A-related licensed patents		
Dependent Variables:	Nominal value	Real value
	(1)	(2)
Reportable	-0.552***	-0.506***
	(-3.72)	(-3.47)
NonReportable	-0.072	-0.080
	(-0.48)	(-0.60)
Controls	Yes	Yes
High-order firm size polynomials	Yes	Yes
Year FE	Yes	Yes
Firm FE	Yes	Yes
N	2,963	2,963
R^2	0.72	0.72

Panel B: Political connection and the value of M&A-related licensed patents				
Dependent variable:	Nominal value	Real value	Nominal value	Real value
Political proxy:	Lobbying		Campaign contribution	
	(1)	(2)	(3)	(4)
Reportable × Politically-connected	-0.726**	-0.653**	-0.338	-0.383**
	(-2.37)	(-2.24)	(-1.62)	(-2.00)
Reportable	-0.325*	-0.302*	-0.375***	-0.301***
	(-1.96)	(-1.81)	(-4.81)	(-3.97)
NonReportable × Politically-connected	0.127	0.018	0.174	0.058
	(0.56)	(0.09)	(0.77)	(0.29)
NonReportable	-0.077	-0.076	-0.070	-0.067
	(-0.46)	(-0.52)	(-0.41)	(-0.45)
Political-connected	0.030	0.028	-0.728***	-0.716***
	(0.20)	(0.18)	(-3.58)	(-3.22)
Controls	Yes	Yes	Yes	Yes
High-order firm size polynomials	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes	Yes
N	2,963	2,963	2,963	2,963
R^2	0.72	0.72	0.72	0.72

Table 8: Contract terms of M&A-related licensing

This table examines how licensing contract terms vary with M&A activity. The unit of observation is a licensing agreement. In column (1), the dependent variable, *Royalty Rate*, is the natural logarithm of the royalty rate specified in the licensing agreement. In column (2), *License Duration* is the natural logarithm of the number of years the licensing agreement remains in effect. In column (3), *Exclusivity* is a dummy equal to one if the license grants exclusive rights to the licensee. *Reportable* (*NonReportable*) is a dummy equal to one if the license agreement becomes effective during the licensor's reportable (non-reportable) M&A transaction window, and zero otherwise. All regressions include licensor-level controls: a fifth-order polynomial in firm size, market-to-book ratio, leverage, ROA, R&D ratio, and high industry concentration. All other variables are defined in Table 2 and Appendix A. All specifications include industry-by-year and firm fixed effects. Standard errors are clustered at the licensor firm level. t-statistics are reported in parentheses. *, **, *** denote significance at the 10%, 5%, and 1% levels, respectively.

Dependent variable:	Royalty rate	License duration	Exclusivity
	(1)	(2)	(3)
Reportable	-0.632***	-0.226	-0.115
	(-3.43)	(-1.33)	(-0.55)
NonReportable	-0.094	0.072	-0.124
	(-0.23)	(0.66)	(-0.43)
Controls	Yes	Yes	Yes
High-order firm size polynomials	Yes	Yes	Yes
Industry-by-year FE	Yes	Yes	Yes
Firm FE	Yes	Yes	Yes
N	351	351	351
R^2	0.59	0.65	0.68

Appendix

Appendix A: Variable Definitions

Variable	Description	Source
Dependent variables		
License	Dummy variable equal to one if a firm enters into one or more licensing agreements as a licensor in a given year, and zero otherwise (firm-year panel regressions).	RoyaltyStat/SDC
$License\ (a, b)$	Dummy variable equal to one if the acquirer enters into one or more licensing agreements as a licensor within the specified event window (a, b) relative to the deal announcement date (deal-level regressions).	RoyaltyStat/SDC
Completion	Dummy variable equal to one if the M&A transaction is completed, and zero if it is withdrawn.	SDC
Duration	The natural logarithm of the number of calendar days between deal announcement and deal completion.	SDC
Price	The logarithm of the quarterly average product price in a given state.	NielsenIQ Retail Scanner Database
Δ Product Price	Percentage change in a product's price within a state, defined as the difference between the average price over the two years following deal completion and the average price over the two years prior to deal announcement, divided by the pre-announcement average price and multiplied by 100. Weekly state average prices are computed as total sales revenue divided by total units sold across all stores.	NielsenIQ Retail Scanner Database
Nominal value	Natural logarithm of the KPSS patent value measured in millions of nominal U.S. dollars.	Kogan et al. (2017)
Real value	Natural logarithm of the KPSS patent value deflated to 1982 (million) U.S. dollars using the CPI.	Kogan et al. (2017)
Royalty rate	Natural logarithm of the royalty rate specified in the license agreement.	RoyaltyStat
License duration	Natural logarithm of the number of years the license agreement remains in effect.	RoyaltyStat
Exclusivity	Dummy variable equal to one if the license grants exclusive rights to the licensee, and zero otherwise.	RoyaltyStat
Independent variables		
Reportable	Dummy variable equal to one if the M&A transaction value exceeds the applicable HSR premerger notification threshold at the time of announcement, and zero otherwise.	SDC
NonReportable	Dummy variable equal to one if the M&A transaction value is below the applicable HSR notification threshold, and zero otherwise.	SDC
Horizontal	Dummy variable equal to one if the acquirer and target operate in the same four-digit SIC industry.	SDC
Challenged	Dummy variable equal to one if the M&A transaction is formally challenged by the DOJ or FTC, and zero otherwise.	DOJ/FTC
Political-connected	Dummy variable equal to one if the acquirer has positive lobbying expenditures (or positive political action committee contributions) in the three years prior to the deal announcement, and zero otherwise.	Lobbyview/FEC

