Mass Exit Under Governance Conflict

A Behavioral Foundation for Theories of State Development[*]

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ABSTRACT. Many influential theories of state development assume that political fragmentation enables subjects to exit across competing rulers, thereby disciplining extraction, yet direct evidence that fragmentation shapes how subjects resist remains scarce. I show this in Tokugawa Japan (1615–1868), where nearly 300 feudal domains operated under the central Tokugawa shogunate, generating sharp variation in villages' exposure to governance conflict. In a difference-in-differences design, the shogunate's conversion of domains to direct rule raised the probability of nearby mass exit by 4 percentage points with no rise in direct confrontation, and a grid-year panel confirms the pattern. Among revolting villages, greater exposure to governance conflict tilted resistance toward exit over confrontation. Domains more exposed to governance conflict levied lower taxes, especially where neighboring domains could absorb migrants. The findings microfound a premise that load-bears across the Great Divergence literature, fiscal-federalism arguments, and the cartel theory of the territorial state system.

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1. Introduction

Many influential theories of state development rest on a common premise: political fragmentation enhances the mobility of subjects across competing rulers, and that mobility disciplines extraction and shapes institutions. This premise carries the central causal weight of leading explanations for the Great Divergence (Cox 2017; Fernández-Villaverde et al. 2023; Jones 2003; Mokyr 2017); for accounts linking Europe's polycentric structure to constrained taxation (Abramson 2017; Dincecco and Wang 2018; Sng and Moriguchi 2014; Stasavage 2016); for fiscal-federalism arguments that competing jurisdictions tame Leviathan (Qian and Weingast 1997; Rodden 2003; Tiebout 1956; Weingast 1995); and for recent theories positing that the modern territorial state system itself arose as a cartel against subject mobility (Acharya and Lee 2018; 2023). In each case, the mechanism is the same: fragmentation offers subjects credible exit options, and the threat or use of exit forces rulers to compete or cooperate. Yet the behavioral foundation of this chain has yet to be directly demonstrated: that the intensity of governance conflict actually shaped how subjects resisted, or put differently, that subjects rationally exploited fragmented polity to check oppression. I show just this.

Direct evidence supporting this premise is scarce for two main reasons. First, data on exit are limited (Bailey 2014, 76): unlike open rebellion, which leaves abundant documentary traces, exit is often subtle and covert. As a result, most existing research infers exit indirectly rather than observing it directly at the community level. Second, it is difficult to find usable variation in governance conflict. Macro-level studies that adopt this premise typically compare entire regions, such as Europe and China, whose polities differ across many dimensions, providing little within-case variation suitable for testing a behavioral mechanism at the level where subjects actually make choices. This paper addresses both challenges.

Tokugawa Japan (1615–1868) provides a clean laboratory for this mechanism. Nearly 300 feudal domains formed a patchwork of autonomous units, each with its own legal code, tax system, and border controls, held together by the Tokugawa shogunate as hegemon, a hybrid regime of domain autonomy combined with supra-domain arbitration, attainder power, and direct peasant-appeal authority retained at the center (Moore 1993; Ravina

1999; White 1995). Three features make the setting exceptional. First, the post-1615 peace shuts off the war-finance channel that Rosenthal and Wong (2011) identify as overwhelming the exit-discipline mechanism in early modern Europe, isolating the mechanism itself (Moore 1993). Second, intense inter-domain competition for agricultural labor made mass migration a credible peasant strategy, and villages varied widely in their geographic exposure to rival rulers. Third, newly digitized records of peasant uprisings compiled by Aoki (1975) document both direct confrontations and mass exits at the village level across the entire period.[3]

The empirical analysis proceeds in three parts. First, exposure to governance conflict raised the *occurrence* of exit. A difference-in-differences design exploits an exogenous shock to local outside options: the shogunate's periodic abolition of domains and conversion of their territories into direct shogunal rule (*tenryo*), widely perceived as less heavily taxed and less coercively governed than daimyō domains. Conversion also installed a shogunal intendant at the former domain capital, sharply lowering the cost of direct appeal to the shogunate for nearby villagers. Tenryo conversion raised the probability of nearby exit by 3.9 percentage points while rates of direct confrontation remained unchanged, and event studies show no pre-trends and a post-conversion spike in exit. A grid-year panel tracking within-cell shifts in the relative distance to the two nearest active capitals as nearby domains are established, abolished, or relocated corroborates this result: greater exposure to governance conflict predicted more frequent exit events across the full panel of cells, with no corresponding rise in direct confrontation.

Second, among villages that did revolt, greater geographic exposure tilted the chosen form of resistance toward exit over direct confrontation. I measure each village's exposure as the ratio of its distance to home-domain capital to its distance to the nearest outside administrative center. Across all recorded revolts, this exit-*share* result holds robustly

[3] Tokugawa mass exit operated as active resistance rather than permanent relocation: peasants presented demands to home-domain officials while seeking permission from neighboring rulers to cross the border, frequently returning once concessions were granted. Moore (1993, p. 255), drawing on Borton's catalog, recorded 106 such mass desertions across the Tokugawa period as evidence of "the solidarity of the Japanese village"; my catalog substantially extends this count and covers the full archipelago. The same active form appears in medieval Europe (Bailey 2014), precolonial Southeast Asia (Adas 1981; Scott 1985), and colonial India (Gandhi and Parel 1997, 95), suggesting the Tokugawa case is one window onto a broader phenomenon.

under home-neighbor dyad fixed effects. Decomposing the ratio into its two absolute distances reveals that proximity to a foreign capital is the active margin, confirming that access to a rival center of power, not limited oversight by the home domain alone, drives villages toward exit.

Third, the exit mechanism leaves a fiscal trace at the domain level: rulers more exposed to alternative authority levied lower tax rates, an association that strengthens where neighboring domains had greater capacity to absorb labor, consistent both with a theoretical prediction that geographic exposure constrains taxation when the outside option is attractive enough to make exit credible, and with the long-standing political-economy claim that interjurisdictional competition disciplines ruler extraction. The association is robust to extensive sensitivity checks, including controls for domain size, geography, and neighboring domains' characteristics; alternative distance measures based on least-cost walking paths and border proximity; Conley standard errors accounting for spatial correlation; and formal tests for omitted variable bias.

The paper makes three core contributions. First, it provides direct, micro-level evidence for a central behavioral premise running through four traditions in political economy, each resting on the same jurisdictional-conflict-to-exit link: the Great Divergence argument that ruler competition for mobile subjects checked extraction and fueled European growth (Jones 2003; Mokyr 2017); state-formation accounts in which exit forced ruler bargains and built representation (Dincecco and Wang 2018; Stasavage 2016); fiscal-federalism claims that competition among jurisdictions tames Leviathan through taxpayer and labor mobility (Rodden 2003; Tiebout 1956; Weingast 1995); and the cartel theory of the territorial state, in which modern borders are erected specifically to suppress subject exit (Acharya and Lee 2018; 2023). The evidence substantiates the crucial first link in this shared causal chain: fragmentation shapes the mode of popular resistance, thereby producing effective constraints on rulers.

Second, the paper brings exit from the background to the center of the study of state development. In Hirschman's (1970) terms, the dominant analytical traditions are built around voice: revolution and rebellion drive regime change (Acemoglu and Robinson 2001; Skocpol et al. 1982), collective mobilization makes states (Hintze 1975; Tilly 1978), and the threat of revolt disciplines extraction (Besley and Persson 2009; Fearon 2011). When exit appears, it shapes incentives (Besley and Persson 2009; Fearon 2011) or long-run sorting (Jones 2003) without being directly observed. Historians of everyday resistance have long documented otherwise: exit often predominated over revolt (Adas 1981; Bailey 2014; Camp 2004; Diouf 2014; Kolchin 1990; Scott 1976; 1985; Sreenivasan 2004). This paper documents exit directly as a strategy of resistance alternative to open rebellion.

Third, this paper contributes the community-level evidence on *horizontal* exit (movement between competing jurisdictions) contrasting with the *vertical* exit (flight beyond state reach) that dominates much of the existing literature on exit (Carneiro 1970; Herbst 2000; Pruett 2024; Scott 2009).[4] Within studies of horizontal exit, the Tiebout tradition interprets mobility primarily as individual preference-sorting rather than collective resistance (Peterson 1992; Tiebout 1956). Other strands conceive of exit as a spur to voice (Hirschman 1993; Pfaff and Kim 2003; Schoppa 2022), analyze government-driven emigration (Lueders 2025; Michel, Miller, and Peters 2023), or infer horizontal mobility indirectly—from the persistence of indigenous identity near jurisdictional havens (Guardado and Franco-Vivanco 2025), from legal efforts to restrict peasant movement (Matranga and Natkhov 2025), or from tax-base reactions to cross-border competition (Kleven et al. 2020)—but do not directly capture community-level exit as resistance. The Tokugawa case uniquely pairs direct observation of mass community exit with the polycentric institutional environment to which macro-level theories of state development refer.

2. A Model of Exit and Fight under Governance Conflict

To derive testable predictions, I develop a model in which geographically heterogeneous villages choose between exit and direct confrontation in response to taxation by competing rulers. Villagers weigh compliance against two forms of resistance: walking away to a neighboring jurisdiction (*exit*) or staying to fight collectively (*fight*). Exit is cheaper for

[4] Moore (1993, p. 230) characterized the late Tokugawa order as marked by "vertical and horizontal fissures… perhaps equally important," foregrounding the horizontal axis without operationalizing it; this paper supplies the community-level measurement.

villages near a rival capital; fight requires same-domain coordination, so every villager who leaves is one fewer body for collective violence. Rulers anticipate these choices and restrain extraction only when the neighboring jurisdiction is attractive enough that their villagers would actually walk. Three predictions follow: (i) greater exposure tilts the composition of resistance toward exit in equilibrium, even after the ruler's offsetting tax cut has absorbed part of the exit pressure; (ii) exposure raises the unconditional frequency of exit events; (iii) rulers more exposed to credible outside options levy lower taxes, but only when those options are attractive enough to be taken.[5]

2.1. Model Setup

Consider a society consisting of two neighboring domains 1 and 2. In each domain, a ruler governs a continuum of villages of measure 1.[6] The domains are indexed by $j \in \{1, 2\}$ and villages in domain j are indexed by $i_j \in [0, 1]$. The game begins with the two rulers simultaneously setting tax rates $\tau_j \in [0, 1]$ for their respective domains under complete information about each other's primitives. After observing the tax rate, each village decides whether to comply or pursue one of two outside options: fight or exit. Let $a_{i_j} \in \{\text{comply}, \text{fight}, \text{exit}\}$ denote village i_j's action. The proportion of villages choosing action $a \in \{\text{comply}, \text{fight}, \text{exit}\}$ in domain j is represented by $\mu_j^a \in [0, 1]$. Each village generates yield $y_j > 0$, which varies by domain. Compliant villages retain $y_j(1 - \tau_j)$, their net-of-tax yield. I assume that a compliant village i_j receives a random utility shock $\varepsilon_{i_j} \in [-\delta, \delta]$ with $\delta > 0$, which smooths the ruler's objective function.

A direct fight against the ruler is modeled as a collective action problem: its success depends on the fraction of villages participating, and specifically, the probability of

[5]The model extends three literatures. Relative to Tiebout fiscal-competition setups where mobility is costless sorting (Tiebout 1956), exit here is costly and heterogeneous across villages. Relative to single-organization exit-voice models (Gehlbach 2006; Hirschman 1970), the two-ruler structure makes the cross-domain best-response an equilibrium object. The closest precursor on the ruler side is Acemoglu and Wolitzky (2011), whose cross-domain best-response I endogenize rather than parameterize. On exit's effect on collective action, the model commits to substitution (Sellars 2019) rather than the complementarity views (Clark, Golder, and Golder 2017; Hirschman 1993; Karadja and Prawitz n.d.; Pfaff and Kim 2003), matching the Tokugawa case in which fight required within-domain coordination.

[6]I treat the village, rather than each subject, as the unit of decision making, consistent with Tokugawa Japan's governance structure, where villages were the fundamental units for taxation, negotiation with lords, and resistance.

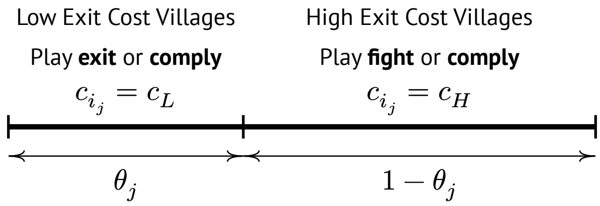


Figure 1: Exit Cost Determines Each Village's Mode of Resistance

success is μ_j^{fight}. If the fight succeeds, each participating village enjoys the full yield y_j without taxation; if it fails, they receive nothing. Choosing to fight incurs a sunk cost of $k > 0$ for each participating village.

Exit payoff is $y_{-j}(1 - \tau_{-j}) - c_{i_j}$, where $y_{-j}(1 - \tau_{-j})$ is the net-of-tax income in the other domain and $c_{i_j} \in \{c_L, c_H\}$ is the exit cost, with $c_H > c_L > 0$. The *outside-option quality* $y_{-j} > 0$ captures how attractive the neighboring domain is for potential migrants, which reflects the neighbor's economic capacity, governance conditions, and willingness to absorb newcomers.

Exit is modeled as a temporary bargaining strategy, not permanent relocation. Rather than modeling the sequential bargaining process between ruler and exiters, which would reduce tractability without additional insight, I assume rulers offer just enough concessions to induce returning villages' payoff to match the exit option, so all exiters return.[7]

The *geographic exposure* parameter $\theta_j \in [0, 1]$ denotes the share of villages in domain j that face the lower exit cost c_L; the remaining $1 - \theta_j$ face c_H. Under Assumption 1 (stated in Appendix A), high-exit-cost villages choose between fighting and compliance, while low-exit-cost villages choose between exit and compliance (see Figure 1). A greater θ_j thus means more villages with exposure to governance conflict.

Overall, the payoff of village i with exit cost $c_i \in \{c_L, c_H\}$ is summarized as

[7]The discussion of peasant resistance below details several historical examples of mass exits used as temporary bargaining tactics in Tokugawa Japan. Under a take-it-or-leave-it bargaining protocol in which the home ruler captures all the bargaining surplus, the village's equilibrium payoff is exactly $y_{-j}(1 - \tau_{-j}) - c_{i_j}$ as written: the ruler's concession is just enough to make the exit-threatening village indifferent, and the village's realized payoff coincides with the outside option by construction. Treating exit-threatening villages as contributing zero revenue to the home ruler is a first-order approximation (exact at the exit cutoff and mildly conservative for infra-marginal exit villages), and preserves the comparative statics of Propositions 1–3.

$$u_{i_j}\left(a_{i_j};\tau_j,\tau_{-j}\right)=\begin{cases} y_j\left(1-\tau_j\right)+\varepsilon_{i_j} & \text{if } a_{i_j}=\text{comply}\\ y_j\mu_j^{\text{fight}}-k & \text{if } a_{i_j}=\text{fight}\\ y_{-j}\left(1-\tau_{-j}\right)-c_{i_j} & \text{if } a_{i_j}=\text{exit}\end{cases}$$

The ruler's objective is to maximize net tax revenue, taking into account losses arising from resistance. To reflect the costs associated with quelling resistance, I assume that a ruler collects taxes only from villages within her own domain that comply. Importantly, the rulers do not collect taxes from the other domain's villages exiting to her territory; as described above, such villages always return to their home domain after gaining concessions from their original ruler. The utility of domain j ruler is thus given by

$$u_j(\tau_j,\tau_{-j})=\left[1-\mu_j^{\text{fight}}(\tau_j)-\mu_j^{\text{exit}}(\tau_j,\tau_{-j})\right]y_j\tau_j=\mu_j^{\text{comply}}(\tau_j,\tau_{-j})y_j\tau_j$$

for each $j\in\{1,2\}$.

For tractability, the baseline treats exit as reversible; permanent exit (exiters absorbed into the neighbor's tax base) imposes a lasting fiscal loss on the ruler, sharpening the tax-base-versus-extraction trade-off. Appendix A.9 develops this extension and shows the main results are robust. It also treats the ruler's tenure as secure; regime-collapse risk (she loses power once revolt passes a tolerance threshold; see Section 3.2) gives her a direct incentive to cap total revolt. Appendix A.7 and Appendix A.8 develop this extension and show the main results hold.

2.2. Empirical Implications

Each proposition below is stated at interior equilibrium and maps onto a specific empirical design developed in the next section; proofs are collected in Appendix A.

2.2.1. Partial-Equilibrium Shock Response

Before tax rates adjust, a shock that makes the neighboring domain more appealing raises exit without touching fight, an asymmetric response driven entirely by where each payoff gets its arguments from.

Proposition 1 (Resistance Composition under Exogenous Shocks): Under Assumption 1 and interiority of village cutoffs, for fixed tax rates $(\tau_j, \tau_{-j}) \in (0,1)^2$: (i) $\frac{\partial \mu_j^{\text{exit}}}{\partial y_{-j}} > 0$ and $\frac{\partial \mu_j^{\text{fight}}}{\partial y_{-j}} = 0$; (ii) $\frac{\partial \mu_j^{\text{exit}}}{\partial \tau_{-j}} < 0$ and $\frac{\partial \mu_j^{\text{fight}}}{\partial \tau_{-j}} = 0$.

The exit payoff $y_{-j}(1-\tau_{-j}) - c_{i_j}$ depends on the neighbor's yield and tax rate, so any improvement there pushes more low-cost villages past the exit-comply threshold; the fight payoff $y_j \mu_j^{\text{fight}} - k$ depends only on own-domain variables, leaving the fight margin unaffected. Proposition 1 is stated at fixed tax rates to match the shock-based design in Section 3.3.1, which exogenously shifts the outside option and recovers exactly this comparative static.

2.2.2. Equilibrium Composition and Exit Rate

In equilibrium, a village's geographic exposure to a rival capital shapes both the composition of resistance and the level of exit.

Proposition 2 (Geographic Exposure and Exit): Under Assumption 1 and the regularity conditions of Appendix A:

(a) *Exit share.* Whenever $y_{-j} > \kappa_j^*$ (the threshold defined in Proposition 3 below), the equilibrium exit share among non-compliant villages,

$$\sigma_j^*(\theta_j) := \frac{\mu_j^{*,\text{exit}}}{\mu_j^{*,\text{exit}} + \mu_j^{*,\text{fight}}}$$

is strictly increasing in θ_j.

(b) *Exit rate.* The equilibrium exit rate $\mu_j^{*,\text{exit}}(\theta_j)$ is strictly increasing in θ_j.

Higher θ_j means more villages face low exit costs: the exit pool expands, the fight pool shrinks. At fixed taxes, exit rises proportionally with θ_j. The ruler anticipates this and cuts τ_j^* (Proposition 3) to keep compliance attractive, but the cut is partial: going further would sacrifice per-compliant-village revenue faster than it would recover retention. The exit rate, exit share, and total revolt therefore all rise in equilibrium.

2.2.3. Equilibrium Tax Response

Whether the ruler's equilibrium tax response to exposure (to governance conflict) is to lower or raise taxes depends on whether the neighbor's outside option is attractive enough to make exit credible.

> Proposition 3 (Governance Conflict and Equilibrium Tax Rate): Under Assumption 1 and standard regularity conditions stated in Appendix A, for each $j \in \{1, 2\}$, there exists a unique threshold $\kappa_j^* > 0$ such that $\frac{\partial \tau_j^*}{\partial \theta_j} < 0$ if and only if $y_{-j} > \kappa_j^*$.

Rising θ_j shifts marginal villages from the fight margin to the exit margin. When $y_{-j} > \kappa_j^*$, exit-type villages have attractive outside options: the exit-constraint tightening dominates and the ruler lowers taxes. When $y_{-j} < \kappa_j^*$, exit is not credible and fight-pool shrinkage dominates, allowing the ruler to raise taxes: the "exit erodes voice" channel (Hirschman 1970; Sellars 2019). The empirical results are consistent with this conditional effect of governance conflict on extraction.

3. Empirical Evidence

3.1. Setting: Tokugawa Japan

During the Tokugawa period (1603–1868), Japan was divided into nearly 300 domains, each governed by a feudal lord (daimyō); Figure 2(a) maps the 264 active domain capitals in 1867, and Figure 2(b) plots the annual number of active capitals across the period as domains were created, merged, transferred, or absorbed into shogunal land. The origins of these domains lay in territories controlled by local warrior leaders during the Warring States period, or even earlier in private estates known as shoen. Formally, daimyō were subordinate to the Tokugawa shogunate, which asserted ultimate authority by establishing national policies, arbitrating inter-domain disputes, and receiving direct petitions from peasants. Yet, they continued to enjoy considerable autonomy within their own domains. In particular, each domain maintained its own government and military forces, and independently set tax rates (Ravina 1999, 1000).

Three features of this setting make it well suited for studying the relationship between governance conflict and exit as resistance. First, domains shared a broadly common administrative architecture that concentrated authority in the castle town serving as the domain capital (Hall 1955), with the village contract system (murauke) making villages directly responsible to the daimyō for tax collection (Saxonhouse 1995; Walthall 1997, 89; White 1995). This uniform, capital-centered structure justifies representing governance conflict parsimoniously by the geometry of domain capitals: a village's distances to its home capital and to the nearest rival capital capture the margin on which exit against alternative authority operated.

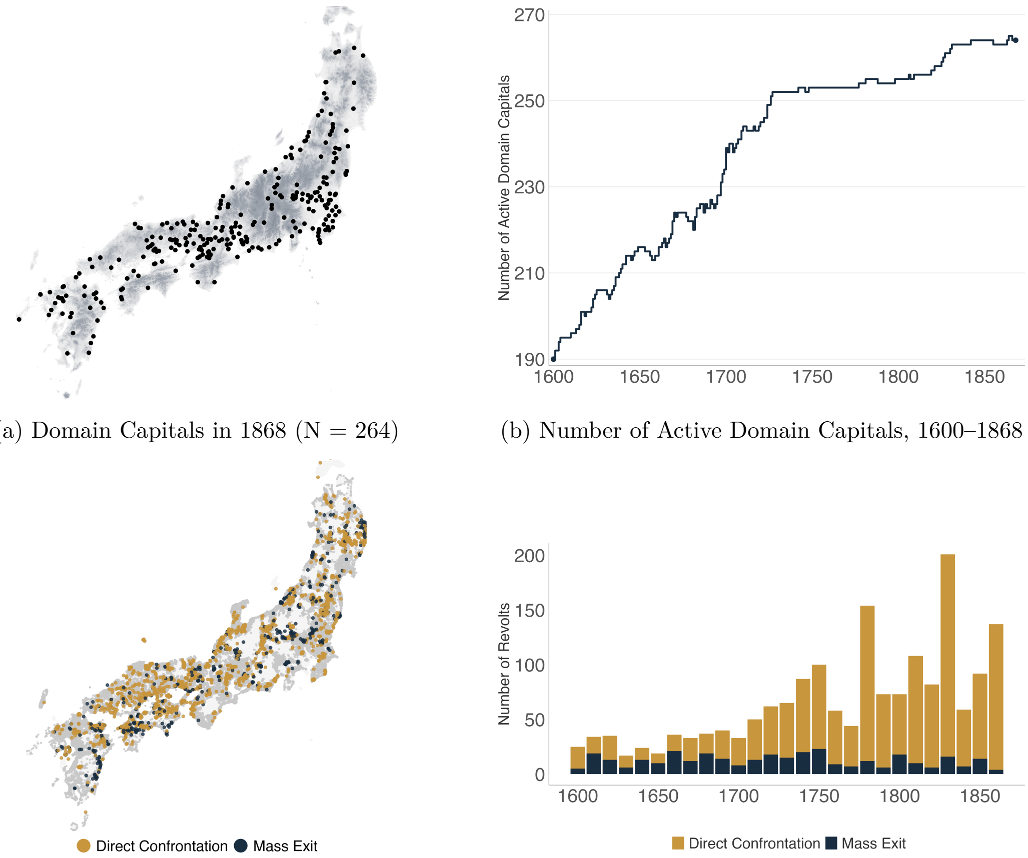
Second, the Tokugawa era saw no major armed conflict between domains until the shogunate's late-period collapse, marked by the Chōshū War (1864–1866) and the Boshin War (1868–1869).[8] Governance conflict among domains played out not through warfare but through competition over subjects and the "everyday forms of resistance" (Scott 1985) that this competition enabled, with external warfare playing a much smaller role than in the cases that dominate the traditional statebuilding literature (Hintze 1975; Tilly 1978; Weber 1919). This peace is a scope condition: Rosenthal and Wong (2011) theorize horizontal exit as a discipline on rulers but argue that, in early modern Europe, the mechanism was overwhelmed by war finance; the post-1615 Tokugawa peace shuts off exactly that war-finance channel, allowing the mechanism itself to be observed cleanly.

Third, the period affords an unusually rich documentary record of resistance, described below.

3.2. Peasant Resistance: Fight and Exit

Aoki's (1975) catalog identifies five forms of resistance: coercive appeals (gōso), destruc-

[8]The Shimabara Rebellion of 1637–38 is a mid-period exception but was an anti-Christian and anti-Matsukura uprising suppressed by a shogunal-led coalition of western daimyō, not a free-form inter-*daimyō* war. The Chōshū War was similarly a shogunate-versus-Chōshū conflict that pulled several daimyō into an ad-hoc coalition, not decentralized inter-domain warfare; the Boshin War that followed was a regime-collapse sequence. Between Shimabara and the 1860s, institutional peace held. My tax rate data are drawn from the end of the Tokugawa period and could in principle reflect Chōshū- or Boshin-era disruptions; reassuringly, historical evidence indicates that tax policies remained highly stable throughout the period, including its final years (Smith 1958).



(a) Domain Capitals in 1868 (N = 264)

(b) Number of Active Domain Capitals, 1600–1868

(c) Spatial Distribution of Peasant Revolts

(d) Temporal Trend of Peasant Revolts

Figure 2: Domain and Peasant Revolt Data

Note: (a) Active domain capitals in 1867 (black points) over Japan's terrain. (b) Annual number of active capitals, 1600–1868. (c) Revolt villages, colored by mode of resistance (exit vs. fight); gray = non-revolting villages extant through end of period. (d) Decade-by-decade revolt frequency.

tion (uchikowashi), rebellion (hōki), desertion (chōsan), and end-run appeals (osso).[9] I consolidate these into two categories: "fight" strategies (direct confrontation through coercive appeals, destruction, and rebellion) and "exit" strategies (desertions and end-run appeals that sought protection from an authority outside the home domain). Figure 2 (c) presents the spatial distribution of peasant revolts, and (d) presents the frequency of revolts over time.

The three fight subcategories differed in intensity (Aoki 1975). Coercive appeals were mass marches on domain offices or castles — crowds of thousands donning straw raincoats

[9]Subsequent studies follow his categorization: White (1995) provides a quantitative treatment of protest repertoires and their evolution; Vlastos (1986) treats end-run appeals (osso) as a distinct protest form that dominated the mid-Tokugawa period; Walthall (1997) translates primary documents from representative episodes across all five categories.

and carrying bamboo spears as ritual symbols of collective resolve (Walthall 1997, 131–132) — to present demands backed by the implicit threat of violence. Destruction involved targeted attacks on the property of officials, merchants, or wealthy villagers perceived as complicit in oppressive policies; participants organized into groups to systematically "smash" (*uchikowashi*) houses and storerooms (Walthall 1997, 89–130). Rebellion was armed uprising against domain authority, the rarest form, accounting for roughly 1% of recorded incidents (White 1995, 142). All three sometimes won concessions but often met severe repression.

The exit coding combines two routes: desertion relied on physical relocation toward a neighboring jurisdiction, whereas end-run appeals involved formal petitioning of an external authority — whether a neighboring lord or the shogunate — to intervene on the villagers' behalf. Both share the defining feature that distinguishes exit from fight: the protester seeks protection from an authority outside the home domain, so the home ruler faces the threat of losing subjects or legitimacy to an external actor.

3.2.1. Revolts Were Consequential

Both exit and fight carried the risk of attainder (*kaieki*) for the home daimyō. The Buke Shohatto (Laws of Military Houses) required domains to be "well-governed" (Steele, Paik, and Tanaka 2017, 3–5), and governance failure served as a recognized pretext for the shogunate to strip a lord of his domain, reduce his territory, or forcibly transfer him to a smaller holding. The two forms of resistance differed, however, in their *publicizing* risk. Fight typically remained a domain-internal matter: information about mass assemblies or property destruction could be contained or reported through channels the daimyō controlled. Exit, by contrast, crossed domain boundaries by construction: fleeing villagers placed themselves under the protection of a rival lord or a shogunal intendant, who then had their own reasons for reporting the episode. A lord whose peasants fled therefore advertised his governance failure directly to rival lords and to the shogunate, making the same underlying attainder risk much more difficult to suppress (Steele, Paik, and Tanaka 2017). This asymmetry in publicity explains why daimyō had especially strong incentives to preempt exit by moderating extraction, particularly in border villages with credible access to outside authority. Consistent with this, mass flight was "what the warrior class feared most" (Kuromasa 1927, 69) and was prohibited by strict decree.

3.2.2. Exit as Deliberate Protest

Both exit routes, desertion and end-run appeals, depended on proximity to outside authority.[10] The two tactics were frequently deployed together in the same episode: village representatives would present petitions to external authorities while other members sought refuge in safe areas (Hosaka 2002; Miyazaki 1995).

Both served as bargaining tactics rather than permanent relocations. In a typical exit protest, some representatives traveled to file a petition with an outside authority while others took refuge in nearby foreign territory (Hosaka 2002; Miyazaki 1995). What exiters sought from outside officials was not immediate permanent resettlement: the outside authority could offer temporary shelter, guards, mediation, or formal recognition of a petition, and the possibility of a durable jurisdictional transfer mattered mainly because it strengthened bargaining leverage. Many episodes therefore ended not with permanent migration but with concessions from the home domain and a negotiated return (Aoki 1975; Miyazaki 1995). This bargaining-tactic character justifies the setup of my theoretical model, in which exit is a threat of departure resolved by concessions from the home ruler rather than a one-way relocation. Desertion and end-run appeals were, in short, deliberate protests, not passive flight (Hosaka 2002; Vlastos 1986).

End-run appeals in practice typically routed through neighboring rulers, who provided shelter, mediation, or escalation to the shogunate, rather than directly to Edo. Direct-to-Edo petitions, the Sakura Sōgorō pattern, are the well-known exception rather than the modal form (Walthall 1997). The three illustrative cases below all follow the neighbor-routed pattern: in Nobeoka 1690, Takanabe first mediated and then escalated to Edo; in Yamashirodani 1837, Imabari sheltered the villagers and offered absorption; in Matsuyama 1741, Ōzu forced concessions directly (Hosaka 2002; Miyazaki 1995). The

[10]Some cases coded as desertion in Aoki (1975) were passive exits known as "hashiri," where small groups fled secretly without political coordination or external recognition (Hosaka 2002). I retain these in the statistical analysis because they share the same strategic feature as deliberate chōsan: even passive flight required recognition from an external authority (without it, domain officials could track down fugitives and forcibly repatriate them), so success hinged on access to a neighboring jurisdiction.

distance to the nearest foreign capital d_i^{FC} is therefore the operative distance for both desertion and end-run appeals, since both depend on access to a rival ruler. Pooling the two tactics under "exit" is consistent with the theoretical mechanism: both depend on access to a rival ruler, not on access to the shogunate in the abstract.

3.2.3. How Exit Worked: Illustrative Cases

Three episodes from the secondary literature on *hyakushō ikki* (peasant uprising) illustrate how exit worked in practice: fleeing villagers secured shelter, mediation, and help escalating appeals to the shogunate from neighboring rulers, which made the threat of permanent relocation credible enough to extract concessions from the home lord.

The Nobeoka case (1690) illustrates how exit could coincide with the collapse of a ruling family. After years of oppressive taxation compounded by repeated flooding, peasants in Nobeoka Domain (Hyūga Province, southeastern Kyushu) fled south to the neighboring Takanabe Domain. Takanabe officials halted the fleeing villagers, offered to mediate, and — when negotiations reached an impasse — helped escalate the case to the shogunate in Edo, providing guards to escort village representatives and shelter for those who remained. Nobeoka acquiesced to the villagers' demands, including tax relief, and imposed no punishment. The following year, the shogunate stripped the Nobeoka lord Arima Kiyozumi of his domain and transferred him to the far smaller Itoigawa Domain in Echigo Province, effectively ending the Arima family's prominence (Ishikawa 1981).

The Yamashirodani case (1837) illustrates how neighboring lords actively facilitated exit, turning inter-domain competition into a credible threat. Villages in a mountainous district on the border between Tokushima Domain (Awa Province, Shikoku) and Iyo Province crossed into Imabari Domain. Rather than turning the peasants away, the Imabari lord dispatched guards to shield them from Tokushima's officials and facilitated negotiations, even warning that he was prepared to accept the villages into his own jurisdiction should Tokushima refuse their petition (Okamura 2009). The receiving domain had its own reasons to cooperate: absorbing productive villages meant more tax revenue

(Miyazaki 1995).[11]

The Matsuyama case (1741) shows exit operating as bargaining leverage rather than permanent relocation. After famine left them unable to pay taxes, several villages in Matsuyama Domain (Iyo Province, Shikoku) fled upriver to the neighboring domain of Ōzu and appealed for permission to migrate. Upon arriving in the Ōzu castle town, their strategy proved effective: the Matsuyama lord ultimately accepted most of their demands. Rather than punishing the revolt's leaders, the domain exiled its own chief minister to an island, attributing the crisis to his mismanagement (Nakagawa 1980). The villagers returned home.

The statistical analysis treats villages as strategic decision units choosing among compliance, fight, and exit based on distance, taxation, and collective-action technology. A substantial specialist literature (Vlastos (1986) on the political economy of benevolence, Walthall (1997) on ritualized mass petitioning, and White (1995) on the "referents of consciousness" invoked by protesters) emphasizes that peasants framed their actions in terms of customary legitimacy, honor, and ritual. The three illustrative cases above involve such elements: oaths, the role of the *otsukai* village representative, and the mino-and-spear symbolism of *gōso*. The strategic framing adopted here is appropriate as a first-cut model for testing aggregate statistical patterns across 910 revolts; it is not a claim about the complete phenomenology of Tokugawa protest. The moral-economy account and the strategic account are complementary: customary legitimacy shapes which actions are available in a village's repertoire, and relative costs among available actions shape which is selected.

3.3. Exposure Raises Exit Occurrence

Two designs test whether exposure to alternative authority raised the occurrence of exit.

[11]The Imabari–Tokushima episode is documented in Okamura (2009) and is consistent with the broader pattern of inter-domain labor competition documented for mountain border districts (Hosaka 2002; Miyazaki 1995).

3.3.1. Shock-Based Test: Tenryō Conversion

I leverage a quasi-experimental shock that introduced a powerful new competitor in close proximity: the Tokugawa shogunate's abolition of certain domains and the subsequent conversion of their territories into direct shogunal land (*tenryo*).

A tenryō conversion heightened governance competition for neighboring domains through several reinforcing channels. Tax burdens in tenryō were, on average, lower than in daimyō domains (Murakami 1996; Sng and Moriguchi 2014), with a few exceptions (Vlastos 1986). The installed intendants held short, rotating tenures and lacked independent military power (White 1995), reducing their investment in aggressive extraction. Contemporary peasants recognized and acted on these differences: White (1995) documents "a general perception that the administrative hand weighed lighter" in shogunal lands, such that calls for shogunal-style governance became "a component of peasant appeals," and records a significant number of incidents in which the people of an area demanded transfer to shogunal control (White 1995, 110). Moore (1993, p. 242) reaches the same conclusion in synthesis: "the hold of the authorities over the peasants was weaker in areas directly controlled by the Shōgun than in some of the outlying fiefs." A recent nearby conversion also served as precedent, signaling that the shogunate was willing to absorb additional territory. End-run appeals also became cheaper: with a shogunal intendant (*daikan*) newly installed at the former domain capital, nearby villagers could petition locally rather than undertake the costly journey to Edo.

I exploit these conversions to test Proposition 1's partial-equilibrium prediction, which captures the effect of a neighbor shock before the home ruler has optimally adjusted his own tax rate in response: a shift that raises y_{-j} and lowers τ_{-j} should raise exit but not fighting in foreign villages near the converted capital. Because conversions were sometimes triggered by internal disorder in the converted domain itself, the null on nearby fight doubles as a placebo that rules out a generic spillover of unrest from converted territories onto their neighbors.

I draw on the revolt data compiled by Aoki (1975), which catalogs 7,664 incidents of peasant resistance from local histories, official documents, and diaries.[12] Each observation records the date, home domain, nearest neighboring domain,[13] and mode of resistance classified as "fight" or "exit" following the categorization described in Section 3.2; I exclude 574 incidents classified as peaceful petitions and inter-village disputes (resource and boundary conflicts between villages rather than against the lord). I georeference each revolt village using several sources,[14] identifying the location of 2,571 villages out of the 2,790 sites recorded in Aoki (1975) the remaining sites are excluded due to imprecise location information. The resulting dataset contains 910 revolt incidents involving 2,571 villages and 2,741 tactic choices (summary statistics in Table A.5, Appendix B). As with most historical datasets, questions of completeness naturally arise; large and consequential mass exits are much more likely to have been preserved in the historical record than smaller-scale flights, so the absence of minor cases is unlikely to bias the core conclusions. I complement the revolt data with capital locations and creation and abolition dates for all 294 domains from Kimura, Fujino, and Murakami (2015) and Kodama and Kitajima (1989) (summary statistics in Table A.6, Appendix B); branches (shi-han) are treated as part of their main domains (hon-pan).

From these data I construct an annual domain-year panel covering 1615 to 1868. For each domain and year, I count exit and fight revolts occurring in villages ruled by other domains but located within a fixed radius of the domain capital. The main specification uses a $r \in \{30\text{km}, 40\text{km}, 50\text{km}\}$ radius around each capital to define the zone in which foreign villages are exposed to the new exit opportunity created by conversion. Figure 3 illustrates the design.

I exclude the four abolitions in which territory passed directly to another domain instead of to tenryo. The resulting 55 treated domains, enumerated with year and stated

[12]Hosaka (2002) characterizes Aoki (1975) as the culmination of postwar Japanese research on peasant uprisings and "one of the largest contributions in Japanese historical studies after WW2."

[13]I used Kodama and Kitajima (1989) to identify the neighboring domains existing at the time of each revolt. The nearest foreign domain is identified using time-varying domain existence data: for each revolt, only domains that were active at the time of the incident are considered as potential alternative authorities.

[14]These include Honda, Natsume, and Nemoto (2022), the geocoding API provided by the Ministry of Land, Infrastructure, Transport and Tourism of Japan, and the Historical Location Name Comprehensive Database provided by the Center for Open Data in the Humanities.

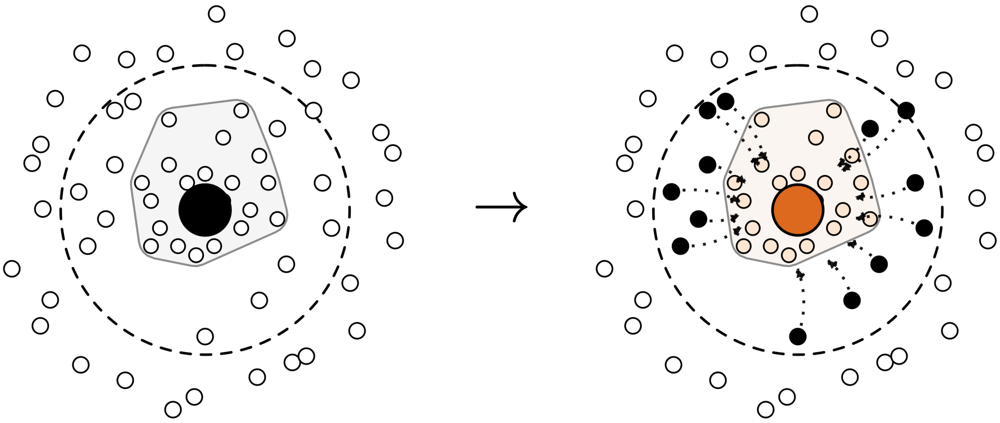


Figure 3: DiD Design: Tenryō Conversion

Note: Stylized illustration. Left: a daimyō domain (gray territory) extracts taxes; foreign villages within radius r of the capital (solid black dots outside the territory) face no attractive alternative authority. Right: abolition converts the same capital into shogunal territory (*tenryō*, orange), which was on average less oppressive; those same foreign villages can now exit toward the converted territory (dotted arrows). The DiD compares changes in exit and fight outcomes in these treated foreign villages against never-converted neighbors.

reason in Appendix C.1, are drawn from 79 dissolutions recorded across the 294-domain universe in Kimura, Fujino, and Murakami (2015) and Kodama and Kitajima (1989). For converted domains, I then restrict the sample to a symmetric event window from five years before to five years after the first tenryo-conversion year, keeping only strict pre-conversion years and years in which the domain is actually dissolved under direct shogunal rule. Never-converted controls are kept only in the same calendar years represented by the converted observations in each subset. I estimate annual domain-level linear probability difference-in-differences with domain and year fixed effects:

$$\text{Revolt}_{j,t} = \beta_1 \big(\text{Converted}_j \times \text{Post}_{j,t}\big) + \varphi_j + \omega_t + \varepsilon_{j,t}$$

In this specification, $\text{Revolt}_{j,t} \in \big\{\text{Exit}_{j,t}, \text{Fight}_{j,t}\big\}$ represents a binary indicator for the occurrence of any nearby exit or fight within r kilometers of domain j's capital in year t. Converted_j denotes domains that ever undergo a tenryo conversion within the sample, while $\text{Post}_{j,t}$ marks years following the first such conversion when the domain is under direct shogunal administration.

A caveat is that the destination of each recorded exit is largely unknown: the sources note where fleeing villagers came from but rarely how and where they traveled. Proximity to the converted capital therefore proxies for the converted territory being a plausible destination rather than directly measuring it. Two features of the results are reassuring. First, tightening the radius from 50km to 40km and 30km preserves the sign and significance of the exit coefficient on progressively more spatially precise exposure bands. Second, the placebo null on nearby fight holds at every radius, which is difficult to reconcile with a story in which the exit coefficient merely picks up geographically mismeasured unrest.

Table 1 shows a positive nearby-exit response. In the main 50km specification, tenryo conversion raises the probability of any nearby exit by 0.0385 ($p = 0.010$), about 3.9 percentage points, or roughly 59% of the control-group pre-conversion mean (0.065), while the corresponding any-fight estimate is only 0.0076 ($p = 0.532$). When a nearby domain was replaced by direct shogunal rule, surrounding villages became more likely to run than to stay and fight.

Outcome:	Any Nearby Exit			Any Nearby Fight		
Sample Radius:	30km	40km	50km	30km	40km	50km
Variables						
Converted × Post	0.0163	0.0262	0.0385	0.0084	0.0033	0.0076
Cluster SE	(0.0078)**	(0.0116)**	(0.0147)***	(0.0093)	(0.0092)	(0.0122)
Conley SE (70km)	(0.0084)*	(0.0107)**	(0.0131)***	(0.0104)	(0.0108)	(0.0136)
Controls						
Domain FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Fit Statistics						
Num. Obs.	33,877	33,877	33,877	33,877	33,877	33,877
Tenryo-Converted/All Domains	55/294	55/294	55/294	55/294	55/294	55/294
R2	0.038	0.052	0.062	0.078	0.101	0.136
Mean Dep. Var.	0.010	0.018	0.026	0.020	0.031	0.044

Table 1: Domain-to-Tenryo Conversions Increase Nearby Exit

Note: LPM estimates on an annual domain-year panel. "Nearby" denotes revolts by villages of *other* domains within the specified radius. Converted domains restricted to rel_year in [−5, +5], matched to never-converted controls in the same calendar years. SEs in parentheses: (i) two-way clustered (domain, year); (ii) Conley (70 km) on capital coordinates; stars reflect that estimator's p-value. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

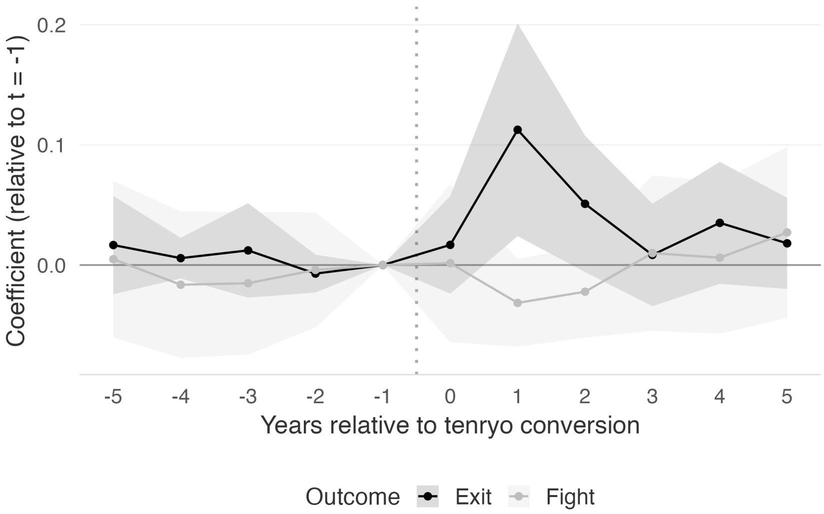


Figure 4: Tenryo-Conversion Event-Study Coefficients

Note: Year-by-year event-study coefficients with $s = -1$ as the reference period for both nearby exit (top) and nearby fight (bottom). Shaded bands show 95% confidence intervals. "Nearby" denotes revolts by villages belonging to *other* domains but located within 50km of the focal domain's capital. Two-way clustered (domain & year) standard errors.

The identification assumption is that converted and never-converted domains would have followed parallel revolt trends absent conversion. I assess this with an event study:

$$\text{Revolt}_{j,t} = \sum_{s \neq -1} \beta_s \big(\text{Converted}_j \times \mathbf{1}\left[t - t_j^* = s\right]\big) + \varphi_j + \omega_t + \varepsilon_{j,t}$$

where Converted_j denotes tenryo-converted domains, t_j^* is the first conversion year, and $s = -1$ is the reference period. As shown in Figure 4, pre-treatment coefficients are individually small and a joint F-test confirms their insignificance ($F = 1.26$, $p = 0.283$; fight: $F = 0.18$, $p = 0.947$). After conversion, the exit series jumps sharply at $t = 1$ ($\hat{\beta}_1 = 0.113$, $p < 0.05$) and remains elevated throughout the window, while the fight series stays flat. A Rambachan and Roth (2023) relative-magnitudes sensitivity analysis yields a breakdown value of $\overline{M}^* = 0.20$: the exit effect tolerates post-period deviations up to 20% of the worst pre-period deviation (Appendix C.7). At canonical benchmarks $\overline{M} \in \{0.5, 1, 2\}$ and under the absolute-smoothness restriction Δ^{SD} evaluated at the pre-period standard deviation σ_{pre}, the signed effect is preserved though confidence intervals widen (Appendix C.8). A complementary identification check models the timing of tenryo conversion directly: lagged nearby exit and fight intensities do not predict which domain the shogunate converts in a given year (joint Wald $p = 0.344$, Appendix C.11).

The results are robust to alternative specifications and estimation strategies. Because the 55 conversions are staggered across more than two centuries, the standard TWFE estimator could be biased if treatment effects vary across cohorts (Baker, Larcker, and Wang 2022). A Goodman-Bacon decomposition places 94.5% of the TWFE estimand weight on clean treated-versus-never-treated comparisons, leaving the "bad-comparison" problem of heterogeneous-timing DiD immaterial here (Appendix C.5). Consistently, the interaction-weighted estimator of Sun and Abraham (2021), which is robust to staggered-timing heterogeneity, yields an aggregate exit ATT of 0.0413 ($p < 0.001$), virtually identical to the TWFE estimate of 0.0385, while the fight ATT remains insignificant (0.0089, $p = 0.262$); the full comparison appears in Appendix C.4. A cohort-split Sun-Abraham estimate concentrates the exit effect in the pre-1750 cohort, which holds 49 of 55 conversions (ATT $= 0.044$, $p < 0.001$), with the two later cohorts too small to be load-bearing (Appendix C.6).

The tighter 30km and 40km bands confirm the same pattern, with exit coefficients significant at 0.0262 ($p = 0.025$) and 0.0163 ($p = 0.038$) and fight coefficients indistinguishable from zero across all radii. Conley standard errors at bandwidths from 30 to 150 km deliver p-values between 0.0002 and 0.0033 for the 50 km exit estimate; the Kelly (2019) data-driven bandwidth selects a 10 km cutoff that produces inference within 0.001 of the hand-picked 70 km bandwidth (Appendix C.12, Appendix C.13). Replacing the binary outcome with exit and fight counts produces consistent estimates (Appendix C.2). A 500-iteration permutation test that randomly reassigns fake conversion years to never-converted domains yields a two-sided p-value of 0.012 for the actual estimate; a more demanding variant that instead draws 55 placebo conversions from the full pool of 294 domains (including the actually-converted 55) yields $p = 0.006$ (Appendix C.10). Both variants confirm the response is specific to the actual timing and location of conversions, not an artifact of the panel's staggered-adoption structure. Because some conversions cluster spatially, I also verify that the result is not driven by interference among overlapping treatment radii: dropping all converted domains within 100 km and 5 years of another conversion, or excluding controls within 50 km of any converted capital, preserves the sign and magnitude of the exit coefficient (Appendix C.9). The result also holds when exit is restricted to desertion alone, a narrower definition that excludes end-run appeals (Appendix C.3).

The effect loads on the regime that best fits the mechanism. The cohort split (Appendix C.6) shows the exit response concentrated in the pre-1750 cohort, which holds 49 of 55 conversions. This is consistent with the specialist view that the attainder threat, the primary publicizing risk created by exit, was most credible in the first century of Tokugawa rule, when the shogunate actively disciplined daimyō (Ravina 1999; White 1995). Later conversions, particularly after the Tenmei and Tempō reforms of the 1780s and 1840s, operated in a different political environment: weakened shogunal authority, rising commercial taxation, and broader fiscal strain. The headline effect therefore loads primarily on the regime in which the attainder-publicity mechanism had the most bite.

3.3.2. General Test: Grid-Year Panel

The DiD identifies exit's response to a specific shock. A complementary general test asks whether geographic exposure to governance conflict, varying continuously across locations and across years as nearby domain capitals are established, abolished, or relocated, predicts exit occurrence, as predicted by Proposition 2(b).

I operationalize exposure to governance conflict at the location level as a *distance ratio*:

$$\tilde{\theta}_{i,t} := \frac{d^1_{i,t}}{d^2_{i,t}}$$

where $d^1_{i,t}$ and $d^2_{i,t}$ are the straight-line distances from location i to its nearest and second-nearest active domain capitals in year t. The ratio rises toward one whenever a new capital appears closer than the existing second-nearest — as when a neighboring domain is dissolved or a new one established — or whenever the closest capital becomes more distant. The ratio thus captures the relative accessibility of alternative authority a location enjoys: a value near zero indicates one dominant nearest capital, while a value near one indicates two equally accessible competing rulers. A village-level analog adapted for the cross-sectional analyses below appears in Section 3.4.1.

A village-level panel would be the natural unit for this test, but domain affiliation is observed only at the end of the Tokugawa period: for most non-revolting villages, the governing domain at any given date is unknown, making a village panel with time-varying distance ratios impossible. A grid-year panel sidesteps this by assigning villages to fixed 5 km hexagons and computing $\tilde{\theta}_{i,t}$ from the two nearest active capitals each year, without requiring domain affiliation. Populated cells are defined from a 46,086-village catalog compiled from the late-Tokugawa *Kyūdaka–Kyūryō Torishirabe Chō* (LFAT), described more fully in Section 3.4; each village is assigned to the 5 km hexagon containing it. The panel retains 2,225,802 cell-years across 8,763 occupied cells, and each observation counts the number of exit and fight revolts within that hexagon in the given year.

The estimating equation is

$$\text{Revolt}_{i,t} = \beta_3 \tilde{\theta}_{i,t} + \lambda_i + \eta_t + \varepsilon_{i,t}$$

with hexagon and year fixed effects. Here $\text{Revolt}_{i,t}$ is either the exit count or the fight count. Identifying variation comes from capital dissolutions, transfers, and reestablishments and the resulting reshuffling of the nearest outside authority — not from a single event. Figure 5 illustrates this variation by mapping $\tilde{\theta}_{i,t}$ across three snapshot years. The overall pattern is stable, but localized changes are visible: in areas where new domains were established or existing ones dissolved, governance conflict shifts as hexagons gain or lose proximity to the closest or the second-closest capital. Hexagon and year fixed effects absorb time-invariant geography and common shocks.