Control variables		
Firm size	The natural logarithm of firm sales.	Compustat-CRSP-Merged
Market-to-book ratio	Market value of total assets divided by book value of total assets.	Compustat-CRSP-Merged
Leverage	Book value of debt divided by book value of total assets.	Compustat-CRSP-Merged
ROA	Operating income before depreciation divided by average total assets.	Compustat-CRSP-Merged
R&D ratio	Research and development expenses divided by sales.	Compustat-CRSP-Merged
High industry concentration	Dummy variable equal to one if the acquirer operates in an industry where the top-four-firm concentration ratio is above the annual sample median, and zero otherwise.	U.S. Census Bureau
Private target	Dummy variable equal toi one if the target is a private firm, and zero otherwise.	SDC
Percent of stock payment	Percentage of total deal value financed with stock.	SDC
Tender	Dummy variable equal to one if the M&A transaction is structured as a tender offer, and zero otherwise.	SDC
Hostile	Dummy variable equal to one if the M&A transaction deal is classified as hostile, zero otherwise.	SDC

Appendix B: Additional Tables

Appendix Table A1: Additional test on M&A antitrust risk and IP licensing using 2001 HSR Reform

This table presents an additional test of the effect of M&A antitrust risk on the likelihood of IP licensing, exploiting the 2001 HSR reform that raised the premerger notification threshold from \$15 million to \$50 million.

$$License_{i,t+1} = \beta_0 + \beta_1 Treat_i \times Post_t + \beta_2 Treat_i + \beta_3 Control_i \times Post_t + \beta_4 Control_i + X_{i,t}\gamma + \psi_i + \theta_t + \varepsilon_{i,t+1}$$

The unit of observation is a firm-year, and the sample spans from 1998 to 2003. We exclude firm-year observations in which a firm announces a deal with transaction value greater than or equal to \$50 million. The dependent variable, *License*, is a dummy variable equal to one if a firm enters into one or more licensing agreements in a given year, and zero otherwise. *Treat* is a dummy variable equal to one if a firm announces an M&A deal with transaction value between \$15 million and \$50 million in a given year, and zero otherwise. *Control* is a dummy variable equal to one if a firm announces a deal with transaction value below \$15 million. *Post* indicates years 2001 and later. All other variables are defined in Table 2 and Appendix A. Standard errors are clustered at the firm level. t-statistics are reported in parentheses. *, **, *** indicate significance at the 10%, 5%, and 1% levels, respectively.

Dependent variable: License	
	(1)
Treat × Post	-0.047*
	(-1.84)
Treat	0.043***
	(2.65)
Control × Post	-0.033
	(-1.51)
Control	0.027*
	(1.74)
Controls	Yes
High-order firm size polynomials	Yes
Industry-by-year FE	Yes
Firm FE	Yes
N	20,967
R^2	0.48

Appendix Table A2: Percentage of HSR-reportable M&A transactions receiving second-request investigations by deal size

Range ($ million):	50-100	100-150	150-200	200-300	300-500	500-1,000	1,000-10,000 (over 1,000)	over 10,000	Total
2010	2.8	2.9	0	5.6	4.8	3.2	10.5		4.1
2011	0.9	0.8	3	2.5	5.7	4.3	15.1		4.1
2012	1.7	1	2.3	3.8	2.5	4.8	11.5		3.5
2013	1.9	1.1	3.3	2.3	2.4	4.8	12		3.7
2014	1.8	0.7	0.5	1.6	2.9	3.3	11.1		3.2
2015	0	1.3	0.9	0.9	1.2	1.5	12		2.7
2016	3.5	0.6	0.4	1.4	2.8	3.2	10		3
2017	0	1.2	0.7	2	1.2	2.3	10.2		2.6
2018	0	0.3	0.7	1.3	1.8	1.5	9.5		2.2
2019	0	1.5	0.7	0.4	2	4	10.3		3
2020	0	1.2	1.9	2.6	2.4	3.8	6.4		3
2021	0	1.2	0.6	1.1	1.7	1.8	4.7		1.9
2022	0	0.7	0.2	0.4	1.4	1.4	4.3		1.6
2023	0	0	0.4	2.7	1.2	2.5	3.8		2.1
2024	0	0.8	1	0.8	1.3	2	6.4	47.1	3