Table 2 shows that a higher distance ratio predicts exit incidence but not fight incidence, and that mere remoteness from any state authority predicts neither. Columns (1)–(3) report LPM estimates with the time-varying distance ratio: a one-unit increase is associated with a 0.035-percentage-point higher probability of any exit in a cell-year, statistically significant at the 5% level under both two-way cell-and-year clustering ($p = 0.048$) and Conley spatial inference ($p = 0.037$). The corresponding fight coefficient is smaller, positive, and not statistically significant. The exit coefficient is small in absolute

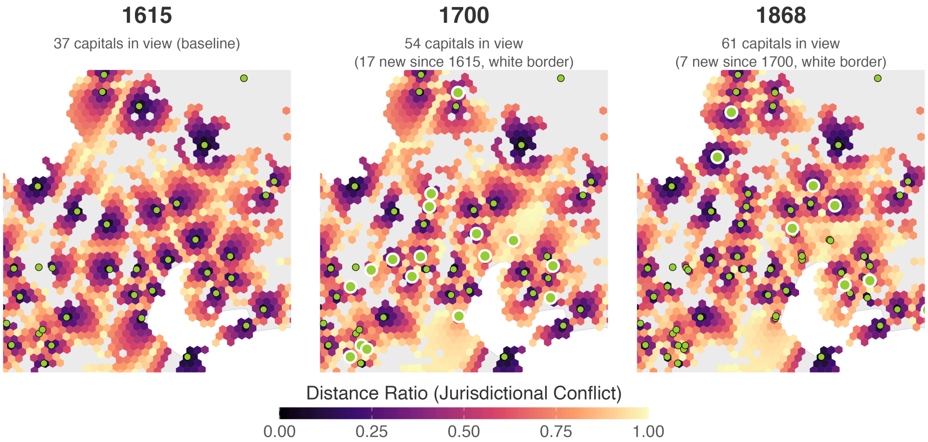


Figure 5: Time-Varying Distance Ratio across 5km Hexagonal Grid: Kinki Region

Note: Each panel maps $\tilde{\theta}_{i,t}$ (distance to nearest active capital / distance to second-nearest) for a given year, zoomed to a $\approx 200\,\text{km} \times 200\,\text{km}$ window in central Japan (lat 34.4–36.2°N, lon 135.5–137.6°E) covering Kyoto, Osaka, Nagoya, and Lake Biwa — the region with the highest concentration of cells whose distance ratio changed across the three snapshot years. Yellowgreen circles mark active capitals; white borders distinguish capitals newly added since the previous snapshot. Higher values (yellow) indicate strong governance conflict; lower values (purple), a dominant nearest capital. Spatial shifts reflect domain dissolutions and creations.

	Distance ratio ($\tilde{\theta}_{i,t}$)			Closest-capital dist. ($d^1_{i,t}$, km)		
Outcome:	(1)	(2)	(3)	(4)	(5)	(6)
	Any exit	Any fight	Any fight (strict)	Any exit	Any fight	Any fight (strict)
Coefficient	0.00035	0.00026	−0.00012	0.00000	0.00000	0.00000
Cluster SE	(0.00017)**	(0.00019)	(0.00011)	(0.00000)*	(0.00000)	(0.00000)
Conley SE (10km)	(0.00017)**	(0.00023)	(0.00012)	(0.00000)	(0.00000)	(0.00000)
Controls						
Grid FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Fit Statistics						
Num.Obs.	2,225,802	2,225,802	2,225,802	2,225,802	2,225,802	2,225,802
R^2	0.006	0.007	0.006	0.006	0.007	0.006
Mean Dep. Var.	0.00022	0.00064	0.00026	0.00022	0.00064	0.00026

Table 2: Governance Conflict Predicts Exit

Note: LPM on binary any-event-of-type indicators at the 5 km hexagon–year level for the full panel of 8,763 cells. Columns (1)–(3) use the time-varying distance ratio $\tilde{\theta}_{i,t} = \frac{d^1_{i,t}}{d^2_{i,t}}$ (computed from the two closest active capitals each year). Columns (4)–(6) replace it with the raw distance to the closest active capital $d^1_{i,t}$ (km), a measure of mere remoteness from any state authority that discards information about the second-nearest. "Fight (strict)" restricts fight to destruction and uprising. Grid and year FE. SEs: (i) two-way clustered (cell, year); (ii) Conley (10 km). The ever-active-cells subsample analysis appears in Appendix D.2.12. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

terms — the baseline probability of any exit in a cell-year is only 0.022% — but large relative to this base: a one-standard-deviation increase in the distance ratio (SD = 0.26) corresponds to roughly 40% of the baseline mean.

Columns (4)–(6) replace the distance ratio with the raw distance to the closest active capital $d^1_{i,t}$, a measure of mere remoteness from any state authority that discards information about the relative position of the second-nearest. The closest-capital coefficient on exit is small and only marginal under clustered SE ($p = 0.092$) and gets null under Conley SE ($p = 0.162$); the fight and fight-strict coefficients are likewise indistinguishable from zero. The contrast between columns (1) and (4) isolates the active ingredient: the relative geometry of competing rulers, not the absolute distance to the nearest one, is estimated precisely enough to predict exit. Restricting the panel to ever-active cells (the 1,466 hexagons with at least one event between 1615 and 1868) preserves the result while statistical significance attenuates somewhat (Appendix D.2.12). The result is robust to count-data estimators (Poisson and Negative Binomial FE; Appendix D.2.11), restricting exit to desertion only (Appendix D.2.10), replacing great-circle distance with a terrain-weighted least-cost-path ratio (Appendix D.2.13), and including daikansho in the distance computation (Appendix D.2.15), though statistical significance attenuates somewhat in each.

3.4. Exposure Shifts Revolt Composition Toward Exit

The two preceding tests establish that exposure to governance conflict raises exit occurrence. A distinct question is whether, among villages that did revolt, greater exposure shifted the chosen form of resistance toward exit over direct confrontation, as predicted by Proposition 2(a).

The village-level analyses here and below (tax rates in Section 3.5) draw on a dataset of 46,086 villages compiled primarily from the *Kyūdaka–Kyūryō Torishirabe Chō* (Ledgers of Former Assessments and Territories, hereafter LFAT), a comprehensive Meiji-government land census that records the home domain for each village at the end of the Tokugawa period; village location data are obtained from Honda, Natsume, and Nemoto (2022). When multiple domains collected taxes from a single village, I assign it to the domain possessing the largest share of its assessed productivity (*kokudaka*, a rice-denominated index of aggregate agricultural output rather than literal rice harvest alone). The resulting data encompass virtually all of Japan's arable land, omitting only high-altitude, non-cultivable regions, as illustrated in Figure 2(a)–(b) and Figure 7.

3.4.1. Measuring Exposure to Alternative Authority

The theoretical parameter θ_j is a domain-level binary share: the fraction of villages facing low exit costs. I operationalize this at the village level as a continuous index $\hat{\theta}_i$, measuring each village's geographic exposure to alternative authority. The logic is that villages closer to an outside authority face lower exit costs because proximity facilitates the external intervention on which mass exit depended (Section 3.2): it reduced the cost of dispatching guards or mediators to protect fleeing peasants, made it easier for villagers to reach neighboring capitals or shogunal offices to voice grievances, and made administrative transfer of allegiance more practicable. I therefore construct a parsimonious distance-based measure:

$$\hat{\theta}_i := \frac{d_i^{\text{HC}}}{d_i^{\text{FC}}}$$

where d_i^{HC} denotes the straight-line (Euclidean) distance from village i to its home-domain capital, and d_i^{FC} is the straight-line distance to the nearest non-home domain capital that was active at the time of the revolt.[15] Both distances are calculated using georeferenced village locations from the revolt data and the time-varying domain-capital panel described in Section 3.3.1. A capital is considered "active" in a given year if it served as the seat of an existing daimyō domain, or if the domain had been dissolved and converted to tenryo, in which case the daimyō was replaced by a shogunal intendant (*daikan*) (Vlastos 1986, 49). In the latter scenario, the capital remains in the choice set for d_i^{FC} throughout the tenryo period.[16]

[15]Robustness checks employing least-cost path distances that incorporate terrain are reported in Appendix D.2.4.

[16]In some cases, the shogunate may have administered converted territory from a nearby *daikansho* (a shogunal intendant's district office) rather than from the former castle town. To address this, Appendix D.2.14, Appendix D.2.15, and Appendix D.3.1 replicate all main analyses with the nearest capitals redefined to include *daikansho* locations. Nevertheless, Aoki's destination data indicate that

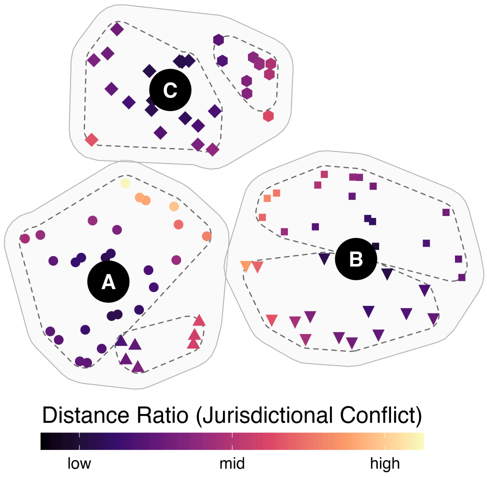


Figure 6: Dyad FEs Enable Within-Pair Comparison

Note: Three stylized domains (A, B, C) with capitals (black circles) and villages (shapes) shaded by the distance ratio ($\hat{\theta}_i$): dark shades indicate low exposure (villages near their home capital), light shades indicate high exposure (villages near a rival capital). Home-neighbor dyad fixed effects restrict comparisons to villages sharing both home and nearest-rival domains (those within the same dashed circles).

3.4.2. Empirical Strategy: Home-Neighbor Dyad Fixed Effects

Each observation corresponds to a village involved in a revolt recorded in Aoki (1975).

To isolate variation in the distance-ratio measure from potential confounders, I employ home-neighbor dyad fixed effects. Each dyad denotes a unique ordered pair of home domain and nearest-rival domain (e.g., A–B, A–C in Figure 6). Because dyad FEs nest both home-domain and neighbor-domain FEs, they absorb every unobserved domain-level feature — each domain's tax rate, coercive capacity, ruler ideology, agricultural endowment, and administrative quality — alongside pair-specific factors such as historical cross-border ties, shared transit corridors, or pair-level border-control arrangements. Identification then comes from villages that share both home and nearest-rival domains but differ in their relative distance to the two capitals — the within-dyad variation Figure 6 depicts inside each dashed circle.

In sum, the main specification is:

$$\text{Exit}_{i(v,j,k,t)} = \beta_2 \ln\left(\hat{\theta}_i\right) + X_v'\boldsymbol{\gamma}_2 + \underbrace{\varphi_{j,k}}_{\text{home-neighbor dyad FE}} + \underbrace{\omega_t}_{\text{decade FE}} + \varepsilon_i$$

daikansho seldom served as actual exit destinations (none appear as targets), and, in line with this, all replications yield attenuated results.

Here, Exit_i is a binary variable indicating whether event i is exit (1) or fight (0). The main predictor, $\hat{\theta}_i$, represents the ratio of the distance from village v to its home domain capital j relative to the distance to the nearest non-home domain capital active at the time of event i, as defined in the preceding subsection. This ratio is log-transformed to stabilize variance.[17] The vector X_v contains village-level controls: latitude, longitude, their interaction, elevation (log), ruggedness (log), distance to the nearer of Edo or Kyoto (where the shogunate's headquarters was located, log), and wet rice suitability (log).[18] The home-neighbor dyad fixed effects ($\varphi_{j,k}$) absorb all unobservables specific to each ordered pair of home and nearest-rival domains. Decade fixed effects (ω_t) account for temporal factors that might influence rebellion patterns throughout the study period.

3.4.3. Geographic Exposure Predicts Exit Share

The estimation results in Table 3 show that the distance ratio measure predicts villages' propensity to opt for exit over fight when rebelling. The coefficients remain positive across all specifications and statistically significant in the preferred controlled specifications, indicating that villagers strategically select their mode of resistance based on their relative distances to home and outside authorities. In the preferred specification with home-neighbor dyad fixed effects and full controls (Model 5), a one-standard-deviation increase in the log distance ratio (≈ 1.58) raises the exit probability by 3.9 percentage points, or about 16% of the 24.6% baseline exit rate. The less restrictive two-way (home + neighbor) FE specification (Model 4) yields a slightly smaller but statistically indistinguishable estimate of 3.4 percentage points, indicating that the headline result does not hinge on the dyadic restriction.

Decomposing the ratio into its two absolute distances reveals that proximity to a *foreign* capital is the active margin on two counts. First, the foreign-capital coefficient exceeds the home-capital coefficient in absolute magnitude: in the preferred dyadic speci-

[17] Using the logged ratio of distances requires caution, as it effectively constrains the coefficients of the two logged distance variables to be equal in magnitude but opposite in sign. Reassuringly, linear hypothesis tests confirm that this restriction is consistent with the data, showing no evidence that the coefficients for the two distances differ statistically from one another, as reported in Appendix D.3.5.

[18] Wet rice suitability is measured as potential output under irrigation and medium input, based on the Caloric Suitability Index developed by Galor and Özak (2016).

Outcome:	Exit$_i$						
	Distance Ratio					*Abs. Distances*	
Models:	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Variables							
Distance ratio ($\ln(\hat{\theta}_i)$)	0.0458	0.0519	0.0201	0.0212	0.0244		
Cluster SE	(0.0154)***	(0.0141)***	(0.0071)***	(0.0072)***	(0.0068)***		
Conley SE (10km)	(0.0117)***	(0.0093)***	(0.0078)***	(0.0078)***	(0.0085)***		
$\ln(d_i^{\mathrm{HC}})$						0.0160	0.0203
Cluster SE						(0.0104)	(0.0083)**
Conley SE (10km)						(0.0106)	(0.0113)*
$\ln(d_i^{\mathrm{FC}})$						−0.0306	−0.0309
Cluster SE						(0.0129)**	(0.0135)**
Conley SE (10km)						(0.0138)**	(0.0148)**
Controls							
Decade FE	No	Yes	Yes	Yes	Yes	Yes	Yes
Home Domain FE	No	No	Yes	Yes	No	Yes	No
Neighbor Domain FE	No	No	Yes	Yes	No	Yes	No
Home-Nbr Dyad FE	No	No	No	No	Yes	No	Yes
Geography	No	No	No	Yes	Yes	Yes	Yes
Fit Statistics							
Num.Obs.	2,739	2,739	2,739	2,739	2,739	2,739	2,739
R2	0.028	0.225	0.621	0.626	0.678	0.627	0.678
Mean Dep. Var.	0.246	0.246	0.246	0.246	0.246	0.246	0.246

Table 3: Exposure to Alternative Authority Predicts Exit (Exit versus Fight)

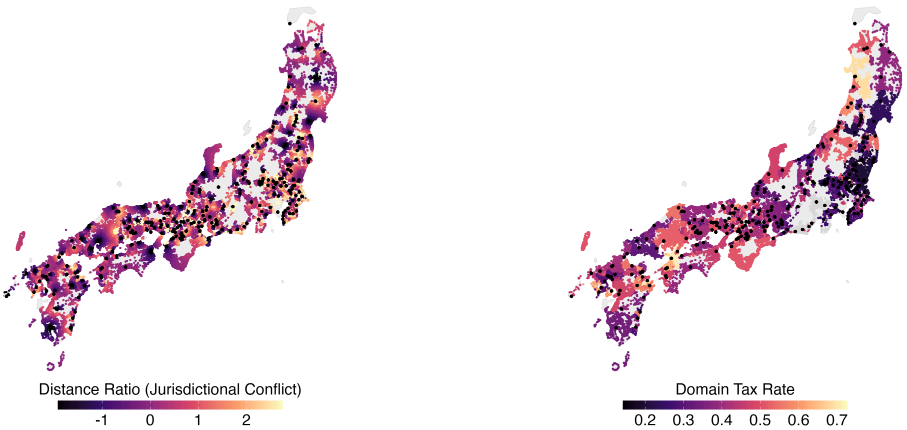
Note: OLS estimates of each village's exit-vs-fight choice conditional on revolting. Models (1)–(5) use the log distance ratio; Models (6)–(7) replace it with the two absolute distances (d_i^{HC}: home capital; d_i^{FC}: nearest foreign capital). Full results in Table A.22. SEs in parentheses: (i) clustered on home, nearest-neighbor, decade; (ii) Conley (10 km) on village coordinates; stars reflect that estimator's p-value. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

fication (Model 7), a one-standard-deviation decrease in $\ln(d_i^{\mathrm{FC}})$ (≈ 0.92) raises the exit probability by 2.8 percentage points, or about 11% of the 24.6% baseline exit rate, with a near-identical magnitude under two-way FE (Model 6). Second, in Model 7, $\ln(d_i^{\mathrm{HC}})$ loses statistical significance once Conley spatial standard errors are imposed, whereas $\ln(d_i^{\mathrm{FC}})$ remains statistically significant under both cluster and Conley standard errors in every specification. Access to a rival center of power, not limited oversight by the home domain in isolation, drives villages toward exit.

The result is robust across alternative specifications, controls, and measurement strategies. The coefficients hold under alternative controls for agricultural productivity and distance to political centers (Appendix D.2.1) and after excluding repeat-rebel villages (Appendix D.2.2). Using least-cost path distances that account for terrain ruggedness produces results consistent with the main analysis, though less robust across specifications (Appendix D.2.4); fleeing villagers likely took circuitous routes to avoid detection rather than optimal paths, weakening the relevance of the theoretically "least-costly" distance. A related concern is that straight-line distance can cut across sea straits, implying walking routes that are physically impossible; dropping village-neighbor dyads whose anchors lie on different islands leaves the coefficient unchanged (Appendix D.2.5). Controlling for each village's minimum distance to the nearest domain border (Appendix D.3.2), the distance ratio continues to predict exit while border distance itself is not statistically associated with exit choice, confirming that access to a rival center of power, not simply location near a boundary, drives the result. Restricting exit to desertion alone preserves the result (Appendix D.2.9). Including shogunal intendant offices (*daikansho*) in the distance computation leaves the preferred dyadic-FE estimate virtually unchanged (Appendix D.2.14). Finally, because Table 3 conditions on revolt, geographic exposure could in principle shape selection into the rebel set itself; a nested-logit decomposition that merges the rebel sample with the comprehensive end-period village ledger and estimates the revolt-versus-comply margin separately from the exit-versus-fight contrast finds a clean null at the selection step and recovers the Table 3 Model 5 coefficient at the form-of-resistance step, so the conditioning here is not a source of selection bias on $\hat{\theta}_i$ (Appendix D.3.3).

3.5. Exposure and Constrained Taxation

The preceding analyses establish that exposure to governance conflict channels resistance toward exit. The model predicts a further consequence: where exit is viable, rulers lower taxes to retain their population (Proposition 3). Testing this prediction directly is difficult because systematic tax data survive only as a late-period cross-section; a pooled regression of tax rates on geographic exposure is therefore associational and on its own indistinguishable from mechanisms such as state-capacity decay. I instead exploit variation in neighbor characteristics that shifts the intensity of governance conflict within the cross-section. The exit mechanism predicts that the negative association between geographic exposure and tax rates should be concentrated in domains whose neighbors are attractive enough to credibly absorb migrants; state-capacity decay implies no such asymmetry. To operationalize the geographic-exposure component, I calculate the average



(a) Village Log Distance Ratio (Governance Conflict) (b) Domain Tax Rates

Figure 7: Geographic Exposure and Tax Rates Across Villages

Note: Panel (a) plots village locations shaded by the log distance ratio $\ln(\hat{\theta}_i) = \ln(d_i^{\text{HC}}/d_i^{\text{FC}})$, the village-level measure of geographic exposure to alternative authority. Panel (b) plots the same village locations shaded by their domain's nominal tax rate; gray points denote villages in domains for which tax data are unavailable. Black dots mark domain capitals.

distance ratio across all villages within each domain. Figure 7 previews the two key variables: villages shaded by their log distance ratio (panel a) and by their domain's nominal tax rate (panel b).

3.5.1. Empirical Strategy

Estimating this association raises two identification concerns. The first is endogeneity of capital location: domain rulers may have strategically located their capitals or exchanged border villages to reduce the risk of mass exit. Capital locations were largely inherited from the Sengoku period (Hall 1955) and frozen by the 1615 one-castle-per-domain edict (*ikkoku ichijō rei*), with subsequent relocations requiring shogunal approval and remaining rare. Deliberate anti-exit capital placement after 1615 was therefore uncommon, limiting the scope for the endogeneity channel. Shogunate-strategic placement of fudai versus tozama daimyō (the shogunate seated hereditary vassals in strategically important positions and kept tozama lords on the periphery (Ravina 1999)) is a separate source of confounding that is partially absorbed by the alliance-status control below. Even if residual adjustments occurred, they bias results conservatively: deliberate capital placement and village transfers would attenuate, not inflate, the observed negative relationship between geographic exposure and tax rates. A separate concern is that a daimyō's ability to relocate a capital or trade border villages may have depended on political ties to the shogunate, since such changes required shogunal approval. This could confound the estimates if those ties also shaped a daimyō's extractive capacity. To address this possibility, I control for each daimyō's alliance status with the shogunate.[19]

The second concern is confounding. The home-capital distance component could proxy for domain scale, state capacity at the periphery (Sng and Moriguchi 2014), or village development stage, all of which might independently affect tax rates. I address these channels with controls for domain population,[20] agricultural productivity, territory size, the *gōshi system* (where lower-ranking samurai resided in villages as local officials, raising monitoring costs), and the proportion of villages with "shinden" (new rice field) in their names, indicating recent development. Tax delivery costs are unlikely confounders because villagers delivered taxes to nearby local storage facilities, not to domain capitals.[21] Tokugawa institutional arrangements also limited local agents' authority (see Section 3.1), mitigating the state-capacity-decay channel. Any remaining confounding through domain scale or coercive capacity should attenuate the estimate, since larger domains could maintain larger armies to enforce higher rates.

The outcome is the domain-level effective tax rate, computed from the same Kodama and Kitajima (1989) domain compendium introduced above (pp. 425–431) as the ratio of *shunodaka* (actual reported tax collections in rice equivalents) to *uchidaka* (the domain's internal record of actual rice production, distinct from the publicly declared *kokudaka* assessment). The ratio therefore captures the share of realized output that the domain actually collected, which is what the model's τ_j represents. The figures are based on statistics compiled by the Meiji government in 1869–1870, when each domain reported average production and collections for the preceding five years. The use of a multi-year average mitigates short-run harvest noise, and the widespread adoption of the *jōmen-*

[19]Daimyō are classified according to their relationship with the shogunate: shinpan (relatives), fudai (hereditary vassals who fought for the shogunate at Sekigahara in 1600), and tozama (outside lords who became allies only after the battle).

[20]Population is measured at the end of the Tokugawa period, so variables that use it should be read as late-period proxies rather than time-invariant characteristics.

[21]For instance, Fukui Prefecture (1994: 335-346) documents this practice in Fukui Domain.

hō (fixed-rate method) meant that many domains operated with administratively stable rates even in the final decades of Tokugawa rule (Iinuma 1974, 20–24; Smith 1958).

The late-period measurement window deserves comment. Both *shunodaka* and *uchidaka* are drawn from Meiji-era reports covering 1865–69, so the cross-sectional pattern they document is a late-period steady state rather than a time-varying outcome. Two features of the Tokugawa fiscal system make this measurement window informative for the underlying mechanism. First, *jōmen-hō* locked nominal rates in place across decades (Smith 1958), so within-domain year-to-year variation in effective rates was limited. Second, because the mechanism operates on long-run extractive stances rather than short-run fiscal shocks, a late-period measurement is appropriate for an equilibrium comparative-statics interpretation of Proposition 3. Mid-Tokugawa tax scatters compiled by Smith (1958) are qualitatively consistent with the late-period cross-section, though comprehensive mid-period data at the domain level are unavailable. Although tax rates varied among villages within each domain, systematic data at the village level are also unavailable; domain-level averages capture the main variation relevant for the analysis.[22]

I exclude non-daimyō territories (such as *goryō* under the shogunate's direct control and *jisharyō* administered by shrines and temples) as outcome units because these territories lack comprehensive fiscal data and possessed limited autonomy in establishing tax policies (Murakami 1996). The domain-level average $\hat{\theta}_j$ is computed from the same distance-ratio measure defined in Section 3.4.1, averaged across all villages within each domain using end-period capital locations.

3.5.2. The Tax–Exposure Link Binds Where Neighbors Attract Labor

I begin with a pooled benchmark that regresses the domain tax rate on geographic exposure and controls:

$$\tau_j = \alpha_4 + \beta_4 \ln\left(\hat{\theta}_j\right) + X_j' \boldsymbol{\gamma}_4 + \varepsilon_j$$

[22]One might wonder whether public goods provision, rather than tax reduction, served as an alternative means for rulers to address peasant grievances. If so, tax rates could be an incomplete measure of oppressiveness. This is unlikely for two reasons. First, public works projects depended on corvée labor (bueki) extracted from peasant villages, which increased rather than alleviated their burdens. Second, evidence from peasant petitions and revolt records indicates that grievances centered primarily on excessive taxes and unfair assessments, not on the lack of infrastructure (Aoki 1975).

Here τ_j is the tax rate in domain j and $\hat{\theta}_j$ is the domain-average distance ratio; I log-transform the distance ratio to address skewness. The vector X_j contains three groups of controls: (1) Core: domain size (log), wet-rice suitability (log),[23] proportion of new-farmland villages, and indicators for shogunate allies and the gōshi system; (2) Geography: elevation (log), ruggedness (log), and average proportions of rainy and snowy days in a month;[24] and (3) Neighbors: weighted means of adjacent domains' characteristics (territory size, population, rice productivity, number of new farmlands, family ties with the focal domain, and shogunate alliance status), with weights based on the number of villages for which each domain's capital is the nearest neighboring capital.

Table 4 documents a negative cross-sectional association between geographic exposure and reported tax rates. A one-standard-deviation increase in the log distance ratio (SD = 1.34) corresponds to a 1.5–3.2 percentage-point decrease in the reported tax rate (Models 1–4), an association that persists with the full set of controls, in a pure spatial-lag model, and marginally in the fully controlled spatial-lag model (Anselin and Rey 2014) (Models 5–6). The direction aligns with Steele, Paik, and Tanaka (2017), who establish at the domain-period level that collective desertions, but not the broader insurrection category, are associated with lower tax rates, the aggregate correlate of the mechanism this paper investigates at the community level. The pooled estimate is robust to alternative covariate specifications (kokudaka, samurai share, per-capita productivity; Appendix E.2.1), a Diegert, Masten, and Poirier (2022) sensitivity analysis (an unobserved confounder would need to exceed half the strength of all included controls to flip the sign; Appendix E.2.2), double/debiased machine learning (Chernozhukov et al. 2018) over $p = 38$ candidate covariates ($\hat{\beta} \in [-0.012, -0.010]$, $p < 0.05$; Appendix E.2.3), and including daikansho in the distance measure (Appendix D.3.1). Because this cross-sectional estimate cannot on its own separate exit from state-capacity decay, the analysis below exploits neighbor heterogeneity to isolate the exit channel.

[23]The wet-rice measure is an agro-climatic potential-yield surface (Galor and Özak 2016) rather than an observed annual output series; I use it only as a proxy for labor-absorption capacity.

[24]Weather variables represent proportion of rainy/snowy days per month, imputed as weighted averages using daily weather data from 40 locations, with distance to domain capital as weight. Data from Historical Weather Database by Professor Minoru Yoshimura.

Outcome:	τ_j					
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
Distance ratio ($\ln(\hat{\theta}_j)$)	-0.0241	-0.0203	-0.0110	-0.0112	-0.0133	-0.0089
HC SE	(0.0053)***	(0.0058)***	(0.0055)**	(0.0055)**	(0.0047)***	(0.0049)*
Conley SE (70km)	(0.0066)***	(0.0067)***	(0.0056)**	(0.0055)**	(0.0057)**	(0.0048)*
Controls						
Core	No	Yes	Yes	Yes	No	Yes
Geography	No	No	Yes	Yes	No	Yes
Neighbors	No	No	No	Yes	No	Yes
τ_j Spatial Lag	No	No	No	No	Yes	Yes
Fit Statistics						
Num.Obs.	232	230	230	230	230	230
R2	0.074	0.141	0.370	0.389	0.309[a]	0.441[a]
Mean Dep. Var.	0.391	0.391	0.391	0.391	0.394	0.394

Table 4: Pooled Benchmark: Domain-Level Geographic Exposure and Tax Rates

Note: OLS estimates (ML in Models 5–6) of domain-average distance ratio on nominal tax rate. Full results in Table A.43. SEs in parentheses: (i) HC1 (ML SEs in 5–6); (ii) Conley (70 km) on capital coordinates, computed via two-step filtered-residual approximation (Anselin and Rey 2014, pp. 34–36); stars reflect that estimator's p-value. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1. [a] Nagelkerke pseudo-R2

To identify the exit channel, I interact geographic exposure with $\hat{y}_{-j}$, a discrete moderator for the quality of neighboring domains' outside option:

$$\tau_j = \alpha_5 + \beta_5 \left[\ln\left(\hat{\theta}_j\right) \times \hat{y}_{-j}\right] + \bar{\beta}_5 \ln\left(\hat{\theta}_j\right) + \tilde{\beta}_5 \hat{y}_{-j} + X_j' \boldsymbol{\gamma}_5 + \varepsilon_j$$

I approximate the outside-option quality y_{-j} using neighboring domains' demand for labor. Rulers confronting labor shortages adopted measures to attract workers, such as land redistribution, provision of tools and rations, and investments in irrigation and other public goods (Hyugano 1957; Miyazaki 1995). Incentives to court migrants intensified with the severity of labor scarcity. In the Kantō (eastern) region, where shortages persisted throughout the period, domains implemented peasant in-migration (recruitment) policies (iri-byakushō seisaku) that offered tax exemptions, food and seed grants, and other settlement inducements to draw cultivators from neighboring territories (Gorai 1950). I construct $\hat{y}_{-j}$ as a binary indicator that equals "Low" when the average labor-to-productivity ratio of neighboring domains is *above* its median, and "High" otherwise. [25] The average is computed by weighting each neighbor by the number of domain j's villages for which that neighbor's capital is the nearest neighboring capital, and each neighbor's labor-to-productivity ratio is its population divided by its potential wet-rice yield from Galor and Özak (2016).[26] $\hat{y}_{-j} = \text{High}$ therefore marks domains whose villages face neighbors with greater labor-absorption capacity and thus stronger economic incentives to extend favorable treatment to potential migrants. In the interaction model, average rice yield and territory size of neighbors are absorbed into $\hat{y}_{-j}$ and dropped from X_j.

Figure 8 reports the main result. Across all specifications, the negative association between geographic exposure and tax rates is concentrated in domains with high $\hat{y}_{-j}$ (neighbors with low labor-to-potential-yield ratio, and thus strong demand for in-migrants). Once geographic characteristics are controlled (Models 3 and 4), the negative association attains statistical significance only for this high-$\hat{y}_{-j}$ group; for domains facing labor-abundant neighbors, the coefficients are small and statistically insignificant. Consistent with Proposition 3, geographic exposure constrains taxation only where the outside option is attractive enough to make exit credible. This asymmetry is the signature that distinguishes exit from state-capacity decay: remoteness-driven extractive slack would not vary with neighboring rulers' labor demand, whereas the exit threat, and its fiscal concession, should bind only where neighbors can credibly absorb migrants. The pattern replicates with population density as an alternative proxy for outside-option quality (Appendix E.3.3). Reconstructing $\hat{y}_{-j}$ with a fixed neighbor set (all domains within 100 km of the focal domain, independent of the nearest-capital weighting) yields the same high/low asymmetry and significance pattern, and a Hainmueller-Mummolo-Xu kernel-smoothed marginal-effect plot on the continuous moderator shows the negative association holds across the full support of neighbor labor-absorption capacity (Appendix E.3.4).

[25]Hainmueller, Mummolo, and Xu (2019) show that continuous multiplicative-interaction models rest on a linear-interaction assumption that frequently fails in practice, and recommend replacing the continuous moderator with a discrete one precisely for settings like this, in which the moderating relationship is inherently nonlinear.

[26]I use the potential yield measure under irrigated conditions with medium inputs.

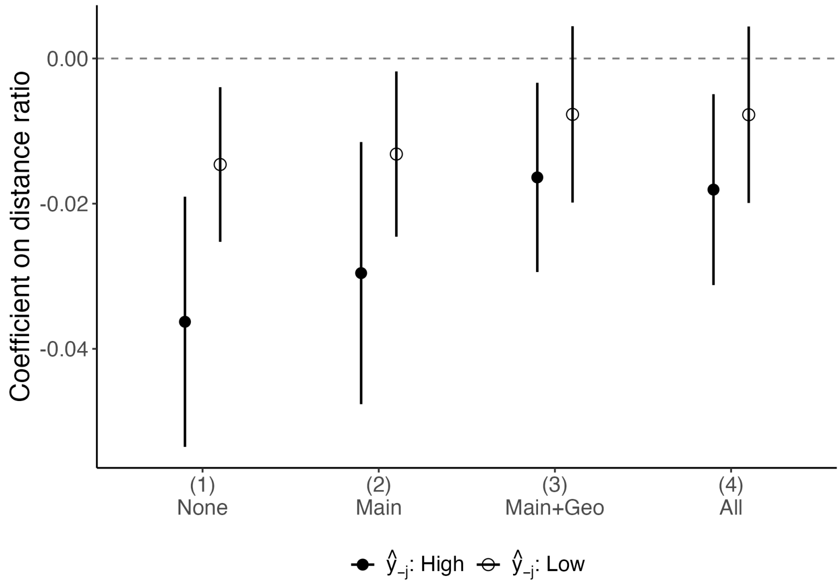


Figure 8: The Tax-Exposure Link Strengthens with Outside-Option Quality

Note: OLS coefficients on $\ln\left(\hat{\theta}_j\right) \times \hat{y}_{-j}$, where $\hat{y}_{-j}$ is above/below the median neighbor labor-to-potential-yield ratio. Filled: High; open: Low. 95% HC CIs. Controls cumulative: (1) none; (2) domain; (3) +geography; (4) +neighbor. Full table in Table A.44.

3.5.3. Additional Analyses

Two further analyses clarify what the geographic-exposure measure captures. First, simple remoteness from the home capital, the alternative predictor implied by state-capacity decay (Sng and Moriguchi 2014), does not explain the pattern: adding the average log distance to the home capital as a control leaves the geographic-exposure coefficient negative (Appendix E.2.4), and replacing the distance ratio with the remoteness measure renders any negative association insignificant once geographic controls enter (Appendix E.3.1). Second, distance to the nearest border, the alternative predictor if exit were about reaching foreign land rather than rival authority, likewise does not predict tax rates (Appendix E.3.2),[27] mirroring the village-level pattern from Section 3.4. Exposure to governance conflict, not mere remoteness or border proximity, is what constrains extraction.

[27] The boundary data were constructed following the procedure described in (Honda, Natsume, and Nemoto 2022).

4. Conclusion

This paper shows that exposure to governance conflict — the combination of geographic proximity to rival rulers and the quality of the outside option those rivals offer — determines whether subjects exit or fight. The empirical analysis proceeds in three parts. First, tenryō conversions varied the outside option, causing nearby exit to spike while fighting did not, and a grid-year panel confirms that greater geographic exposure predicts higher exit frequency. Second, among villages that did revolt, greater geographic exposure shifted the chosen form of resistance toward exit over confrontation. Third, the exit mechanism left a fiscal trace: geographic exposure predicted lower domain tax rates, and this association held only where neighboring domains offered attractive outside options, consistent with exit being credible only when both conditions are met.

Exit was not passive flight; it was a bargaining strategy that relied on shelter, mediation, and recognition from outside authorities. What enables exit as a distinct form of resistance is alternative authority, not mere border proximity or local disorder. Where rival rulers are willing and able to absorb migrants, the menu of resistance expands beyond direct confrontation, and rulers who ignore it risk losing their tax base.

The logic extends beyond Tokugawa Japan. Medieval European lordships competing for peasant tenants (Bailey 2014), precolonial African polities losing subjects to neighboring kingdoms, and modern federal systems share the same structure: fragmented authority with cross-jurisdictional allegiance shifts. In each, geographic access to a rival jurisdiction and its willingness to absorb newcomers should shape whether exit becomes viable. Whether the premise scales into the institutional outcomes those literatures emphasize lies beyond these data. The paper supplies direct community-level evidence for a premise load-bearing across the Great Divergence literature, fiscal-federalism theory, and the cartel theory of the territorial state: political fragmentation gives subjects usable leverage over rulers through competing authorities. The Tokugawa peasantry is a hard case for this premise (hegemonically constrained, administratively bound, and legally prohibited from migration), making the evidence a conservative test.

Data Availability Statement

Replication data and code will be deposited in the International Organization Dataverse (https://dataverse.harvard.edu/dataverse/IOJ) upon conditional acceptance. The full pipeline reproduces every table and figure from the raw inputs documented in the codebook.

Ethical Statement

This research uses publicly available historical records and archival data on Tokugawa-era Japan and does not involve human subjects. No IRB review was required.

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Appendix

Contents

A. Omitted Formal Analysis

The following assumption ensures a clean partition of resistance strategies: villages with high exit costs (c_H) always prefer compliance to exit, while those with low exit costs (c_L) always prefer exit to fighting.

Assumption 1: For each $j \in \{1, 2\}$:

$$\frac{k - c_L}{\delta - c_L} > \frac{y_j(1 - \theta_j)}{2\delta} \quad \text{and} \quad c_H > y_{-j} + \delta$$

The proofs of Propositions 2 and 3 additionally require the following regularity conditions, which ensure interior equilibrium play: (R1) $2\delta > y_j(1 - \theta_j)$ (positive denominators); (R2) $0 < \mu_j^{\text{fight}}, \mu_j^{\text{exit}} < 1$ (interior village cutoffs); (R3) $0 < \tau_j^* < 1$ (interior equilibrium taxes); (R4) $0 < p_j < 1$ and $0 < p_j + q_j < 1$ for each j (best-response interiority); and (R5) $P_j > 0$ for each j (the compliance intercept responds to θ_j). The non-vacuousness proposition below verifies that all conditions are jointly satisfiable.

A.1. Resistance Shares

Under Assumption 1, the resistance shares are

$$\mu_j^{\text{fight}}(\tau_j) = \frac{(1 - \theta_j)\left[\delta - k - y_j(1 - \tau_j)\right]}{2\delta - y_j(1 - \theta_j)}$$

$$\mu_j^{\text{exit}}(\tau_j, \tau_{-j}) = \frac{\theta_j}{2\delta}\left[y_{-j}(1 - \tau_{-j}) - c_L - y_j(1 - \tau_j) + \delta\right]$$

Derivation of the fight share. A high-cost village fights if $y_j\mu_j^{\text{fight}} - k \geq y_j(1 - \tau_j) + \varepsilon_{i_j}$, i.e., $\varepsilon_{i_j} \leq y_j\mu_j^{\text{fight}} - k - y_j(1 - \tau_j)$. Integrating over $\varepsilon_{i_j} \sim \text{Uniform}[-\delta, \delta]$ and multiplying by the share $1 - \theta_j$ of high-cost villages yields the fixed-point equation

$$\mu = f(\mu) := \frac{(1 - \theta_j)(y_j\mu - k - y_j(1 - \tau_j) + \delta)}{2\delta}$$

The map f is affine in μ with slope $(1 - \theta_j)y_j/(2\delta)$, which is strictly less than one under the positive-denominator condition $2\delta > y_j(1 - \theta_j)$ of Assumption 1. By the Banach contraction principle on $[0, 1 - \theta_j]$, f has a unique fixed point. When the interior formula above yields a positive value — equivalently, when $\delta > k + y_j(1 - \tau_j)$, so that a positive-measure subset of high-cost villages strictly prefers fighting even at low aggregate participation — the interior solution is that unique fixed point and is adopted throughout. Otherwise the unique fixed point is the corner $\mu_j^{\text{fight}} = 0$, which falls outside the interiority region maintained in the propositions.

A.2. Proof of Proposition 1

Proof. Fix $j \in \{1, 2\}$. Maintain Assumption 1 and assume interiority of village cutoffs. Fix tax rates $(\tau_j, \tau_{-j}) \in (0, 1)^2$.

(i) $\boldsymbol{y_{-j}}$ channel. Exit depends on y_{-j} through the exit payoff:

$$\frac{\partial \mu_j^{\text{exit}}}{\partial y_{-j}} = \theta_j \frac{1 - \tau_{-j}}{2\delta} > 0$$

Fight does not depend on y_{-j}, so $\partial \mu_j^{\text{fight}} / \partial y_{-j} = 0$.

(ii) $\boldsymbol{\tau_{-j}}$ channel. A lower neighboring tax raises the exit payoff:

$$\frac{\partial \mu_j^{\text{exit}}}{\partial \tau_{-j}} = -\theta_j \frac{y_{-j}}{2\delta} < 0$$

Fight does not depend on τ_{-j}, so $\partial \mu_j^{\text{fight}} / \partial \tau_{-j} = 0$. □

A.3. Proof of Proposition 2 (Geographic Exposure and Exit)

Proof. Maintain Assumption 1 and all interiority conditions. First note that at fixed tax rates, θ_j enters μ_j^{exit} as a multiplicative factor (so $\partial \mu_j^{\text{exit}} / \partial \theta_j = \mu_j^{\text{exit}} / \theta_j > 0$) and enters μ_j^{fight} through $1 - \theta_j$ in both numerator and denominator (giving $\partial \mu_j^{\text{fight}} / \partial \theta_j = -2\delta N_j / D_j^2 < 0$ where $D_j = 2\delta - y_j(1 - \theta_j)$ and $N_j = \delta - k - y_j(1 - \tau_j) > 0$).

The equilibrium proof proceeds by case analysis on the sign of $\frac{d\tau_j^*}{d\theta_j}$.

Case 1: $\boldsymbol{y_{-j} \le \kappa_j^*}$. Proposition 3 gives $\frac{d\tau_j^*}{d\theta_j} \ge 0$. Since the best response of domain $-j$ does not depend on θ_j, the cross-domain tax response satisfies $\frac{d\tau_{-j}^*}{d\theta_j} = q_{-j} \frac{d\tau_j^*}{d\theta_j} \ge 0$. Define $W_j := y_j - y_{-j} q_{-j}$. Since $A_{-j} \ge \theta_{-j} y_{-j} / (2\delta)$, we have $y_{-j} q_{-j} = \theta_{-j} y_j y_{-j} / (4\delta A_{-j}) \le y_j / 2$, so $W_j \ge y_j / 2 > 0$. The exit gap $E_j := y_{-j}(1 - \tau_{-j}^*) - c_L - y_j(1 - \tau_j^*) + \delta$ satisfies $\frac{dE_j}{d\theta_j} = W_j \frac{d\tau_j^*}{d\theta_j} \ge 0$. Hence the exit rate $\mu_j^{*,\text{exit}} = \theta_j E_j / (2\delta)$ has derivative

$$\frac{d\mu_j^{*,\text{exit}}}{d\theta_j} = \frac{E_j}{2\delta} + \frac{\theta_j W_j}{2\delta}\frac{d\tau_j^*}{d\theta_j} > 0$$

since the first term is strictly positive and the second is non-negative. This establishes Part (b) when $y_{-j} \leq \kappa_j^*$.

Case 2: $y_{-j} > \kappa_j^*$. Proposition 3 gives $\frac{d\tau_j^*}{d\theta_j} < 0$. Three structural observations deliver the result.

Total resistance increases. At the equilibrium, compliance satisfies $\mu_j^{\text{comply}} = A_j\tau_j^*$ (from the first-order condition). Differentiating:

$$\frac{d}{d\theta_j}\left(\mu_j^{\text{fight}} + \mu_j^{\text{exit}}\right) = -A_j'\tau_j^* - A_j\frac{d\tau_j^*}{d\theta_j} > 0$$

since $A_j' < 0$ (shown in the equilibrium analysis) and $\frac{d\tau_j^*}{d\theta_j} < 0$, making both terms positive.

Fight rate decreases. Write $\mu_j^{\text{fight}} = \left(1-\theta_j\right)N_j/D_j$ where $N_j = \delta - k - y_j\left(1-\tau_j^*\right) > 0$ (interiority) and $D_j = 2\delta - y_j\left(1-\theta_j\right) > 0$. Then

$$\frac{d\mu_j^{\text{fight}}}{d\theta_j} = \frac{-2\delta N_j + \left(1-\theta_j\right)y_j D_j \frac{d\tau_j^*}{d\theta_j}}{D_j^2} < 0$$

Both terms in the numerator are negative: the first because $N_j > 0$, the second because $\frac{d\tau_j^*}{d\theta_j} < 0$.

Part (b) follows. Since total resistance increases and fight decreases:

$$\frac{d\mu_j^{*,\text{exit}}}{d\theta_j} = \underbrace{\frac{d}{d\theta_j}\left(\mu_j^{\text{fight}} + \mu_j^{\text{exit}}\right)}_{>0} - \underbrace{\frac{d\mu_j^{\text{fight}}}{d\theta_j}}_{<0} > 0$$

Part (a) follows. By the quotient rule, $\frac{d\sigma_j^*}{d\theta_j}$ has numerator $\frac{d\mu_j^{*,\text{exit}}}{d\theta_j}\mu_j^{*,\text{fight}} - \mu_j^{*,\text{exit}}\frac{d\mu_j^{*,\text{fight}}}{d\theta_j}$. Since $\frac{d\mu_j^{*,\text{exit}}}{d\theta_j} > 0$, $\mu_j^{*,\text{fight}} > 0$, $\mu_j^{*,\text{exit}} > 0$, and $\frac{d\mu_j^{*,\text{fight}}}{d\theta_j} < 0$, every term is positive. □

A.4. Equilibrium Analysis

Define

$$A_j := \frac{(1-\theta_j)y_j}{2\delta - y_j(1-\theta_j)} + \frac{\theta_j y_j}{2\delta}$$

$$B_j := \frac{(1-\theta_j)(\delta - y_j - k)}{2\delta - y_j(1-\theta_j)} + \frac{\theta_j}{2\delta}\left[y_{-j} - c_L - y_j + \delta\right]$$

$$C_j := \frac{\theta_j y_{-j}}{2\delta}$$

so that the total resistance share is $\mu_j^{\text{fight}} + \mu_j^{\text{exit}} = B_j + A_j\tau_j - C_j\tau_{-j}$ and ruler j's utility becomes

$$u_j(\tau_j, \tau_{-j}) = y_j\tau_j\left[1 - B_j + C_j\tau_{-j} - A_j\tau_j\right]$$

Since $\frac{\partial^2 u_j}{\partial \tau_j^2} = -2y_jA_j < 0$, utility is strictly concave in own tax and the best response is unique. The first-order condition gives

$$\tau_j = p_j + q_j\tau_{-j}$$

with

$$p_j = \frac{1-B_j}{2A_j}, \quad q_j = \frac{C_j}{2A_j} = \frac{\theta_j y_{-j}}{4\delta A_j}$$

Impose interiority: $0 < p_\ell < 1$ and $0 < p_\ell + q_\ell < 1$ for each $\ell \in \{j, -j\}$.

Since $A_j \geq y_j/(2\delta)$, we have

$$q_1q_2 = \frac{\theta_1\theta_2 y_1 y_2}{16\delta^2 A_1 A_2} \leq \frac{\theta_1\theta_2}{4} \leq \frac{1}{4} < 1$$

The two affine best responses therefore intersect exactly once, yielding a unique Nash equilibrium in $(0,1)^2$:

$$\tau_j^* = \frac{p_j + q_j p_{-j}}{1 - q_j q_{-j}} \quad \text{for each } j \in \{1, 2\}$$

A.5. Proof of Proposition 3 (Tax Rate)

Proof. Fix $j \in \{1, 2\}$. Maintain Assumption 1 and assume all interiority conditions hold. Let $\Omega_j := y_{-j}$. The comparative static proceeds as in the standard two-domain tax-competition model. Since A_j does not depend on Ω_j, while B_j is affine in Ω_j, the equilibrium tax τ_j^* is a smooth function of Ω_j whose sign is determined by the numerator

$$\Xi_j(\Omega_j) := (1 - q_j q_{-j})\frac{\partial p_j}{\partial \theta_j} + (p_{-j} + p_j q_{-j})\frac{\partial q_j}{\partial \theta_j}$$

Define $\tilde{B}_j$ as the Ω_j-free part of B_j:

$$\tilde{B}_j := \frac{(1-\theta_j)(\delta - y_j - k)}{2\delta - y_j(1-\theta_j)} + \left(\frac{\theta_j}{2\delta}\right)(\delta - c_L - y_j)$$

so $B_j = \tilde{B}_j + \theta_j \Omega_j/(2\delta)$ and $\tilde{p}_j := (1 - \tilde{B}_j)/(2A_j)$ gives $p_j = \tilde{p}_j - q_j$.

Since

$$A_j' := \frac{\partial A_j}{\partial \theta_j} = \frac{y_j}{2\delta} - \frac{2\delta y_j}{\left[2\delta - y_j(1-\theta_j)\right]^2} \leq 0$$

we have $\frac{\partial q_j}{\partial \theta_j} = \Omega_j K_j > 0$ where $K_j := (A_j - \theta_j A_j')/(4\delta A_j^2) > 0$.

Defining $P_j := \partial \tilde{p}_j/\partial \theta_j$ and $R_j := \theta_j q_{-j}/(4\delta A_j) \geq 0$, the numerator becomes

$$\Xi_j(\Omega_j) = (1 - R_j \Omega_j)P_j + \Omega_j K_j \left[p_{-j} + q_{-j}\tilde{p}_j - 1\right]$$

This is affine in Ω_j with a strictly negative slope (since $p_{-j} + q_{-j}\tilde{p}_j < 1$ by interiority and $P_j > 0$). Since $\Xi_j(0) = P_j > 0$, there exists a unique threshold

$$\kappa_j^* := \frac{P_j}{\left[R_j P_j - K_j(p_{-j} + q_{-j}\tilde{p}_j - 1)\right]} > 0$$

at which $\Xi_j(\kappa_j^*) = 0$. For $\Omega_j > \kappa_j^*$, $\Xi_j < 0$ and $\frac{\partial \tau_j^*}{\partial \theta_j} < 0$. Conversely, for $\Omega_j < \kappa_j^*$, $\Xi_j > 0$ and $\frac{\partial \tau_j^*}{\partial \theta_j} > 0$; at $\Omega_j = \kappa_j^*$, the derivative is zero. This establishes the "if and only if" claim in Proposition 3. □

A.6. Non-Vacuousness

PROPOSITION 4 (Non-Vacuousness): The parameter space satisfying Assumption 1, the positive-denominator condition $2\delta > y_j(1-\theta_j)$, interiority of village cutoffs ($0 < \mu_j^{\text{fight}}, \mu_j^{\text{exit}} < 1$ and $0 < \tau_j^* < 1$), the best-response interiority conditions $0 < p_j < 1$ and $0 < p_j + q_j < 1$ for each $j \in \{1,2\}$, $P_j > 0$ for each j, and $y_{-j} > \kappa_j^*$, is non-empty. Moreover, Propositions 1–3 all hold at the exhibited parameter values.

Proof. Set $\delta = 10$, $y_1 = y_2 = 18$, $k = 7/2$, $c_L = 1/2$, $c_H = 29$, and $\theta_1 = \theta_2 = 5/6$. All basic restrictions hold by inspection. By symmetry it suffices to verify the remaining conditions for one domain.

Positive denominators. $2\delta = 20 > 3 = y_j(1 - \theta_j)$.

Assumption 1. $(k - c_L)/(\delta - c_L) = 6/19 > 3/20 = y_j(1 - \theta_j)/(2\delta)$ and $c_H = 29 > 28 = y_{-j} + \delta$.

Best-response interiority. Using the formulas above, $A_j = 63/68$, $B_j = 77/272$, $C_j = 3/4$, whence $p_j = 65/168$ and $q_j = 17/42$. Thus $0 < 65/168 < 1$ and $0 < 19/24 = p_j + q_j < 1$.

Equilibrium and village cutoffs. The equilibrium tax is $\tau_j^* = p_j/(1 - q_j) = 13/20 \in (0, 1)$. At this tax rate, $\mu_j^{\text{fight}} = 1/510$ and $\mu_j^{\text{exit}} = 19/48$, both in $(0, 1)$, with $\mu_j^{\text{fight}} + \mu_j^{\text{exit}} = 541/1360 < 1$.

$P_j > 0$. The quotient rule applied to $\tilde{p}_j = (1 - \tilde{B}_j)/(2A_j)$ gives $P_j = 2/21 > 0$.

Proposition 1 (resistance composition). Both $\mu_j^{\text{fight}} = 1/510 > 0$ and $\mu_j^{\text{exit}} = 19/48 > 0$, so interiority holds and the derivatives in Proposition 1 have the claimed signs. The y_{-j} and τ_{-j} channels follow directly from the formulas.

Proposition 2 (geographic exposure and exit). Since $\Omega_j = y_{-j} = 18 > \kappa_j^* \approx 8.49$, Case 2 of the proof applies. $A_j' < 0$ holds by inspection ($D_j = 2\delta - y_j(1 - \theta_j) = 17 < 2\delta = 20$). Interiority gives $N_j = \delta - k - y_j(1 - \tau_j^*) = 1/5 > 0$ and $\mu_j^{\text{fight}} = 1/510 > 0$. The three structural inequalities (total resistance increasing, fight decreasing, exit increasing) follow, establishing both parts (a) and (b).

Proposition 3 (tax rate). Since $\Omega_j = y_{-j} = 18$ and $\kappa_j^* \approx 8.49$, we have $\Omega_j > \kappa_j^*$, so $\frac{\partial \tau_j^*}{\partial \theta_j} < 0$.

Every condition listed in the proposition holds as a strict inequality at the exhibited parameter vector, so by continuity each condition holds on an open neighborhood of $(\delta, y_1, y_2, k, c_L, c_H, \theta_1, \theta_2) = (10, 18, 18, 7/2, 1/2, 29, 5/6, 5/6)$. The non-vacuous parameter region therefore has positive Lebesgue measure rather than being a knife-edge point. □

A.7. Extension: Attainder-Constrained Taxation

The baseline model treats the ruler as a pure revenue maximizer. Tokugawa daimyō, however, faced a distinct constraint: the shogunate's authority to strip a domain through attainder (*kaieki*) when governance failed publicly (see Section 3.2). The Buke Shohatto required domains to be "well-governed" (Steele, Paik, and Tanaka 2017), and mass revolt — especially cross-border exit that advertised governance failure directly to rival lords and to Edo — was a recognized pretext. The Nobeoka case (1690) illustrates: after exit-driven appeals reached Edo, the shogunate stripped the Arima lord of his domain. This subsection formalizes the attainder constraint as a hard cap $\bar{\mu}_j$ on the aggregate revolt share the daimyō could tolerate without triggering kaieki. When the cap binds, the ruler's optimization is constrained rather than Laffer-determined, and the comparative statics of Proposition 3 operate through a different channel: the ruler absorbs rising autonomous resistance by cutting taxes to keep total revolt at $\bar{\mu}_j$, rather than by balancing the Laffer envelope.

A.7.1. Modified Ruler Problem

Replace the ruler's unconstrained objective with

$$\max_{\tau_j \in [0,1]} \quad y_j \tau_j \left[1 - \mu_j^{\text{fight}}(\tau_j) - \mu_j^{\text{exit}}(\tau_j, \tau_{-j})\right] \quad \text{s.t.} \quad \mu_j^{\text{fight}}(\tau_j) + \mu_j^{\text{exit}}(\tau_j, \tau_{-j}) \leq \bar{\mu}_j$$

where $\bar{\mu}_j \in (0,1)$ is an exogenous primitive measuring the attainder tolerance of the shogunate toward domain j. All other primitives, including Assumption 1 and interior village cutoffs, are unchanged. The village-side analysis is therefore identical to the baseline: the resistance-share derivations deliver the same $\mu_j^{\text{fight}}(\tau_j)$ and $\mu_j^{\text{exit}}(\tau_j, \tau_{-j})$ ().

Using the baseline A_j, B_j, C_j coefficients (), the total resistance share at the best-response pair is affine:

$$\mu_j^{\text{fight}} + \mu_j^{\text{exit}} = B_j + A_j \tau_j - C_j \tau_{-j}$$

A.7.2. Regime Characterization

The Karush–Kuhn–Tucker conditions yield two regimes.

Unconstrained regime. When the constraint is slack ($\mu_j^{\text{fight}} + \mu_j^{\text{exit}} < \bar{\mu}_j$), the multiplier is zero and the ruler's best response coincides with the baseline first-order condition (): $\tau_j^* = p_j + q_j \tau_{-j}^*$, yielding the baseline equilibrium tax .

Binding regime. When the constraint binds, total revolt equals the cap and the ruler's tax is pinned by the constraint itself:

$$B_j + A_j \tau_j^* - C_j \tau_{-j}^* = \bar{\mu}_j \iff \tau_j^* = \frac{\bar{\mu}_j - B_j + C_j \tau_{-j}^*}{A_j}$$

In symmetric binding with $\bar{\mu}_1 = \bar{\mu}_2 = \bar{\mu}$ and identical primitives across j:

$$\tau_j^* = \frac{\bar{\mu} - B_j}{A_j - C_j}$$

The denominator is strictly positive: $A_j - C_j = (1 - \theta_j) y_j / (2\delta - y_j(1 - \theta_j)) > 0$, which follows automatically from the positive-denominator condition and $\theta_j \in [0, 1)$. No new regularity is imposed.

Regime boundary. At symmetric play, the unconstrained Laffer total revolt equals $(1 + B_j - C_j \tau_{-j}^*)/2$ (by substituting the Laffer FOC into the revolt expression). Hence the constraint binds iff $\bar{\mu}_j < (1 + B_j - C_j \tau_{-j}^*)/2 =: \bar{\mu}_j^\dagger$. For $\bar{\mu}_j \geq \bar{\mu}_j^\dagger$ the baseline equilibrium applies unchanged.

A.7.3. Three Propositions in the Binding Regime

PROPOSITION 5 (Revolt Constancy under Attainder): In the binding regime, the equilibrium total revolt share satisfies $\mu_j^{*,\text{fight}} + \mu_j^{*,\text{exit}} = \bar{\mu}_j$ identically, so

$$\frac{d}{d\theta_j}\left(\mu_j^{*,\text{fight}} + \mu_j^{*,\text{exit}}\right) = 0$$

Proof. Immediate from the definition of the binding regime: the constraint is saturated, and $\bar{\mu}_j$ does not depend on θ_j. □

Proposition 6 (Composition Shift under Attainder): In the binding regime under symmetric equilibrium, the equilibrium exit share $\sigma_j^* := \mu_j^{*,\text{exit}}/\bar{\mu}_j$ is strictly increasing in θ_j, and the fight share $1-\sigma_j^*$ is strictly decreasing in θ_j.

Proof. At symmetric play $\tau_j = \tau_{-j}$, the exit-share formula () collapses to

$$\mu_j^{*,\text{exit}} = \frac{\theta_j}{2\delta}\left[y_{-j}(1-\tau_{-j}^*) - c_L - y_j(1-\tau_j^*) + \delta\right] = \theta_j \frac{\delta - c_L}{2\delta}$$

because the $(1-\tau)$ terms cancel under $y_j = y_{-j}$ and $\tau_j^* = \tau_{-j}^*$. Hence $\mu_j^{*,\text{exit}}$ is a linear function of θ_j with positive slope $(\delta - c_L)/(2\delta) > 0$ (since $\delta > c_L$ by interiority). Dividing by the θ_j-invariant cap $\bar{\mu}_j$ preserves strict monotonicity:

$$\frac{d\sigma_j^*}{d\theta_j} = \frac{\delta - c_L}{2\delta\bar{\mu}_j} > 0$$

The fight share is $1-\sigma_j^*$, with opposite sign. □

Proposition 7 (Fiscal Response under Attainder): In the binding regime under symmetric equilibrium, τ_j^* is given by Equation. Let $\tau_j^\dagger := -B_j'(\theta_j)/[A_j'(\theta_j) - C_j'(\theta_j)]$ whenever $A_j'(\theta_j) - C_j'(\theta_j) < 0$. Then

$$\frac{d\tau_j^*}{d\theta_j} < 0 \quad \Longleftrightarrow \quad \tau_j^* < \tau_j^\dagger$$

At the non-vacuousness parameter vector exhibited below, $A_j' - C_j' < 0$, $\tau_j^\dagger > \tau_j^*$, and therefore $d\tau_j^*/d\theta_j < 0$.

Proof. Differentiate with respect to θ_j:

$$\frac{d\tau_j^*}{d\theta_j} = \frac{-B_j'(A_j - C_j) - (\bar{\mu}_j - B_j)(A_j' - C_j')}{(A_j - C_j)^2}$$

Using $\bar{\mu}_j - B_j = \tau_j^*(A_j - C_j)$ from ,

$$\frac{d\tau_j^*}{d\theta_j} = \frac{-B_j' - \tau_j^*(A_j' - C_j')}{A_j - C_j}$$

Since $A_j - C_j > 0$, the sign is $\operatorname{sign}\big(-B_j' - \tau_j^*\big(A_j' - C_j'\big)\big)$. Differentiating the definitions of A_j, B_j, C_j in (with $y_j = y_{-j} = y$) gives

$$A_j' = \frac{y}{2\delta} - \frac{2\delta y}{\big[2\delta - y\big(1-\theta_j\big)\big]^2}, \quad C_j' = \frac{y}{2\delta}, \quad B_j' = -\frac{2\delta(\delta - y - k)}{\big[2\delta - y\big(1-\theta_j\big)\big]^2} + \frac{\delta - c_L}{2\delta}$$

At the non-vacuousness parameter vector below, $A_j' - C_j' < 0$ and $B_j' > 0$, so $-B_j' - \tau_j^*\big(A_j' - C_j'\big) < 0$ iff $\tau_j^*\big(A_j' - C_j'\big) > -B_j'$, i.e., $\tau_j^* < -B_j'/\big(A_j' - C_j'\big) = \tau_j^\dagger$ (the inequality direction flips because $A_j' - C_j' < 0$). □

The mechanism is interpretive: rising θ_j raises autonomous exit pressure (larger μ_j^{exit} for any fixed τ_j) and lowers fight pressure (smaller μ_j^{fight}). At the baseline Laffer optimum the net effect on total revolt can go either way; under the attainder constraint, however, total revolt is pinned at $\bar{\mu}_j$, and the ruler uses the only instrument she has — her own tax — to maintain the cap. At interior parameter values where the exit channel dominates, she cuts τ_j.

A.7.4. Non-Vacuousness

Proposition 8 (Non-Vacuousness of the Binding Regime): The parameter space satisfying Assumption 1, the positive-denominator condition, interior village cutoffs, $0 < \tau_j^* < 1$, $A_j > C_j$ (automatic), the binding inequality $\bar{\mu}_j < \bar{\mu}_j^\dagger$, $\mu_j^{*,\text{fight}} > 0$ and $\mu_j^{*,\text{exit}} > 0$, and the sign conditions of Propositions Proposition 3, Proposition 4, and Proposition 5 is non-empty. Moreover, those three propositions hold at the exhibited parameter values.

Proof. Set $\delta = 10$, $y_1 = y_2 = 18$, $k = 5/2$, $c_L = 1/2$, $c_H = 29$, $\theta_1 = \theta_2 = 5/6$, and $\bar{\mu}_1 = \bar{\mu}_2 = 101/250$. The only deviation from the baseline non-vacuousness vector in Proposition 2 is $k = 5/2$ instead of $k = 7/2$; the other primitives are unchanged. The lower k widens the interval of admissible $\bar{\mu}$ values from a knife-edge band near the original NV Laffer revolt (about $(0.3958, 0.3978)$) to a window of roughly 0.01, which is required for interior binding ($\mu_j^{*,\text{fight}} > 0$) to hold on an open neighborhood rather than at a limit.

Assumption 1. $(k - c_L)/(\delta - c_L) = (2)/(19/2) = 4/19 \approx 0.211$ exceeds $y_j\big(1 - \theta_j\big)/(2\delta) = 3/20 = 0.15$. And $c_H = 29 > 28 = y_{-j} + \delta$. Both parts hold strictly.

Positive denominator. $2\delta = 20 > 3 = y_j(1 - \theta_j)$.

Baseline coefficients. $A_j = 63/68$, $B_j = 239/816$, $C_j = 3/4$. Thus $A_j - C_j = 3/17 > 0$ (so $A_j > C_j$ is automatic).

Unconstrained Laffer revolt. With $p_j = \frac{1-B_j}{2A_j} = 577/1512$ and $q_j = \frac{C_j}{2A_j} = 17/42$, the symmetric baseline tax is $\tau_j^{*,\text{Laffer}} = \frac{p_j}{1-q_j} = 577/900 \approx 0.641$, yielding total revolt $\frac{1+B_j - C_j \tau_j^{*,\text{Laffer}}}{2} = 2761/6800 \approx 0.406$.

Binding. Since $\bar{\mu} = 101/250 = 0.404 < 2761/6800 \approx 0.406 = \bar{\mu}_j^{\dagger}$, the constraint binds.

Binding equilibrium tax. From , $\tau_j^* = \frac{101/250 - 239/816}{3/17} = 11333/18000 \approx 0.630 \in (0, 1)$.

Resistance shares at binding. $\mu_j^{*,\text{exit}} = \theta_j \frac{\delta - c_L}{2\delta} = 19/48 \approx 0.3958 > 0$ (independent of τ_j at symmetric play). $\mu_j^{*,\text{fight}} = \bar{\mu} - \mu_j^{*,\text{exit}} = 101/250 - 19/48 = 49/6000 \approx 0.00817 > 0$. Both are strictly interior and sum to $\bar{\mu} = 101/250$.

Sign of $A_j' - C_j'$ and B_j'. $A_j' - C_j' = -2\delta y / \left[2\delta - y(1 - \theta_j)\right]^2 = -360/289 < 0$. $B_j' = -2\delta \frac{\delta - y - k}{[2\delta - y(1-\theta_j)]^2} + \frac{\delta - c_L}{2\delta} = (20)\frac{10.5}{289} + 19/40 = 210/289 + 19/40 = 13891/11560 \approx 1.20 > 0$.

Fiscal threshold. $\tau_j^{\dagger} = -B_j' / (A_j' - C_j') = (13891/11560)(289/360) = 13891/14400 \approx 0.965$. Since $\tau_j^* = 11333/18000 \approx 0.630 < 0.965 = \tau_j^{\dagger}$, Proposition 5 delivers $d\tau_j^*/d\theta_j < 0$. Direct computation: $d\tau_j^*/d\theta_j = -473/200 = -2.365$.

Composition shift. $d\sigma_j^*/d\theta_j = (\delta - c_L)/(2\delta\bar{\mu}) = (19/40)/(101/250) = 475/404 \approx 1.176 > 0$.

Revolt constancy. $d\left(\mu_j^{*,\text{fight}} + \mu_j^{*,\text{exit}}\right)/d\theta_j = 0$ by definition of the binding regime. Direct computation: $d\mu_j^{*,\text{exit}}/d\theta_j = \frac{\delta - c_L}{2\delta} = 19/40 = +0.475$ and $d\mu_j^{*,\text{fight}}/d\theta_j = -19/40 = -0.475$, summing to zero.

All listed conditions hold as strict inequalities at the exhibited parameter vector, so by continuity each holds on an open neighborhood, giving the non-vacuous region positive Lebesgue measure. □

A.7.5. Comparison with Baseline Predictions

The attainder extension delivers one distinctive prediction and preserves three others. The contrast with the baseline is the point: at the baseline Laffer optimum, total revolt *rises* with θ_j whenever the outside option is credible (Case 2 of the Proposition 2 proof). Under the binding attainder constraint, total revolt is mechanically pinned at $\bar{\mu}_j$ and therefore does *not* respond to θ_j.

New: total revolt is invariant to θ_j. Proposition 3: $d\left(\mu_j^{*,\text{fight}} + \mu_j^{*,\text{exit}}\right)/d\theta_j = 0$ identically. This zero-response is the empirical signature that distinguishes the attainder regime from the baseline.

Preserved: composition. Proposition 2 (a) holds in the binding regime via Proposition 4: $\sigma_j^* = \theta_j(\delta - c_L)/\left(2\delta\bar{\mu}_j\right)$ is strictly increasing in θ_j. Mechanism: the composition shift now operates through fight attrition (the fight rate *falls* by exactly the amount the exit rate rises) rather than through overall revolt expansion.

Preserved: exit rate. Proposition 2 (b) holds: $\mu_j^{*,\text{exit}} = \theta_j(\delta - c_L)/(2\delta)$ at symmetric play, strictly increasing in θ_j.

Preserved: tax direction, different channel. Proposition 3 holds via Proposition 5 at interior parameters. The baseline cuts τ_j because exit-constraint tightening dominates fight-pool shrinkage (the κ_j^* threshold). The attainder regime cuts τ_j because the ruler absorbs rising exit pressure to hold total revolt at the cap. Same sign of $d\tau_j^*/d\theta_j$, distinct mechanisms.

Unchanged: partial-equilibrium composition. Proposition 1, stated at fixed tax rates, depends only on village-side optimization (proof of Proposition 1) and is unaffected by the ruler's objective.

A.7.6. Connection to the Empirical Results

The binding regime delivers a sharper prediction for the Step-1 result of Appendix D.3.4 than the baseline Laffer model. Under the baseline, greater geographic exposure raises total revolt (Case 2 of the Proposition 2 proof) through rising autonomous exit pressure; a small negative Step-1 coefficient is then an empirically observed but theoretically underdetermined magnitude. Under the attainder extension, the ruler's tax response absorbs exactly the rise in resistance pressure that would otherwise increase total revolt, pinning $\mu_j^{*,\text{fight}} + \mu_j^{*,\text{exit}} = \bar{\mu}_j$; the small-magnitude empirical coefficient is then the natural prediction rather than a quantitative surprise. The extension is consistent with the paper's main-text description of attainder as "what the warrior class feared most" (Section 3.2) and with the Nobeoka, Yamashirodani, and Matsuyama cases, in which exit-driven revolts triggered kaieki or near-kaieki outcomes. The baseline Laffer model remains relevant for rulers facing loose shogunal scrutiny and low baseline revolt; the attainder extension captures rulers whose governance visibility places them close to the kaieki margin.

A.8. Extension: Stochastic Kaieki Risk

The hard-cap extension of Appendix A.7 is the sharpest specification of attainder discipline. A smoother alternative lets the ruler trade expected tax revenue against an expected-attainder penalty that grows with revolt above the shogunate's tolerance. This subsection embeds the baseline, the hard-cap extension, and all intermediate cases within a single two-parameter family: a tolerance threshold $\bar{\mu}_j \in (0, 1)$ (the same primitive as in Appendix A.7) and a risk-intensity parameter $\Lambda \geq 0$. As $\Lambda \to \infty$ the family collapses onto the hard-cap extension; at $\Lambda = 0$ it returns the baseline. Between the two extremes, the ruler's optimization strengthens continuously with Λ, and equilibrium revolt's sensitivity to geographic exposure θ_j shrinks monotonically to zero.

A.8.1. Modified Ruler Problem

Replace the ruler objective with

$$u_j^{\text{stoch}}(\tau_j, \tau_{-j}; \Lambda, \bar{\mu}_j) = y_j \tau_j \left[1 - M_j\right] - \left(\frac{\Lambda}{2}\right) \left[\max(0, M_j - \bar{\mu}_j)\right]^2$$

where $M_j := \mu_j^{\text{fight}}(\tau_j) + \mu_j^{\text{exit}}(\tau_j, \tau_{-j})$. The quadratic one-sided penalty activates only when revolt exceeds the shogunate's tolerance $\bar{\mu}_j$; $\Lambda \geq 0$ scales the expected-attainder cost (the product of the probability-per-unit excess revolt and the forfeited-domain value). The village-side analysis () is unchanged, so $M_j = B_j + A_j \tau_j - C_j \tau_{-j}$ using the baseline coefficients .

A.8.2. Two Regimes

The FOC bifurcates at the threshold.

Slack regime ($M_j^* \le \bar{\mu}_j$). The penalty derivative vanishes, the baseline FOC applies, and τ_j^*, M_j^* coincide with the baseline equilibrium . The slack regime is active iff the baseline revolt $M_j^{*,\text{base}} := \left[A_j\left(1+B_j\right)-C_j\right] / \left(2A_j - C_j\right)$ at symmetric play satisfies $M_j^{*,\text{base}} \le \bar{\mu}_j$; i.e., when the shogunate's tolerance exceeds the ruler's unconstrained preferred revolt.

Binding regime ($M_j^* > \bar{\mu}_j$). Strict concavity in own tax is preserved: $\partial^2 u_j^{\text{stoch}} / \partial \tau_j^2 = -2y_j A_j - \Lambda A_j^2 < 0$. The FOC is

$$y_j\left(1-M_j\right) = A_j\left[y_j \tau_j + \Lambda\left(M_j - \bar{\mu}_j\right)\right]$$

yielding the affine best response $\tau_j = p_j\left(\Lambda, \bar{\mu}_j\right) + q_j(\Lambda)\tau_{-j}$ with

$$p_j\left(\Lambda, \bar{\mu}_j\right) = \frac{y_j\left(1-B_j\right) + \Lambda A_j\left(\bar{\mu}_j - B_j\right)}{A_j\left(2y_j + \Lambda A_j\right)}, \quad q_j(\Lambda) = \frac{C_j\left(y_j + \Lambda A_j\right)}{A_j\left(2y_j + \Lambda A_j\right)}$$

At $\Lambda = 0$ both reduce to the baseline p_j, q_j , regardless of $\bar{\mu}_j$.

A.8.3. Symmetric Binding Equilibrium

Set $y_1 = y_2 = y$, collapse $\left(A_j, B_j, C_j\right)$ to (A, B, C), and set $\bar{\mu}_1 = \bar{\mu}_2 = \bar{\mu}$. The symmetric Nash equilibrium in the binding regime is

$$\tau_j^*(\Lambda, \bar{\mu}) = \frac{y(1-B) + \Lambda A(\bar{\mu} - B)}{y(2A-C) + \Lambda A(A-C)}$$

and, substituting into $M_j = B + (A-C)\tau_j$,

$$M_j^*(\Lambda, \bar{\mu}) = \frac{y[A(1+B)-C] + \Lambda A(A-C)\bar{\mu}}{y(2A-C) + \Lambda A(A-C)}$$

nests three benchmarks:

- *Baseline ($\Lambda = 0$).* $M_j^*(0, \bar{\mu}) = [A(1+B)-C]/(2A-C) = M_j^{*,\text{base}}$, independent of $\bar{\mu}$.
- *Hard cap ($\Lambda \to \infty$).* Both numerator and denominator grow linearly in Λ through the same factor $A(A-C)$, and $\lim_{\Lambda\to\infty} M_j^*(\Lambda, \bar{\mu}) = \bar{\mu}$. The $\Lambda \to \infty$ limit is therefore the binding regime of Appendix A.7 at the same tolerance $\bar{\mu}$.
- *Zero-tolerance special case ($\bar{\mu} = 0$).* $M_j^*(\Lambda, 0) = y[A(1+B)-C]/[y(2A-C) + \Lambda A(A-C)] \to 0$ as $\Lambda \to \infty$.

A.8.4. Main Result

PROPOSITION 9 (Convergence and Vanishing Sensitivity under Stochastic Kaieki Risk): Under Assumption 1, symmetric interior play, and the binding regime ($M_j^{*,\text{base}} > \bar{\mu}$):

(a) *Convergence to the cap.* $M_j^*(\Lambda, \bar{\mu})$ is strictly decreasing in Λ, with $\lim_{\Lambda\to\infty} M_j^*(\Lambda, \bar{\mu}) = \bar{\mu}$.

(b) *Vanishing sensitivity.* $\lim_{\Lambda\to\infty} \partial M_j^*/\partial\theta_j = 0$.

Proof. Write $M_j^* = N(\theta_j; \bar{\mu}, \Lambda)/D(\theta_j; \Lambda)$ with

$$N(\theta_j; \bar{\mu}, \Lambda) = y[A(1+B) - C] + \Lambda A(A-C)\bar{\mu}, \quad D(\theta_j; \Lambda) = y(2A - C) + \Lambda A(A-C)$$

(a) Differentiating in Λ:

$$\frac{\partial M_j^*}{\partial \Lambda} = \frac{A(A-C)[\bar{\mu}D - N]}{D^2} = \frac{A(A-C)\cdot y(2A-C)\left[\bar{\mu} - M_j^{*,\text{base}}\right]}{D^2}$$

since $\bar{\mu}D - N = \bar{\mu}y(2A-C) + \Lambda A(A-C)\bar{\mu} - y[A(1+B)-C] - \Lambda A(A-C)\bar{\mu} = y(2A-C)\left[\bar{\mu} - M_j^{*,\text{base}}\right]$. In the binding regime $\bar{\mu} < M_j^{*,\text{base}}$, so $\partial M_j^*/\partial\Lambda < 0$. For the limit, write $M_j^* = [N/\Lambda]/[D/\Lambda] \to [A(A-C)\bar{\mu}]/[A(A-C)] = \bar{\mu}$ as $\Lambda \to \infty$.

(b) Differentiating in θ_j:

$$\frac{\partial M_j^*}{\partial \theta_j} = \frac{\partial N/\partial\theta_j}{D} - \frac{N \cdot \partial D/\partial\theta_j}{D^2}$$

Both D and $\partial D/\partial\theta_j$ are $O(\Lambda)$ at large Λ (through $A(A-C)$ and its θ_j-derivative), and N, $\partial N/\partial\theta_j$ are also $O(\Lambda)$ (through the $\bar{\mu}$ term). The first right-hand-side term is thus $O(1)$, the second $O(\Lambda\cdot\Lambda)/O(\Lambda^2) = O(1)$. Their difference is nevertheless $o(1)$: substituting the asymptotic forms,

$$\frac{\partial M_j^*}{\partial \theta_j} \to \frac{\bar{\mu}\cdot\partial[A(A-C)]/\partial\theta_j \cdot A(A-C) - A(A-C)\bar{\mu}\cdot\partial[A(A-C)]/\partial\theta_j}{[A(A-C)]^2} = 0$$

as $\Lambda \to \infty$ (the leading $O(1)$ terms cancel, leaving $O(1/\Lambda)$ residuals). □

A.8.5. Interpretation

The stochastic kaieki specification nests the three benchmarks cleanly: the baseline ($\Lambda = 0$), the hard-cap extension ($\Lambda \to \infty$), and a zero-tolerance limit ($\bar{\mu} = 0$). The intermediate family, indexed by $\Lambda \in (0, \infty)$, describes rulers who internalize attainder risk partially rather than in the two extremes of ignoring it or respecting it as a hard constraint.

As Λ grows, the ruler's optimization against kaieki strengthens. Part (a) shows that equilibrium total revolt is pushed monotonically toward the shogunate's tolerance $\bar{\mu}$, and part (b) shows that its sensitivity to geographic exposure θ_j vanishes in the limit: cross-domain variation in θ_j is absorbed by the ruler's anticipatory tax response rather than by differential equilibrium revolt. Both extensions (Appendix A.7 and the present one) therefore rationalize the attenuated Step-1 coefficient in Appendix D.3.4 through the same channel, the ruler's anticipatory response to attainder risk; but the stochastic version places the comparative static on a continuum that recovers the hard-cap result as the large-Λ limit.

A.9. Extension: Permanent Exit

The baseline treats exit as a temporary bargaining tactic: exiters obtain concessions and return, so the ruler's one-shot revenue problem internalizes no lost tax base beyond within-period non-compliance. A realistic alternative lets a fraction of exiters be absorbed permanently into the neighbor's domain, so each unit of equilibrium exit erodes the ruler's future tax base as well. This subsection formalizes that feature by augmenting the ruler's objective with a linear permanent-exit penalty governed by an exogenous primitive $\rho_j \geq 0$. The penalty shifts the ruler's best-response intercept downward without altering its slope, preserves strict concavity and Nash uniqueness, and strengthens the tax-base-versus-extraction trade-off already operating in the baseline. The three main comparative statics retain their signs, and the tax response to geographic exposure is *strictly intensified*, reinforcing, not replacing, the Case 2 mechanism of Proposition 2.

A.9.1. Modified Ruler Problem

Replace with

$$u_j^{\text{perm}}(\tau_j, \tau_{-j}; \rho_j) = y_j\tau_j\left[1 - \mu_j^{\text{fight}}(\tau_j) - \mu_j^{\text{exit}}(\tau_j, \tau_{-j})\right] - \rho_j\mu_j^{\text{exit}}(\tau_j, \tau_{-j})$$

where $\rho_j \geq 0$ denotes the capitalized per-villager present value of tax base permanently lost when a village exits. The baseline corresponds to $\rho_j = 0$ (all exiters return after concessions); $\rho_j > 0$ captures partial or full permanence. All other primitives, including Assumption 1 and interior village cutoffs, are unchanged. Village-side analysis is therefore identical to the baseline: and the coefficients A_j, B_j, C_j of carry over verbatim. Introduce the auxiliary coefficient

$$A_j^{\text{exit}} := \frac{\theta_j y_j}{2\delta}$$

which equals both the tax-sensitivity of μ_j^{exit} ($\partial\mu_j^{\text{exit}}/\partial\tau_j = A_j^{\text{exit}}$) and the exit-channel contribution to A_j ($A_j = A_j^{\text{fight}} + A_j^{\text{exit}}$, where $A_j^{\text{fight}} := (1-\theta_j)y_j/[2\delta - y_j(1-\theta_j)]$).

A.9.2. Equilibrium Characterization

Strict concavity in own tax is preserved: $\partial^2 u_j^{\text{perm}}/\partial\tau_j^2 = -2y_jA_j < 0$, independent of ρ_j. Differentiating u_j^{perm} in τ_j and using $\partial\mu_j^{\text{exit}}/\partial\tau_j = A_j^{\text{exit}}$, the first-order condition is

$$y_j(1 - B_j + C_j\tau_{-j}) - 2y_jA_j\tau_j - \rho_jA_j^{\text{exit}} = 0$$

yielding the affine best response

$$\tau_j = \tilde{p}_j(\theta_j, \rho_j) + q_j(\theta_j)\tau_{-j}, \quad \tilde{p}_j := p_j - \frac{\rho_jA_j^{\text{exit}}}{2y_jA_j}, \quad q_j := \frac{C_j}{2A_j}$$

whose slope q_j is identical to the baseline . Since q_j is unchanged, the contraction bound $q_1q_2 \leq 1/4 < 1$ derived in the baseline equilibrium analysis carries over, so the two affine best responses intersect exactly once. The unique symmetric Nash equilibrium is

$$\tau_j^*(\rho_j, \rho_{-j}) = \frac{\tilde{p}_j + q_j\tilde{p}_{-j}}{1 - q_jq_{-j}}$$

At $\rho_j = \rho_{-j} = 0$, $\tilde{p}_j = p_j$ and collapses to the baseline equilibrium .

A.9.3. Comparative Statics

The three baseline comparative-statics results carry over under permanent exit, and the tax-direction result is strictly intensified. I state each as a separate proposition.

Proposition 10 (Fiscal Intensification under Permanent Exit): Impose symmetric primitives $\rho_1 = \rho_2 = \rho$ and evaluate at a symmetric interior point $\theta_1 = \theta_2 = \theta$ satisfying all interiority conditions of Proposition 2, and additionally $y_{-j} > \kappa_j^*$ (so Case 2 of Proposition 3 delivers $\partial\tau_j^*/\partial\theta_j < 0$ at $\rho = 0$). Then (a) $\partial\tau_j^*/\partial\theta_j < 0$ for every $\rho \geq 0$ at which the equilibrium remains interior. (b) $\left|\partial\tau_j^*/\partial\theta_j\right|$ is strictly increasing in ρ.

Proof. At symmetric primitives the equilibrium reduces to

$$\tau^*(\theta, \rho) = \tau^*(\theta, 0) - \rho \cdot \frac{S(\theta)}{1 - q(\theta)}$$

where $S(\theta) := A^{\text{exit}}(\theta)/[2yA(\theta)] = \theta/[4\delta A(\theta)]$ (using $A^{\text{exit}} = \theta y/(2\delta)$).

Part (a). Differentiate in θ at fixed ρ:

$$\frac{\partial\tau^*}{\partial\theta}(\theta, \rho) = \frac{\partial\tau^*}{\partial\theta}(\theta, 0) - \rho\frac{d}{d\theta}\left[\frac{S(\theta)}{1 - q(\theta)}\right]$$

The first term is strictly negative by Case 2 of Proposition 3. For the second, $A'(\theta) \leq 0$ (shown at line , as $A' = y/(2\delta) - 2\delta y/[2\delta - y(1-\theta)]^2$) and $q(\theta) = \theta y/[4\delta A(\theta)]$ has $q'(\theta) > 0$ (the numerator θy is strictly increasing and the denominator $4\delta A(\theta)$ is non-increasing, since $A' \leq 0$). Similarly $S(\theta) = \theta/[4\delta A(\theta)]$ is the product of a strictly increasing function and a non-decreasing function, so $S'(\theta) > 0$. Hence $d[S/(1-q)]/d\theta > 0$: the numerator's derivative S' is positive, and $-(1-q)' = q' > 0$ adds a further positive term. The second-term contribution $-\rho \cdot d[S/(1-q)]/d\theta$ is therefore strictly negative for every $\rho > 0$, and the sum of two strictly negative terms is strictly negative. At $\rho = 0$ the statement reduces to Case 2 of Proposition 3.

Part (b). From the same decomposition,

$$\frac{\partial}{\partial\rho}\frac{\partial\tau^*}{\partial\theta} = -\frac{d}{d\theta}\left[\frac{S(\theta)}{1 - q(\theta)}\right] < 0$$

Since $\partial\tau^*/\partial\theta < 0$, $|\partial\tau^*/\partial\theta|$ is strictly increasing in ρ. □

Proposition 11 (Preserved Signs for Composition and Exit under Permanent Exit): Under Assumption 1, interiority, and Case 2 primitives ($y_{-j} > \kappa_j^*$), there exists $\bar{\rho}_j > 0$ such that for all $\rho_j \in [0, \bar{\rho}_j)$ the signs in Proposition 1, in Proposition 2 (a), and in Proposition 2 (b) are preserved. Proposition 1 is preserved identically for every $\rho_j \geq 0$.

Proof. Proposition 1 is a partial-equilibrium statement. The derivatives in Proposition 1 are evaluated at fixed (τ_j, τ_{-j}) and depend only on the village-side optimization. Since μ_j^{fight} and μ_j^{exit} are identical to the baseline at any fixed tax pair, the signs in Proposition 1 carry over for every $\rho_j \geq 0$ identically.

Fight rate decreases in θ_j. Write $\mu_j^{\text{fight}} = (1-\theta_j)N_j/D_j$ with $N_j := \delta - k - y_j(1-\tau_j^*) > 0$ (by interiority) and $D_j := 2\delta - y_j(1-\theta_j) > 0$. The same derivative computation as in Case 2 of the Proposition 2 proof (line) gives

$$\frac{d\mu_j^{\text{fight}}}{d\theta_j} = \frac{-2\delta N_j + (1-\theta_j)y_j D_j \frac{d\tau_j^*}{d\theta_j}}{D_j^2} < 0$$

where $d\tau_j^*/d\theta_j < 0$ is preserved under permanent exit by Proposition 8 (a). Both numerator terms are negative.

Total resistance increases in θ_j for small ρ_j. The modified first-order condition implies

$$\mu_j^{\text{comply}} = 1 - \left(B_j + A_j\tau_j^* - C_j\tau_{-j}^*\right) = A_j\tau_j^* + \rho_j A_j^{\text{exit}}/y_j$$

at symmetric play, so the total-resistance derivative in θ_j is

$$\frac{d}{d\theta_j}\left(\mu_j^{\text{fight}} + \mu_j^{\text{exit}}\right) = -A_j'\tau_j^* - A_j\frac{d\tau_j^*}{d\theta_j} - \rho_j\left(A_j^{\text{exit}}\right)'/y_j = -A_j'\tau_j^* - A_j\frac{d\tau_j^*}{d\theta_j} - \frac{\rho_j}{2\delta}$$

where the last equality uses $\left(A_j^{\text{exit}}\right)' = y_j/(2\delta)$. At $\rho_j = 0$ this equals $-A_j'\tau_j^* - A_j(d\tau_j^*/d\theta_j)$, which is *strictly positive* by Case 2 of the Proposition 2 proof (both terms positive). By continuity in ρ_j there exists $\bar{\rho}_j > 0$ such that the sum remains strictly positive for all $\rho_j \in [0, \bar{\rho}_j)$.

Exit rate and composition. Combining the two preceding steps, for $\rho_j \in [0, \bar{\rho}_j)$:

$$\frac{d\mu_j^{*,\text{exit}}}{d\theta_j} = \underbrace{\frac{d}{d\theta_j}\left(\mu_j^{\text{fight}} + \mu_j^{\text{exit}}\right)}_{>0} - \underbrace{\frac{d\mu_j^{\text{fight}}}{d\theta_j}}_{<0} > 0$$

establishing Proposition 2 (b). The composition result Proposition 2 (a) follows by the quotient-rule argument at line : the numerator of $d\sigma_j^*/d\theta_j$ equals $\left(d\mu_j^{*,\text{exit}}/d\theta_j\right)\mu_j^{*,\text{fight}} - \mu_j^{*,\text{exit}}\left(d\mu_j^{*,\text{fight}}/d\theta_j\right)$, which is positive since each factor has the sign already established and $\mu_j^{*,\text{fight}}, \mu_j^{*,\text{exit}} > 0$ at interior play. □

A.9.4. Non-Vacuousness

Proposition 12 (Non-Vacuousness of the Permanent-Exit Regime): The parameter space satisfying Assumption 1, the positive-denominator condition, interiority of village cutoffs and equilibrium taxes, the best-response interiority of Proposition 2, $y_{-j} > \kappa_j^*$, and the sign conditions of Proposition 8 and Proposition 9 is non-empty. Moreover, those propositions hold at the exhibited parameter values.

Proof. Use the baseline non-vacuousness vector of Proposition 2: $\delta = 10$, $y_1 = y_2 = 18$, $k = 7/2$, $c_L = 1/2$, $c_H = 29$, $\theta_1 = \theta_2 = 5/6$. All baseline-only conditions (Assumption 1, positive denominator, $A_j > 0$, $P_j > 0$, $\kappa_j^* \approx 8.49 < y_{-j} = 18$) hold exactly as in Proposition 2. Set the permanent-exit primitive $\rho_1 = \rho_2 = \rho = 1/10$.

Baseline quantities at $\rho = 0$. From Proposition 2: $A_j = 63/68$, $B_j = 77/272$, $C_j = 3/4$, $p_j = 65/168$, $q_j = 17/42$, $\tau_j^* = 13/20$.

Permanent-exit coefficients. $A_j^{\text{exit}} = \theta_j y_j/(2\delta) = (5/6)(18)/20 = 3/4$ (which happens to equal C_j at symmetric $y_j = y_{-j}$, since both reduce to $\theta y/(2\delta)$). The intercept shift per unit ρ is

$$\frac{S_j}{1-q_j} = \frac{A_j^{\text{exit}}}{2y_jA_j} \cdot \frac{1}{1-q_j} = \frac{3/4}{2 \cdot 18 \cdot 63/68} \cdot \frac{42}{25} = \frac{17}{756} \cdot \frac{42}{25} = \frac{17}{450}$$

so $S_j/(1-q_j) \approx 0.0378$.

Equilibrium tax at $\rho = 1/10$. $\tau_j^*(\rho) = 13/20 - (1/10)(17/450) = 13/20 - 17/4500 = 2908/4500 = 727/1125 \approx 0.6462 \in (0,1)$; interior.

Interior village cutoffs at $\rho = 1/10$. At $\tau_j^* = 727/1125$, $\mu_j^{\text{fight}} = (1-\theta_j)\left[\delta - k - y_j(1-\tau_j^*)\right]/\left[2\delta - y_j(1-\theta_j)\right] = (1/6)(33/250)/(17) = 33/25500 = 11/8500 \approx 0.00129 > 0$, and $\mu_j^{\text{exit}} = \theta_j(\delta - c_L)/(2\delta) = (5/6)(19/2)/20 = 19/48 \approx 0.396 > 0$ (independent of τ at symmetric play). Both strictly interior; total $\mu_j^{\text{fight}} + \mu_j^{\text{exit}} \approx 0.397 < 1$.

Proposition 8 (a): sign preservation. Direct differentiation of at the exhibited vector (numerical evaluation by symbolic differentiation) gives $\partial\tau_j^*/\partial\theta_j \approx -0.2244 < 0$. At $\rho = 0$ the analogous derivative is $\approx -0.2144 < 0$. Both strictly negative.

Proposition 8 (b): intensification. $d\left[S_j/(1-q_j)\right]/d\theta_j \approx 0.0998 > 0$ at the exhibited vector (direct evaluation using the closed form $S/(1-q) = \theta/[4\delta A(\theta) - \theta y]$ and $A' < 0$ at $\theta = 5/6$). Hence $\partial^2\tau_j^*/\partial\rho\partial\theta_j \approx -0.0998 < 0$, so $\left|\partial\tau_j^*/\partial\theta_j\right|$ strictly increases from ≈ 0.2144 at $\rho = 0$ to ≈ 0.2244 at $\rho = 1/10$.

Proposition 9: preserved signs. The key inequality $-A_j'\tau_j^* - A_j\left(d\tau_j^*/d\theta_j\right) - \rho_j/(2\delta) > 0$ must hold at the exhibited vector. Direct evaluation: $A_j' = y/(2\delta) - 2\delta y/[2\delta - y(1-\theta)]^2 = 9/10 - 360/289 = -999/2890 \approx -0.3457$; $\tau_j^* \approx 0.6462$; $-A_j'\tau_j^* \approx 0.2234$; $A_j\left(d\tau_j^*/d\theta_j\right) \approx (63/68)(-0.2244) \approx -0.2079$; $-\rho/(2\delta) = -1/200 = -0.005$. Sum $\approx 0.2234 + 0.2079 - 0.005 \approx 0.4263 > 0$. Continuity in ρ then gives a neighborhood $\rho \in [0, \bar{\rho})$ on which the total-resistance derivative remains strictly positive; $\rho = 1/10$ is inside this neighborhood. Fight rate strictly decreases and exit rate strictly increases by the arithmetic in the Proposition 9 proof.

All listed conditions hold as strict inequalities at the exhibited parameter vector, so by continuity each holds on an open neighborhood of $(\delta, y_1, y_2, k, c_L, c_H, \theta_1, \theta_2, \rho_1, \rho_2) = (10, 18, 18, 7/2, 1/2, 29, 5/6, 5/6, 1/10, 1/10)$, giving the non-vacuous region positive Lebesgue measure. □

A.9.5. Comparison with Baseline Predictions

The permanent-exit extension preserves every sign of the baseline and strictly intensifies one magnitude. In contrast to the two attainder extensions (Appendix A.7 and Appendix A.8), which overturn Proposition 2 (b) as a by-product of pinning total revolt to the shogunate's tolerance, permanent exit leaves the full Proposition-1-through-3 vector intact at ρ_j near zero.

Preserved and intensified: tax direction (Proposition 3). Proposition 8 (a) delivers $\partial\tau_j^*/\partial\theta_j < 0$ for every $\rho \geq 0$ at which the equilibrium remains interior, and Proposition 8 (b) delivers strict magnitude intensification: $\left|\partial\tau_j^*/\partial\theta_j\right|$ rises with ρ.

Preserved: composition and exit rate (Proposition 2). Proposition 9 delivers the signs of Proposition 2 (a) and (b) on a neighborhood of $\rho = 0$. The Case 2 mechanism (exit-constraint tightening dominating fight-pool shrinkage, so total resistance rises in θ_j) survives continuously in ρ.

Preserved: partial-equilibrium composition (Proposition 1). Unchanged identically, since Proposition 1 depends only on village-side optimization.

Mechanism: under permanent exit each marginal unit of extraction generates two losses, within-period non-compliance (as in the baseline) and permanent tax-base erosion (new). The ruler internalizes the second loss through the $\rho_j\mu_j^{\text{exit}}$ penalty, which shifts the best-response intercept $\tilde{p}_j = p_j - \rho_j A_j^{\text{exit}}/\left(2y_jA_j\right)$ *down* relative to baseline while leaving the slope q_j unchanged. A more exit-exposed domain (θ_j higher) suffers a sharper intercept shift because A_j^{exit} grows linearly in θ_j, and the negative cross-derivative of the intercept shift in θ_j translates one-for-one into a more negative $\partial\tau_j^*/\partial\theta_j$. This is the "tax-base-versus-extraction trade-off" of the main-text footnote: the same force that drives $\partial\tau^*/\partial\theta < 0$ in Case 2 of the baseline (where rising exit pressure makes marginal extraction more costly than at the Laffer optimum) is reinforced, not replaced, under permanent exit. Every sign remains intact; one magnitude strictly grows.

B. Descriptive Statistics

B.1. Village Revolt Data

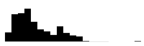
	Unique	Missing %	Mean	SD	Min	Median	Max	Histogram	Source
Basic									
Home Ruler	168	0	–	–	–	–	–	–	Aoki (1975)
Nbr Ruler	224	0	–	–	–	–	–	–	Aoki (1975)
Year	217	0	1,777.3	59.1	1,615.0	1,782.0	1,867.0		Aoki (1975)
Decade	26	0	1,772.6	59.3	1,610.0	1,780.0	1,860.0		Aoki (1975)
Outcome									
Exit	2	0	0.2	0.4	0.0	0.0	1.0		Aoki (1975)
Fight	2	0	0.8	0.4	0.0	1.0	1.0		Aoki (1975)
Predictor									
$\hat{\theta}_i$	2,529	0	7.1	61.3	0.0	1.0	2,792.0		–
$\ln(\hat{\theta}_i)$	2,529	0	0.2	1.5	−5.6	0.0	7.9		–
Dist. to Home Cap	2,517	0	45.8	85.5	0.0	21.2	1,096.0		–
Dist. to Nbr Cap	2,508	0	24.0	29.7	0.0	16.4	494.5		–
$\hat{\theta}_i^{\mathrm{LCP}}$	2,406	0	17.3	172.1	0.0	1.0	7,377.9		–
Controls									
Latitude	2,458	0	35.7	2.0	27.7	35.4	42.0		Honda, Natsume, and Nemoto (2022) and other sources
Longitude	2,445	0	136.5	3.3	128.9	136.6	142.0		Honda, Natsume, and Nemoto (2022) and other sources
Dist. to Edo	2,484	0	413.6	235.6	29.4	413.2	1,356.8		–
Dist. to Edo or Kyoto	2,484	0	234.0	152.4	0.0	191.1	1,041.7		–
Elevation	103	0	19.9	20.9	0.0	12.0	108.0		Geospatial Information Authority of Japan
Ruggedness	146	0	3.7	3.8	0.0	2.5	22.6		–
Wet-Rice Suitability	716	0	18,549.9	3,417.3	0.0	19,003.2	22,570.0		Galor and Özak (2016)
Avg. Caloric Suitability	1,133	0	2,221.3	242.7	941.1	2,297.5	2,531.2		Galor and Özak (2016)

Table A.5: Summary Statistics of the Village Revolt Data ($N = 2,741$)

B.2. Domain Data

	Unique	NA %	Mean	SD	Min	Median	Max	Hist.	Source
Outcome									
Exits	23	0	2.2	5.5	0.0	0.0	37.0		Aoki (1975)
Fights	39	0	7.0	14.4	0.0	1.0	164.0		Aoki (1975)
Revolts	45	0	9.2	17.7	0.0	2.0	201.0		Aoki (1975)
Exit Rate	56	40	0.3	0.4	0.0	0.1	1.0		Aoki (1975)
Tax Rate (τ_j)	256	7	0.4	0.1	0.1	0.4	0.7		Kodama and Kitajima (1989)
Predictor									
$\hat{\theta}_j$	241	12	7.9	16.6	0.1	2.3	138.5		–
$\ln(\hat{\theta}_j)$	241	12	1.0	1.3	−2.0	0.8	4.9		–
Dist. to Home Cap	241	12	49.1	66.8	2.5	22.4	443.0		–
Dist. to Nbr Cap	242	12	15.3	10.7	3.9	11.8	80.3		–
$\hat{\theta}_j^{\mathrm{LCP}}$	242	12	239.0	1,988.8	0.1	2.7	24,956.3		–
Controls: Main									
Wet-Rice Suitability	242	12	3,011,700.3	4,937,787.1	48,040.8	1,232,259.1	41,226,133.2		Galor and Özak (2016)
Territory Size	258	6	811.0	1,785.9	9.5	257.1	16,262.8		Honda, Natsume, and Nemoto (2022) and other sources
Population	255	7	87,781.0	159,087.3	2,466.0	30,833.0	1,065,910.0		LFAT
Pop. Density	249	9	166.5	125.1	11.9	150.2	1,285.0		–
# New Land	32	12	7.8	27.9	0.0	1.0	232.0		–
Shogunate Kin (0/1)	2	0	0.6	0.5	0.0	1.0	1.0		Kimura, Fujino, and Murakami (2015)
Goushi System (0/1)	2	0	0.1	0.3	0.0	0.0	1.0		Kimura, Fujino, and Murakami (2015)

Table A.6: Summary Statistics of the Domain Data ($N = 275$)

	Unique	Missing %	Mean	SD	Min	Median	Max	Histogram	Source
Controls: Geography									
Latitude	265	1	35.5	1.7	31.6	35.3	41.4		Kimura, Fujino, and Murakami (2015)
Longitude	265	1	136.5	3.3	128.8	136.6	141.5		Kimura, Fujino, and Murakami (2015)
Mean Elevation	242	12	16.1	13.6	2.6	12.0	88.4		–
Mean TRI	240	12	3.2	1.8	0.2	2.9	8.9		–
Dist. to Edo or Kyoto	242	12	188.5	148.9	25.2	135.0	671.3		–
% Rainy Days	251	3	0.1	0.0	0.1	0.1	0.2		Yoshimura HWDB[28]
% Snowy Days	251	3	0.0	0.0	0.0	0.0	0.1		Yoshimura HWDB
Controls: Neighbor									
Prop. Same Fam.	65	12	0.1	0.2	0.0	0.0	1.0		Kimura, Fujino, and Murakami (2015)
% New Farmland	78	12	32.9	69.2	0.0	7.5	421.0		–
Avg. Wet-Rice Suitability	242	12	2,135,287.7	3,139,273.3	288,682.9	1,350,670.1	41,057,071.8		Galor and Özak (2016)
Avg. Territory Size	242	12	433.8	602.5	12.4	278.0	6,926.3		Honda, Natsume, and Nemoto (2022) and other sources
Avg. Population	239	13	56,178.5	84,147.0	6,264.6	31,723.2	1,061,486.4		LFAT
Avg. Pop. Density	239	13	175.5	107.1	39.9	161.0	1,028.7		–
Avg. Pop. Density (Prod.)	239	13	0.0	0.0	0.0	0.0	0.2		–
Controls: Alternative									
Samurai Pop.	230	16	70,945.3	134,463.8	614.0	21,519.5	896,808.0		Kimura, Fujino, and Murakami (2015)
# Villages	170	6	196.5	304.5	1.0	82.5	2,243.0		LFAT
Dist. to Edo	242	12	370.8	244.7	42.2	309.4	1,039.5		–

Table A.7: Summary Statistics of the Domain Data (Cont.)

[28] Historical Weather Database made available online by Minoru Yoshimura: http://tk2-202-10627.vs.sakura.ne.jp/

C. Event Study Robustness

C.1. Tenryo-Conversion Universe

The treatment set is constructed from the event seed `analysis/data/domain/domain_events.csv`, itself built from (Kimura, Fujino, and Murakami 2015) and (Kodama and Kitajima 1989). Across the 1615–1868 panel window, these sources record 79 domain dissolutions. Of these, 75 (95%) were followed by direct shogunal administration (tenryō); the remaining four passed directly to another daimyō and are excluded from the treatment group per the rule stated in Section 3.3.1. The 75 dissolve-to-tenryō events correspond to 72 unique domains, three of which experienced a second dissolve-to-tenryō event in a later spell. Restricting to domains active in the panel and retaining each domain's first qualifying dissolution drops four domains whose earliest in-panel dissolution was a daimyō-to-daimyō transfer (their later tenryō conversion occurs outside the matched-control window). These filters yield the 55 treated domains listed in Table A.8.

The list spans the full sample window, with 49 cases concentrated before 1750 and 6 after. The stated reasons cluster on "no heir" (succession failure, 13 cases), "political" (9 cases, often shogunal reshuffles such as Kōfu 1704 or Tatebayashi 1683), and governance failures ("misgovernment" or "misconduct", 7 cases), with small numbers under "family dispute," "insanity," "violence," and "military" (Bakumatsu-era seizures). A handful of cases (rows 1–2, 43, 49 in Table A.8) involve lord transfers that the source classifies as dissolve-to-tenryō because the vacated capital was administered by the shogunate pending reassignment; I retain them for consistency with the source coding. The cohort split in Appendix C.6 confirms that the headline exit response loads on the pre-1750 subsample that holds 49 of the 55 cases.

No.	Year	Domain	Province	Stated Reason
1	1617	Itabashi (板橋藩)	Shimotsuke	transfer
2	1618	Nagamine (長峰藩)	Echigo	transfer
3	1619	Musashi Komuro (武蔵小室藩)	Musashi	attainder (Ina house)
4	1632	Ōmori (大森藩)	Ōmi	kokudaka halved
5	1639	Fukōzu (深溝藩)	Mikawa	residence moved
6	1642	Nariwa (成羽藩)	Bitchū	tenryō after transfer
7	1645	Fuchū (府中藩)	Hitachi	no heir
8	1650	Kaibara (柏原藩)	Tanba	family dispute (oie sōdō)
9	1653	Toyooka (豊岡藩)	Tajima	no heir
10	1656	Funai (府内藩)	Bungo	no heir
11	1657	Marugame (丸亀藩)	Sanuki	no heir
12	1658	Kakegawa (掛川藩)	Tōtōmi	no heir
13	1658	Yashima (矢島藩)	Dewa	early-period abolition
14	1660	Sakura (佐倉藩)	Shimōsa	misconduct
15	1665	Mibu (壬生藩)	Shimotsuke	no heir
16	1665	Saijō (西条藩)	Iyo	misconduct (negligence)
17	1666	Miyazu (宮津藩)	Tango	family dispute (oie sōdō)
18	1668	Shimabara (島原藩)	Hizen	misgovernment
19	1670	Harima Shingū (播磨新宮藩)	Harima	no heir
20	1671	Ichinoseki (一関藩)	Mutsu	Date affair
21	1676	Asō (麻生藩)	Hitachi	no heir
22	1677	Omigawa (小見川藩)	Shimōsa	no heir (sub-daimyō)
23	1678	Yamazaki (山崎藩)	Harima	no heir
24	1679	Kururi (久留里藩)	Kazusa	insanity
25	1679	Niwase (庭瀬藩)	Bitchū	no heir
26	1681	Numata (沼田藩)	Kōzuke	misgovernment (illicit construction)
27	1681	Takada (高田藩)	Echigo	Echigo sōdō
28	1683	Tatebayashi (館林藩)	Kōzuke	political (Tsunayoshi enthronement)

Table A.8: The 55 Tenryō-Conversion Events Used as Treatment in Table 1 (Part 1 of 2)

Note: Cases drawn from (Kimura, Fujino, and Murakami 2015) and (Kodama and Kitajima 1989) via the event seed in `analysis/data/domain/domain_events.csv`. The universe of abolitions recorded in the sources is 79; removing the 4 cases in which territory passed directly to another daimyō leaves 75 dissolve-to-tenryō events (72 unique domains, three of which experienced a second conversion in a later spell). Restricting to domains active in the 1615–1868 panel window and retaining the first qualifying dissolution per domain yields the 55 treated domains listed in Parts 1 and 2. "Stated Reason" condenses the primary motive recorded in the source notes; rows 1–2, 43, and 49 involve lord transfers that the source classifies as dissolve-to-tenryō because the vacated capital was administered by the shogunate pending reassignment, and are retained for consistency with the source coding. Pre-1750 cohort: 49 cases; post-1750 cohort: 6 cases.

No.	Year	Domain	Province	Stated Reason
29	1684	Sanuki (佐貫藩)	Kazusa	misconduct
30	1687	Fukumoto (福本藩)	Harima	partition-abolition
31	1687	Karasuyama (烏山藩)	Shimotsuke	violation of bakufu law
32	1687	Sōmi (沢海藩)	Echigo	territory confiscated
33	1689	Fukami (深見藩)	Sagami	attainder (Genroku 2)
34	1689	Kitami (喜多見藩)	Musashi	attainder (Genroku 2)
35	1692	Hachiman (八幡藩)	Mino	no heir
36	1692	Takayama (高山藩)	Hida	political (Kanamori transfer)
37	1693	Matsuyama-Bitchū (備中松山藩)	Bitchū	no heir
38	1698	Fukuyama (福山藩)	Bingo	no heir
39	1698	Sano (佐野藩)	Shimotsuke	political (Hotta transfer)
40	1698	Yoshii (吉井藩)	Kōzuke	political (Hotta transfer)
41	1701	Akō (赤穂藩)	Harima	Asano incident (violence)
42	1702	Nagashima (長島藩)	Ise	violence
43	1703	Tamanawa (玉縄藩)	Sagami	transfer to Ōtaki
44	1704	Kōfu (甲府藩)	Kai	political (Ienobu enthronement)
45	1704	Yamura (谷村藩)	Kai	political (linked to Kōfu)
46	1712	Hōjō (北条藩)	Awa	misgovernment (Man'goku uprising)
47	1724	Numazu (沼津藩)	Suruga	political (Mizuno transfer)
48	1725	Matsumoto (松本藩)	Shinano	insanity (assault in Edo castle)
49	1746	Takayanagi (高柳藩)	Echigo	transfer to Mikusa
50	1767	Obata (小幡藩)	Kōzuke	Meiwa incident (Yamagata Daini)
51	1786	Sagara (相良藩)	Tōtōmi	political (Tanuma fall)
52	1788	Ōmi Komuro (近江小室藩)	Ōmi	territory confiscated
53	1807	Matsumae (松前藩)	Ezo	Ezo defence (temporary seizure)
54	1866	Hamada (浜田藩)	Iwami	Chōshū War defeat
55	1868	Kaifuchi (貝淵藩)	Kazusa	Boshin War (pro-bakufu)

Table A.9: The 55 Tenryō-Conversion Events Used as Treatment in Table 1 (Part 2 of 2)

Note: Continuation of Table A.8. Same source definitions apply.

C.2. Tenryo-Conversion Count Robustness

Outcome:	Nearby Exit Count			Nearby Fight Count		
Radius:	30km	40km	50km	30km	40km	50km
Converted × Post	0.0359	0.0592	0.0770	0.0057	0.0103	0.0398
Cluster SE	(0.0238)	(0.0323)*	(0.0396)*	(0.0162)	(0.0228)	(0.0284)
Conley SE (70km)	(0.0291)	(0.0390)	(0.0448)*	(0.0171)	(0.0256)	(0.0314)
Num. Obs.	33,877	33,877	33,877	33,877	33,877	33,877
Domain FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Tenryo-Converted/All Domains	55/294	55/294	55/294	55/294	55/294	55/294

Table A.10: Count Robustness for Domain-to-Tenryo Conversions

Note: OLS fixed-effects estimates with event counts (rather than binary indicators) as outcomes, using the same annual domain-year panels as Table 1. Tenryo-converted domains are restricted to rel_year in [−5, +5] and matched to never-converted controls observed in the same calendar years represented by the conversion window. Two SE rows: (i) two-way clustered on domain and year; (ii) Conley spatial standard errors with a 70 km cutoff on domain capital coordinates. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

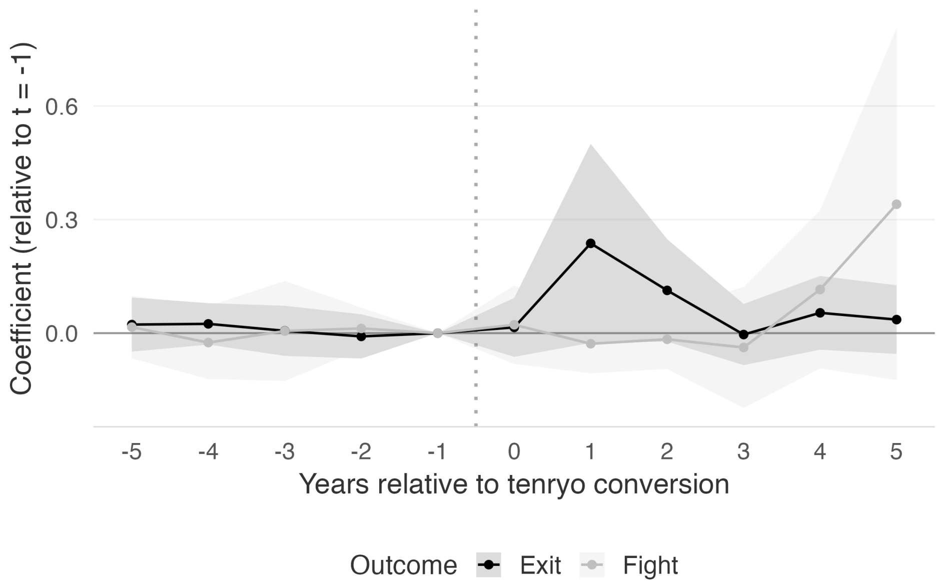


Figure A.9: OLS Count Event-Study Coefficients (50km)

Note: Year-by-year OLS coefficients from the specification $\text{count}_{j,t} = \sum_{s\neq -1} \beta_s \big(\mathbf{1}\big[t - t_j^* = s\big]\big) + \varphi_j + \omega_t + \varepsilon_{j,t}$, where the outcome is the count of nearby exit or fight events within 50km of domain j's capital. Two-way clustered (domain & year) standard errors. Shaded bands show 95% confidence intervals.

C.3. Disaggregating Exit into Desertion and End-Run Appeal

Table A.11 replicates Table 1 with the exit outcome restricted to desertion only (*chōsan*), excluding end-run appeals.

Outcome:	Any Nearby Desertion		
Sample Radius:	30km	40km	50km
Variables			
Converted × Post	0.0114	0.0176	0.0266
Cluster SE	(0.0063)*	(0.0101)*	(0.0122)**
Controls			
Domain & Year FE	Yes	Yes	Yes
Fit Statistics			
Num. Obs.	33,877	33,877	33,877
R2	0.033	0.044	0.055
Mean Dep. Var.	0.004	0.006	0.010

Table A.11: Desertion-Only DiD

Note: Replicates Table 1 with the exit outcome restricted to desertion (*chōsan*), excluding end-run appeals. Specification matches Table 1. Two-way clustered (domain & year) standard errors. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

C.4. Staggered-Robust Estimation

Table A.12 compares the main TWFE pooled estimate with the interaction-weighted estimator of Sun and Abraham (2021), which is robust to heterogeneous treatment effects under staggered adoption. The Sun and Abraham aggregate ATT for exit is 0.0413 ($p < 0.001$), virtually identical to the TWFE pooled estimate of 0.0385; the fight ATT remains insignificant at 0.0089 ($p = 0.262$). The near-identical TWFE and Sun-Abraham point estimates are consistent with the Goodman-Bacon decomposition in Table A.13, in which 94.5% of the TWFE estimand weight falls on clean treated-versus-never-treated 2×2 comparisons; the bad-comparison problem of heterogeneous-timing DiD is quantitatively small in this panel because never-treated domains make up the overwhelming majority of the sample.[29]

Estimator	Nearby Exit			Nearby Fight		
	ATT	SE	p	ATT	SE	p
TWFE	0.0385	(0.0147)**	0.01	0.0076	(0.0122)	0.532
Sun & Abraham (2021)	0.0413	(0.0045)***	< 0.001	0.0089	(0.0079)	0.262

Table A.12: TWFE vs. Sun and Abraham Staggered-Robust Estimates

Note: The TWFE row reproduces the main 50km specification from Table 1. The Sun and Abraham row reports the interaction-weighted estimator that is robust to heterogeneous treatment effects under staggered adoption. Never-converted domains serve as the comparison group. Two-way clustered (domain & year) standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

[29]The Callaway and Sant'Anna (2021) estimator, another widely used staggered-robust method, is preferred on many panels but is ill-suited here because of cell sparsity: several of the 55 distinct conversion cohorts contain only one or two treated domains, which breaks the estimator's influence-function-based inference. Sun and Abraham's interaction-weighted approach, which operates within the standard fixed-effects regression framework, accommodates sparse cohorts more naturally.

C.5. Goodman-Bacon Decomposition of the TWFE Estimator

Table A.13 reports the Goodman-Bacon (2021) decomposition of the two-way fixed-effects estimator on the balanced 50 km panel used for the tenryo DiD. The weight on clean treated-versus-never-treated 2×2 comparisons is 94.5%; only 3.8% of the weight falls on the "bad comparison" of already-treated units serving as controls for later-treated units, and 1.7% on yet-to-be-treated comparisons. The dominant never-treated weight means that the near-identical TWFE and Sun-Abraham point estimates reported in Table A.12 are a mechanical consequence of sample composition, not a test of cohort-effect homogeneity.

Comparison type	Weight	Avg. 2×2 estimate
Treated vs. never-treated	94.5%	0.0101
Treated vs. already-treated	3.8%	-0.0024
Treated vs. yet-to-be-treated	1.7%	0.0151

Table A.13: Goodman-Bacon Decomposition of the 50 km Tenryo DiD

Note: The TWFE-on-balanced-panel estimate is 0.0097. Never-treated units make up 82.6% of the sample. The dominant weight on treated-vs-never-treated 2×2 comparisons means the TWFE-vs-Sun-Abraham agreement reported in Table A.12 is mechanical, not evidence of cohort-effect homogeneity.

C.6. Cohort-Split Sun and Abraham ATT

To inspect heterogeneity directly, Table A.14 re-runs the Sun and Abraham (2021) interaction-weighted estimator separately for treated domains whose tenryo conversion year falls in each of three bins: pre-1750, 1750–1799, and post-1800. The pre-1750 cohort holds 49 of the 55 conversions and delivers an aggregate ATT of 0.0443 ($p < 0.001$), essentially reproducing the pooled result. The post-1800 cohort contains only three conversions but also yields a positive ATT of 0.0367 ($p < 0.001$). The small 1750–1799 bin (three conversions) shows a marginally negative point estimate (−0.0064) that is not load-bearing for the aggregate effect. The exit response is therefore concentrated in the period when the tenryo-conversion instrument has the most statistical weight, and no cohort reverses the sign of the effect.

Cohort bin	Cohorts	Treated units	ATT	SE	p
All cohorts	44	55	0.0413	(0.0045)***	< 0.001
pre-1750	38	49	0.0443	(0.0049)***	< 0.001
1750–1799	3	3	−0.0064	(0.0019)***	0.001
post-1800	3	3	0.0367	(0.0016)***	< 0.001

Table A.14: Cohort-Split Sun and Abraham (2021) ATT on Nearby Exit (50 km)

Note: Each row re-runs the interaction-weighted estimator on the sub-sample of treated domains whose tenryo-conversion year falls in the indicated bin (plus all never-treated controls). Two-way clustered (domain and year) standard errors in parentheses. Signif. codes: .***: 0.01, .**: 0.05, .*: 0.1.

C.7. Rambachan-Roth Relative-Magnitudes Breakdown

Table A.15 reports the full relative-magnitudes sensitivity curve of Rambachan and Roth (2023) for the average post-conversion exit ATT at 50 km. $\overline{M}$ is the ratio of the largest post-period deviation from parallel trends to the largest pre-period deviation. The 95% robust CI on the average post-conversion exit effect excludes zero for every $\overline{M} \leq 0.20$; this is the breakdown value $\overline{M}^*$. The exit pre-trend joint Wald test (the basis for the relative-magnitudes scale) is reassuring in its own right: $F = 1.26$, $p = 0.283$ on four pre-period coefficients, consistent with genuinely flat pre-trends. The sensitivity exercise therefore confirms that the main exit effect withstands post-period deviations up to one-fifth of the largest pre-period deviation, on top of a pre-trend joint test that does not reject parallel trends in the first place.

Panel A: Δ^{RM} (Relative Magnitudes)		
$\overline{M}$	95% robust CI	Conclusion
0.00	$[0.0108, 0.0694]$	excludes 0
0.20	$[0.0022, 0.0935]$	excludes 0
0.50	$[-0.0355, 0.1403]$	includes 0
1.00	$[-0.1157, 0.2224]$	includes 0
2.00	$[-0.2798, 0.3081]$	includes 0
Panel B: Δ^{SD} (Smoothness)		
M	95% robust CI	Conclusion
0.0000	$[0.0227, 0.0773]$	excludes 0
0.0010	$[0.0108, 0.0855]$	excludes 0
0.0016	$[0.0011, 0.0929]$	excludes 0
0.0020	$[-0.0054, 0.0978]$	includes 0
0.0030	$[-0.0215, 0.1098]$	includes 0
0.0050	$[-0.0535, 0.1323]$	includes 0
0.0100	$[-0.1453, 0.1599]$	includes 0

Table A.15: Rambachan-Roth (2023) Sensitivity Analysis

Note: Panel A: Δ^{RM} (relative magnitudes). $\overline{M}$ bounds the maximum post-period deviation from parallel trends as a fraction of the largest pre-period deviation. Breakdown: $\overline{M} = 0.20$. Panel B: Δ^{SD} (smoothness). M bounds the maximum second difference of the trend. Breakdown: $M = 0.0016$. Both panels report 95% robust CIs for the average post-conversion exit ATT at 50 km. Exit pre-trend joint Wald: $F = 1.26$, $p = 0.283$ (df = 4).

C.8. Rambachan-Roth Sensitivity at Canonical Benchmarks

As a companion to Table A.15, Table A.16 reports the robust 95% confidence interval for the average post-conversion exit ATT at the canonical Rambachan and Roth (2023) benchmarks $\overline{M} \in \{0.5, 1, 2\}$, so that readers can read the sensitivity at the scale RR itself recommends for comparison. The CI excludes zero at $\overline{M} \leq 0.20$; at $\overline{M} = 0.5$ the CI is $[-0.036, 0.140]$, at $\overline{M} = 1$ it is $[-0.116, 0.222]$, and at $\overline{M} = 2$ it is $[-0.280, 0.308]$. Read together with the pre-period Wald test ($F = 1.26$, $p = 0.283$, which does not reject parallel trends), the effect withstands post-period deviations up to one-fifth of the largest pre-period deviation; the $\overline{M} = 0.5, 1, 2$ columns transparently report where that envelope widens. The smoothness restriction Δ^{SD} tells the same story on a different scale: the CI excludes zero for second-difference bounds up to $M = 0.0016$, and at the empirical pre-period dispersion $\sigma_{\text{pre}} = 0.0103$ (the standard deviation of the four pre-period event-time coefficients) the CI is $[-0.148, 0.164]$. The sensitivity curve is monotone, and the breakdown values $\overline{M}^* = 0.20$ and $M^* = 0.0016$ are the sharp envelope on how much untested post-period differential trend the sign of the effect tolerates.

Panel A: Δ^{RM} (Relative Magnitudes)		
$\overline{M}$	95% robust CI	Conclusion
0.00	[0.0108, 0.0694]	excludes 0
0.20	[0.0022, 0.0935]	excludes 0
0.50	[-0.0355, 0.1403]	includes 0
1.00	[-0.1157, 0.2224]	includes 0
2.00	[-0.2798, 0.3081]	includes 0
Panel B: Δ^{SD} (Smoothness)		
M	95% robust CI	Conclusion
0.0000 (tight)	[0.0227, 0.0773]	excludes 0
0.0016 (breakdown)	[0.0011, 0.0929]	excludes 0
$\sigma_{\text{pre}} = 0.0103$	[-0.1484, 0.1635]	includes 0

Table A.16: Extended Rambachan-Roth (2023) Sensitivity Analysis

Note: Extension of Table A.15 that reports the full relative-magnitudes curve at the canonical benchmarks $\overline{M} \in \{0, 0.2, 0.5, 1, 2\}$ and the smoothness restriction Δ^{SD} at both the breakdown value ($M = 0.0016$) and the empirical pre-period dispersion ($\sigma_{\text{pre}} = 0.0103$, the SD of the four pre-period event-time coefficients). The average post-conversion exit ATT at 50 km tolerates post-period deviations up to 20% of the largest pre-period deviation under Δ^{RM} and up to a second-difference bound of $M = 0.0016$ under Δ^{SD}; larger bounds include zero. Exit pre-trend joint Wald: $F = 1.26$, $p = 0.283$.

C.9. SUTVA and Spatial Interference

Because tenryo conversions cluster in time and space, some domain-years in the 50 km DiD panel fall within the treatment radius of multiple contemporaneous conversions, potentially violating the stable unit treatment value assumption (SUTVA). We bound this concern with three exclusion exercises. First, we map the overlap: 9.0% of domain-year observations have at least one other conversion within 50 km and ±5 years. Second, we drop all converted domains whose capitals lie within 100 km of another conversion with a conversion date within 5 years ("contaminated" treated units), removing 30 of the 55 converted domains. Third, we exclude all never-converted control domains whose capitals fall within 50 km of any converted capital, as these controls may be indirectly treated (147 of 239 controls excluded). Table A.17 reports the results. The exit coefficient retains its positive sign and comparable magnitude across all four samples. The point estimates range from 0.027 to 0.038; the reduction in statistical significance in the more restrictive samples reflects the loss of treated units rather than a change in the underlying effect. The fight coefficient remains small and statistically insignificant throughout.

	Nearby Exit		Nearby Fight	
Sample	$\hat{\beta}$	SE	$\hat{\beta}$	SE
Full sample (N = 33,877)	0.0385	(0.0147)***	0.0076	(0.0122)
	(0.0131)***			(0.0136)
Drop overlap (N = 33,547)	0.0302	(0.0207)	0.0067	(0.0203)
	(0.0207)			(0.0227)
Clean controls (N = 13,886)	0.0382	(0.0155)**	0.0053	(0.0125)
	(0.0132)***			(0.0145)
Strictest (N = 13,556)	0.0270	(0.0216)	0.0031	(0.0214)
	(0.0213)			(0.0249)
Domain FE: Yes Year FE: Yes				

Table A.17: SUTVA / Spatial Interference Robustness

Note: Each row re-estimates the main 50 km DiD specification on progressively restricted samples. "Drop overlap" removes converted domains within 100 km and 5 years of another conversion. "Clean controls" removes never-converted domains within 50 km of any converted capital. "Strictest" applies both exclusions. Two SE rows are reported in parentheses: (i) two-way clustered on domain and year and (ii) Conley spatial standard errors with a 70 km cutoff. Significance stars on each SE row reflect that estimator's p-value. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

C.10. Permutation Test (Full-Pool Randomization)

As a companion to the narrower permutation test reported in the main text, which randomly reassigns fake conversion events among the 239 never-converted domains and serves as a panel-structure placebo, I also run the stricter Fisher-style randomization inference in which the placebo conversion events can land on any of the 294 domains, including the actual 55. This version asks the identification-oriented question: was the actual treatment assignment unusual among all possible 55-out-of-294 assignments? I draw 55 placebo conversion events from the full 294-domain pool in each of 500 iterations, with conversion-year timing drawn from the empirical distribution of the 55 actual conversion years. Each iteration re-applies the same $[-5, +5]$ event window and matched-control-years restriction used in Table 1 and re-estimates the 50 km pooled TWFE exit coefficient. The permutation distribution is centered at zero (mean -0.0001, SD 0.0140) and the actual estimate of 0.0385 sits far into the right tail: the two-sided p-value is 0.006 and the one-sided right-tail p-value is 0.002. Both improve on the panel-structure version ($p = 0.012$) and confirm that the exit response is specific to the actual identity and timing of tenryo conversions rather than an artifact of the staggered-adoption panel.

Full-Pool Permutation Test (50 km, Any Nearby Exit)	
Actual $\hat{\beta}$ (Converted × Post)	0.0385
Cluster SE p-value (actual)	0.0098
Permutation iterations	500
Permutation pool	All 294 domains
Placebo events per iteration	55 (sampled without replacement from 294)
Conversion-year draw	Empirical distribution of actual 55 conversion years
Permutation mean $\hat{\beta}$	-0.0001
Permutation SD	0.0140
Permutation 2.5% quantile	0.0281
Permutation 97.5% quantile	-0.0001
Two-sided p-value ($\lvert\text{perm}\rvert \geq \lvert\text{actual}\rvert$)	**0.0060**
One-sided right-tail p (perm $\geq$ actual)	0.0020

Table A.18: Full-Pool Permutation Test for the Exit Coefficient

Note: Permutation test in which the 55 placebo conversion events are drawn from the full 294-domain pool (rather than only the 239 never-converted controls, which is the narrower "panel-structure" placebo reported in the main text). Conversion-year timing is drawn from the empirical distribution of the 55 actual conversion years. Each iteration assigns 55 randomly selected domains as "converted," applies the same $[-5, +5]$ event window and matched-control-years restriction used in Table 1, and re-estimates the 50 km pooled TWFE exit coefficient. The reported two-sided p-value is the fraction of permutation estimates whose absolute value weakly exceeds the actual absolute estimate.

C.11. Hazard Model for Conversion Timing

The F-test on pre-period event-time coefficients reported in the main text (Table 1) bounds one threat to identification: a secular differential trend between treated and control domains over the five-year window leading up to conversion. A distinct and more focused concern is whether the shogunate's attainder decisions responded to **recent** local unrest, in which case the conversion year itself is endogenous to nearby exit or fight intensity. I test for this directly with a hazard model on the panel of 294 domains over 1615-1868. The outcome is an indicator equal to one if the domain is converted to tenryo in year t and zero otherwise; domains are censored at conversion (treated) or 1868 (never-converted). The covariates are 5-year moving averages (lagged by one year) of any-nearby-exit and any-nearby-fight within 50 km, plus time-invariant controls for territory size, distance to the Edo-Kyoto axis, and alliance status (fudai/kin versus tozama). Panel B of Table A.19 reports a discrete-time logit, and Panel A reports a Cox proportional-hazards estimator. Neither lagged nearby-exit intensity ($p = 0.1447$ Cox, $p = 0.3608$ logit) nor lagged nearby-fight intensity ($p = 0.9495$ Cox, $p = 0.5485$ logit) predicts the timing of conversion; the joint Wald test that both coefficients are zero fails to reject at $p = 0.3439$. The one significant predictor is the fudai/kin dummy (hazard ratio 3.8364, $p = 0.0077$): the shogunate disproportionately converted domains held by its own allies, which is consistent with the historiography on attainder politics (Ravina 1999) and unrelated to the argument in this paper. The hazard evidence, together with the pre-trend F-test and the full-pool permutation, combines into a substantial defense against endogenous treatment timing.

Panel A: Cox proportional hazards				
Variable	$\hat{\beta}$	SE	p	Hazard ratio
Nearby exit MA(5)	-5.1838	3.5539	0.1447	0.0056
Nearby fight MA(5)	0.1308	2.0674	0.9495	1.1398
log(area)	-0.2334	0.1615	0.1484	0.7918
log(dist to Edo-Kyoto axis)	0.1912	0.2951	0.5171	1.2106
Fudai/kin (vs. tozama)	1.3445	0.5047	0.0077	3.8364
Joint Wald (exit + fight = 0)	chi-square = 2.135, df = 2			$p = 0.3439$
Panel B: Discrete-time logit				
Variable	$\hat{\beta}$	SE	p	Odds ratio
Nearby exit MA(5)	-3.1981	3.4998	0.3608	0.0408
Nearby fight MA(5)	-1.1433	1.9052	0.5485	0.3188
log(area)	-0.1733	0.1575	0.2712	0.8409
log(dist to Edo-Kyoto axis)	0.1746	0.2901	0.5473	1.1907
Fudai/kin (vs. tozama)	1.3142	0.5070	0.0095	3.7218

Table A.19: Hazard Models for Tenryo Conversion Timing

Note: Cox proportional-hazards and discrete-time logit estimates of the tenryo conversion hazard. The panel contains 51,849 domain-years across 294 domains from 1615 to 1868, with 34 tenryo-conversion events in the estimation sample (34 of the 55 conversions fall after the first-five-year MA warm-up). Nearby-exit and nearby-fight variables are lagged 5-year moving averages of any_exit and any_fight within 50 km of the focal domain's capital. Neither lagged nearby exit nor lagged nearby fight intensity predicts conversion timing at any conventional level under either estimator; the joint Wald test of their sum fails to reject at $p = 0.3439$. The only significant predictor is fudai/kin alliance status, consistent with the shogunate's well-documented tendency to convert allied domains where it held attainder authority and not with a story in which conversions responded to local unrest.

C.12. Conley Bandwidth Sweep (DiD)

The 70 km Conley cutoff used for the DiD main-text standard errors is approximately one first-order spatial neighbor distance, but is a researcher degree of freedom. Table A.20 sweeps the cutoff across five standard values (30, 50, 70, 100, 150 km). The point estimate is invariant to the inference method; the Conley SE varies modestly (0.0103 to 0.0131) with all p-values remaining below 0.005.

Inference	Cutoff	$\hat{\beta}$	SE	p	N
Cluster (domain+year)	–	0.0385	0.0147	0.0098	33877
Conley	30 km	0.0385	0.0123	0.0017	33877
Conley	50 km	0.0385	0.0127	0.0024	33877
Conley (paper)	70 km	0.0385	0.0131	0.0033	33877
Conley	100 km	0.0385	0.0114	0.0007	33877
Conley	150 km	0.0385	0.0103	0.0002	33877

Table A.20: Conley Bandwidth Sweep for the Tenryo DiD (50 km)

Note: Tenryo DiD exit coefficient with standard errors computed under the cluster estimator (the main-text default) and Conley spatial HAC standard errors at five cutoffs. The point estimate does not depend on the inference method; only the SE and p-value vary. The hand-picked 70 km cutoff used in Table 1 sits near the middle of the sweep.

C.13. Kelly (2020) Data-Driven Conley Bandwidth

As an alternative to the hand-picked cutoffs used in the main text (10 km at the village and grid levels, 70 km at the domain and DiD levels), I apply the Kelly (2019) data-driven rule, which selects the bandwidth from the empirical residual correlogram: the smallest distance at which the mean pairwise residual correlation falls below $\exp(-1) \approx 0.37$. Table A.21 reports this bandwidth alongside the hand-picked choice for three analyses: the village cross-section (Table 3 Model 5), the domain-level tax regression (Table 4 Model 4), and the tenryo DiD (Table 1 50 km). The data-driven choice is tighter than the hand-picked one in all three cases (6 km for villages, 10 km for tax, 10 km for DiD). Conley SEs change by less than 0.001 between the two choices in every regression, and all three analyses remain significant at the $p < 0.01$ level.

Analysis	Bandwidth rule	Cutoff	$\hat{\beta}$	Conley SE	Conley p	N
Village-level exit share (Table 3 Model 5 dyadic-FE)						
	hand-picked	10km	0.0212	0.0078	0.0068	2739
	Kelly (2020) data-driven	6km	0.0212	0.0079	0.0074	2739
Domain tax regression (Table 4 Model 4)						
	hand-picked	70km	-0.0112	0.0055	0.0432	230
	Kelly (2020) data-driven	10km	-0.0112	0.0055	0.0449	230
Tenryo DiD (Table 1 50 km)						
	hand-picked	70km	0.0385	0.0131	0.0033	33877
	Kelly (2020) data-driven	10km	0.0385	0.0126	0.0022	33877

Table A.21: Kelly (2020) Data-Driven vs. Hand-Picked Conley Bandwidth

Note: The Kelly bandwidth is the smallest distance at which the empirical residual correlogram (binned in 20 distance bins, pairs aggregated across the regression sample) falls below $\exp(-1) \approx 0.37$. For the village and DiD regressions the Kelly bandwidth is tighter than the hand-picked cutoff; for the tax regression it is substantially tighter (10 km vs. 70 km). Point estimates are identical across bandwidth choices; only SEs vary. The data-driven choice moves SEs by less than 0.001 in absolute terms across all three regressions, so the main-text p-values are robust to the bandwidth selection rule.

D. Village Exit Choice Analysis

D.1. Full Table

Outcome:	exit$_i$					
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
$\ln\left(\hat{\theta}_i\right)$	0.0458	0.0235	0.0143	0.0201	0.0212	0.0244
	(0.0154)***	(0.0117)*	(0.0089)	(0.0071)***	(0.0072)***	(0.0068)***
lat					-2.417^{***}	-2.413^{*}
					(0.8551)	(1.460)
lon					-0.5462^{**}	-0.5291
					(0.2119)	(0.3918)
ln(Dist. Edo)					0.2731^{**}	0.2241
					(0.1094)	(0.1395)
ln(Rice)					-0.0048	-0.0027
					(0.0046)	(0.0048)
ln(Elev.)					-0.0216	-0.0291
					(0.0212)	(0.0214)
ln(Rugg.)					0.0276^{**}	0.0328^{**}
					(0.0150)	(0.0152)
Lat $\times$ Lon					0.0174^{**}	0.0171
					(0.0064)	(0.0112)
Controls						
Home (j) Domain FE	No	Yes	Yes	Yes	Yes	No
Neighbor (k) Domain FE	No	No	Yes	Yes	Yes	No
Decade (t) FE	No	No	No	Yes	Yes	Yes
Home-Neighbor Dyad (j, k) FE	No	No	No	No	No	Yes
Fit Statistics						
Num.Obs.	2,739	2,739	2,739	2,739	2,739	2,739
R2	0.028	0.390	0.543	0.621	0.626	0.678

Table A.22: Villages Exposed to Alternative Authority Opt for Exit Over Fight (Full Table)

Note: The coefficients represent OLS estimates of the relationship between the distance ratio measure of each village's exposure to external authority and each village's choice of exit upon revolting. Clustered (Home & nearest neighboring Domains & Decade) standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2. Robustness Checks

D.2.1. Alternative Covariates

Outcome:	exit_i							
	Cal. Suit.				Dist. to Edo			
Models:	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Variables								
$\ln(\hat{\theta}_i)$	**0.0434**	**0.0162**	**0.0209**	**0.0234**	**0.0439**	**0.0135**	**0.0199**	**0.0240**
	(0.0150)***	**(0.0090)***	**(0.0072)*****	**(0.0069)*****	**(0.0149)*****	**(0.0092)**	**(0.0075)****	**(0.0068)*****
lat	−1.3685	−0.8994	−2.5470	−2.5127	−1.7680	−3.0658	−3.5340	−3.6532
	(0.2726)	(0.9761)	(0.8580)	(1.3351)	(0.5527)	(1.0682)	(0.9003)	(1.2780)
lon	−0.3040	−0.0596	−0.5621	−0.5293	−0.4533	−0.7529	−0.8904	−0.9021
	(0.0711)	(0.2402)	(0.2012)	(0.3450)	(0.1633)	(0.3249)	(0.2710)	(0.3792)
ln(Dist. Edo)	−0.0238	0.4825	0.2846	0.2497				
	(0.0449)	(0.2092)	(0.1153)	(0.1346)				
ln(Cal.)	−0.0307	0.1270	0.1926	0.2602				
	(0.2051)	(0.1825)	(0.1566)	(0.1590)				
ln(Elev.)	0.0132	−0.0137	−0.0268	−0.0369	0.0009	−0.0204	−0.0262	−0.0329
	(0.0250)	(0.0193)	(0.0188)	(0.0180)	(0.0196)	(0.0215)	(0.0208)	(0.0212)
ln(Rugg.)	0.0324	0.0210	0.0274	0.0319	0.0445	0.0316	0.0331	0.0368
	(0.0246)	(0.0161)	(0.0153)	(0.0154)	(0.0209)	(0.0160)	(0.0154)	(0.0158)
Lat × Lon	0.0096	0.0057	0.0183	0.0178	0.0127	0.0223	0.0260	0.0266
	(0.0019)	(0.0073)	(0.0064)	(0.0102)	(0.0041)	(0.0083)	(0.0070)	(0.0100)
ln(Dist. Edo)					−0.1417	−0.1751	−0.0181	−0.1370
					(0.0723)	(0.3853)	(0.2963)	(0.3973)
ln(Rice)					−0.0082	−0.0079	−0.0046	−0.0027
					(0.0069)	(0.0064)	(0.0045)	(0.0048)
Const.	44.25***				64.05***			
	(9.996)				(22.24)			
Controls								
Home (j) Domain FE	No	Yes	Yes	No	No	Yes	Yes	No
Neighbor (k) Domain FE	No	Yes	Yes	No	No	Yes	Yes	No
Decade (t) FE	No	No	Yes	Yes	No	No	Yes	Yes
Home-Neighbor Dyad (j,k) FE	No	No	No	Yes	No	No	No	Yes
Fit Statistics								
Num.Obs.	2,739	2,739	2,739	2,739	2,739	2,739	2,739	2,739
R2	0.061	0.551	0.627	0.679	0.069	0.549	0.626	0.678

Table A.23: Villages Exposed to Alternative Authority Opt for Exit Over Fight (Alternative Covariates)

Note: Clustered (Home & Neighbor Domains & Decade) standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2.2. Excluding Villages Rebelling More Than Once

Outcome:	exit$_i$					
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
$\ln(\hat{\theta}_i)$	**0.0488**	**0.0270**	**0.0228**	**0.0229**	**0.0234**	**0.0231**
	(0.0153)***	**(0.0119)****	**(0.0104)****	**(0.0094)****	**(0.0094)****	**(0.0080)*****
lat					−2.6613	−2.8835
					(0.9798)	(1.5600)
lon					−0.6092	−0.6572
					(0.2415)	(0.4214)
ln(Dist. Edo)					0.2075	0.1153
					(0.0702)	(0.0881)
ln(Rice)					−0.0061	−0.0040
					(0.0049)	(0.0051)
ln(Elev.)					−0.0267	−0.0371
					(0.0214)	(0.0222)
ln(Rugg.)					0.0309*	0.0360**
					(0.0157)	(0.0161)
Lat × Lon					0.0194	0.0208
					(0.0073)	(0.0120)
Const.	0.2299***					
	(0.0348)					
Controls						
Home (j) Domain FE	No	Yes	Yes	Yes	Yes	No
Neighbor (k) Domain FE	No	No	Yes	Yes	Yes	No
Home-Neighbor Dyad (j, k) FE	No	No	No	No	No	Yes
Decade (t) FE	No	No	No	Yes	Yes	Yes
Geography	No	No	No	No	Yes	Yes
Fit Statistics						
Num.Obs.	2,295	2,295	2,295	2,295	2,295	2,295
R2	0.032	0.431	0.590	0.672	0.675	0.732

Table A.24: Villages Exposed to Alternative Authority Opt for Exit Over Fight (Excluding Repeat Villages)

Note: Clustered (Home & Neighbor Domains & Decade) standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2.3. Absolute Distances

Outcome:	exit_i					
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
ln(Dist. to Home Capital$_i$)	**0.0399**	**0.0248**	**0.0127**	**0.0231**	**0.0160**	**0.0203**
	(0.0223)*	**(0.0161)**	**(0.0109)**	**(0.0096)****	**(0.0104)**	**(0.0083)****
ln(Dist. to Foreign Capital$_i$)	**−0.0596**	**−0.0203**	**−0.0172**	**−0.0147**	**−0.0306**	**−0.0309**
	(0.0140)***	**(0.0218)**	**(0.0199)**	**(0.0147)**	**(0.0129)****	**(0.0135)****
lat					−2.5043	−2.5427
					(0.8401)	(1.4573)
lon					−0.5607	−0.5495
					(0.2091)	(0.3908)
ln(Dist. Edo)					0.2839	0.2405
					(0.1131)	(0.1354)
ln(Rice)					−0.0047	−0.0024
					(0.0047)	(0.0048)
ln(Elev.)					−0.0192	−0.0252
					(0.0217)	(0.0220)
ln(Rugg.)					0.0292	0.0350
					(0.0157)	(0.0155)
Lat × Lon					0.0180	0.0181
					(0.0063)	(0.0112)
Const.	0.2692					
	(0.0683)					
Controls						
Home (j) Domain FE	No	Yes	Yes	Yes	Yes	No
Neighbor (k) Domain FE	No	No	Yes	Yes	Yes	No
Home-Neighbor Dyad FE	No	No	No	No	No	Yes
Decade (t) FE	No	No	No	Yes	Yes	Yes
Geography	No	No	No	No	Yes	Yes
Fit Statistics						
Num.Obs.	2,739	2,739	2,739	2,739	2,739	2,739
R2	0.030	0.390	0.543	0.621	0.627	0.678

Table A.25: Villages Exposed to Alternative Authority Opt for Exit Over Fight (Abs. Dist. Measures)

Note: Clustered (Home & Neighbor Domains & Decade) standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2.4. Least-Costly Path Distance

Outcome:	exit$_i$					
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
$\ln\left(\hat{\theta}_i^{\text{LCP}}\right)$	**0.0358**	**0.0215**	**0.0137**	**0.0190**	**0.0197**	**0.0112**
	(0.0120)	**(0.0109)**	**(0.0076)**	**(0.0075)**	**(0.0074)**	**(0.0066)**
lat					−2.5400	−2.4649
					(0.7116)	(1.5386)
lon					−0.5941	−0.5511
					(0.2002)	(0.3477)
ln(Dist. Edo)					0.2177	0.1888
					(0.0530)	(0.1475)
ln(Rice)					−0.0046	−0.0025
					(0.0045)	(0.0043)
ln(Elev.)					−0.0230	−0.0298
					(0.0205)	(0.0210)
ln(Rugg.)					0.0299	0.0344
					(0.0148)	(0.0150)
Lat × Lon					0.0184	0.0176
					(0.0055)	(0.0115)
Const.	0.2442					
	(0.0343)					
Controls						
Home (j) Domain FE	No	Yes	Yes	Yes	Yes	No
Neighbor (k) Domain FE	No	No	Yes	Yes	Yes	No
Home-Neighbor Dyad FE	No	No	No	No	No	Yes
Decade (t) FE	No	No	No	Yes	Yes	Yes
Geography	No	No	No	No	Yes	Yes
Fit Statistics						
Num.Obs.	2,764	2,764	2,764	2,764	2,764	2,764
R2	0.028	0.387	0.533	0.615	0.622	0.686

Table A.26: Villages Exposed to Alternative Authority Opt for Exit Over Fight (LCP Distance)

Note: The coefficients represent OLS estimates of the relationship between the least-costly path distance and each village's choice of exit upon rebelling. Clustered (Home & Neighbor Domains & Decade) standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2.5. Water-Crossing Subsample

The main village distance-ratio regression uses straight-line Euclidean distances from each village to its home-domain capital and to its nearest foreign-domain capital, and straight lines can in principle traverse sea straits such as Kanmon, Bungo, Naruto, or the Sado strait. As a non-parametric complement to the least-costly path check of Appendix D.2.4, I drop every village-revolt whose village, home capital, or nearest-foreign capital lies on a different main island (Honshu, Kyushu, Shikoku, Sado, or Awaji), classified from lat/lon bounding boxes; this removes 153 of 2,741 village-revolt rows (5.6 percent). On this water-safe subsample the coefficient on $\log\left(\hat{\theta}_i\right)$ remains positive, stable, and statistically significant across all six specifications, including in the dyadic home-neighbor FE column that mirrors Table 3 Model 5 ($\hat{\beta} = 0.0222$, clustered $p = 0.002$; baseline 0.0197, $p = 0.014$).

Outcome:			exit$_i$			
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Distance ratio ($\ln(\hat{\theta}_i)$)	0.0487	0.0229	0.0194	0.0262	0.0283	0.0222
Cluster SE	(0.0162)***	(0.0146)	(0.0122)	(0.0089)***	(0.0085)***	(0.0063)***
Conley SE (10km)	(0.0127)***	(0.0119)*	(0.0109)*	(0.0088)***	(0.0090)***	(0.0098)**
Controls						
Home Domain FE	No	Yes	Yes	Yes	Yes	No
Neighbor Domain FE	No	No	Yes	Yes	Yes	No
Home-Nbr Dyad FE	No	No	No	No	No	Yes
Decade FE	No	No	No	Yes	Yes	Yes
Geography	No	No	No	No	Yes	Yes
Fit Statistics						
Num. Obs.	2,586	2,586	2,586	2,586	2,586	2,586
R^2	0.029	0.391	0.546	0.632	0.637	0.692
Mean Dep. Var.	0.251	0.251	0.251	0.251	0.251	0.251

Table A.27: Village Distance-Ratio Regression: Water-Crossing Subsample

Note: Re-runs the six village distance-ratio specifications on the subsample that drops village-revolts whose straight-line village-to-home-capital or village-to-foreign-capital line crosses a sea strait (i.e., the village, the home capital, and the nearest foreign capital are not all on the same main island). Island membership (Honshu, Kyushu, Shikoku, Sado, Awaji) is assigned from lat/lon bounding boxes. 153 (5.6%) of 2,741 village-revolt rows are classified as sea-crossing and dropped. The coefficient on $\log(\hat{\theta}_i)$ is stable across all six specifications and statistical significance is preserved, including in Model 6 (dyadic home-neighbor FE plus geography controls), which corresponds to the paper's headline column (Table 3 Model 5). SE rows: (i) three-way clustered on (home, neighbor, decade); (ii) Conley spatial SE with a 10 km cutoff on village coordinates. Significance stars on each SE row reflect that estimator's p-value. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2.6. Conley Standard Error Bandwidth Sensitivity

Table A.28 reports the bandwidth sensitivity of Conley spatial standard errors for the main village distance-ratio regression (Table 3 Model 5, dyadic home-neighbor FE). The coefficient on $\log\left(\hat{\theta}_i\right)$ is stable across all bandwidths from 25 to 150 km and remains significant at the 5% level in every specification.

Standard error	Estimate	SE	p
Cluster (home + nbr + decade, paper)	0.0197	(0.0074)	0.014
Conley 25 km	0.0197	(0.0093)	0.034
Conley 50 km	0.0197	(0.0093)	0.034
Conley 75 km	0.0197	(0.0101)	0.05
Conley 100 km	0.0197	(0.0102)	0.052
Conley 150 km	0.0197	(0.0083)	0.018

Table A.28: Conley SE Bandwidth Sensitivity: Village Distance-Ratio Regression

Note: The coefficient on $\log\left(\hat{\theta}_i\right)$ from Table 3 Model 5 (dyadic home-neighbor FE) is stable across all bandwidths from 25 to 150 km and remains significant at the 5% level in every specification. The first row reports the three-way cluster standard error used in the main table.

D.2.7. Small-G Cluster Sensitivity of the Decade Dimension

Table A.29 reports a small-G cluster-sensitivity check for the village distance-ratio regression (Table 3 Model 5, dyadic home-neighbor FE). The decade dimension has $G = 26$ clusters, below the conventional safe threshold for multi-way cluster-robust inference, motivating a check of whether this dimension is load-bearing for the reported significance. Dropping the decade dimension entirely and clustering only on the two higher-G dimensions (home ruler and neighbor ruler) yields $p = 0.020$, essentially indistinguishable from the three-way number ($p = 0.014$) and comfortably below the 5% threshold; clustering on decade alone yields a tighter $p = 0.008$. The main-text significance is therefore preserved under every cluster specification considered.

Cluster specification	Estimate	SE	p
Three-way (home + nbr + decade, paper)	0.0197	(0.0074)	0.014
Two-way (home + nbr, dropping decade)	0.0197	(0.0084)	0.02
One-way on decade only ($G = 26$)	0.0197	(0.0068)	0.008

Table A.29: Small-G Cluster Sensitivity of the Decade Dimension

Note: Re-runs Table 3 Model 5 (dyadic home-neighbor FE) under three cluster specifications. The decade dimension has $G = 26$ clusters, below the conventional safe threshold for multi-way clustering. The key finding is that the paper's three-way cluster SE is robust to dropping the decade dimension entirely — two-way clustering on home and neighbor gives nearly identical inference — indicating the decade cluster is not load-bearing for the reported p-value.

D.2.8. Hatamoto Kantō Exclusion

Morris (1990) documents that hatamoto (direct shogunal retainers holding small fiefs) were particularly dense in four Kantō provinces: Musashi, Sagami, Kazusa, and Shimōsa. Because the distance-ratio measure in Section 3.4.2 is built on daimyō capitals and shogunal *daikansho* and does not incorporate individual hatamoto fiefs, a village in these four provinces may have a hatamoto landlord closer than its nearest measured outside authority, under-stating its true proximity to an alternative authority. As a robustness check, I re-estimate the main village regression on the sub-sample that excludes the 40 village-revolts (1.5% of the sample) in these four provinces. The coefficient on $\log\left(\hat{\theta}_i\right)$ is essentially unchanged (0.0197 → 0.0198) and remains significant at $p = 0.014$, confirming that the hatamoto omission does not drive the main result (Table A.30).

Sample	Num. Obs.	Estimate	SE	p
Full sample (paper)	2,739	0.0197	(0.0074)	0.014
Excluding Musashi, Sagami, Kazusa, Shimōsa	2,699	0.0198	(0.0075)	0.014

Table A.30: Village Distance-Ratio Regression: Excluding Dense-Hatamoto Kantō Provinces

Note: Re-runs Table 3 Model 5 (dyadic home-neighbor FE) on the sub-sample that excludes village-revolts in the four Kantō provinces with the highest documented hatamoto density – Musashi, Sagami, Kazusa, and Shimōsa (Morris 1990) – where the distance-ratio measure may under-capture the nearest outside authority because hatamoto fiefs are not in the daimyō + daikansho geography. The coefficient on $\log\left(\hat{\theta}_i\right)$ is stable and statistical significance is preserved.

D.2.9. Desertion-Only Village Cross-Section

Table A.31 re-estimates the preferred specifications with the exit outcome restricted to desertion only (*chōsan*), excluding end-run appeals. The coefficient is positive and statistically significant, confirming that the result holds under this narrower definition of exit.

Outcome:	Desertion_i	
Models:	(4)	(5)
Variables		
Distance ratio ($\ln(\hat{\theta}_i)$)	0.0127	0.0121
Cluster SE	(0.0060)**	(0.0054)**
Conley SE (10km)	(0.0062)**	(0.0062)*
Controls		
Home + Neighbor + Decade FE + Geography	Yes	No
Dyadic + Decade FE + Geography	No	Yes
Fit Statistics		
Num.Obs.	2,727	2,727
R2	0.598	0.602

Table A.31: Desertion-Only Village Cross-Section

Note: Replicates Table 3 Models 4–5 with exit restricted to desertion (*chōsan*), excluding end-run appeals. Two SE rows: (i) three-way clustered; (ii) Conley spatial SE (10 km). Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2.10. Desertion-Only Grid-Year Panel

Table A.32 replicates Table 2 with exit restricted to desertion (*chōsan*) only. The desertion-specific coefficients track the pooled exit estimates closely in both sign and significance, across the full panel and the ever-active subsample, confirming that the within-cell exit response is driven by the prototypical desertion cases rather than by end-run appeals.

	Full sample		Ever-active cells	
Outcome:	(1)	(2)	(3)	(4)
	Desertion	Pooled exit	Desertion	Pooled exit
Distance ratio ($\tilde{\theta}_{i,t}$)	0.00024	0.00035	0.00152	0.00198
Cluster SE	(0.00014)*	(0.00017)**	(0.00087)*	(0.00109)*
Conley SE (10km)	(0.00012)**	(0.00017)**	(0.00075)**	(0.00103)*
Controls				
Grid & Year FE	Yes	Yes	Yes	Yes
Fit Statistics				
Num.Obs.	2,225,802	2,225,802	372,364	372,364
R^2	0.005	0.006	0.006	0.008
Mean Dep. Var.	0.00009	0.00022	0.00054	0.00131

Table A.32: Desertion-Only Grid-Year Panel

Note: LPM on binary any-event-of-type indicators. Replicates Table 2 with exit restricted to desertion (*chōsan*), excluding end-run appeals. Pooled exit columns reproduce the main specification for comparison. Columns (1)–(2) use the full panel of 8,763 cells; columns (3)–(4) restrict to the 1,466 ever-active hexagons. Grid and year FE. SEs: (i) two-way clustered (cell, year); (ii) Conley (10 km). Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2.11. Grid-Panel Count Models: Poisson and Negative Binomial

Table A.33 replicates the main-text Table 2 using two count-data estimators with hexagon and year fixed effects: Poisson (columns 1–3) and the negative binomial (columns 4–6). Because both estimators use a log link ($\log \mathbb{E}\left[y_{i,t} \mid \mathrm{X}_{i,t}\right] = \beta \mathrm{X}_{i,t} + \alpha_i + \tau_t$), the raw coefficient $\hat{\beta}$ sits on the log-count scale. The *incidence rate ratio*, $\mathrm{IRR} \equiv \exp\left(\hat{\beta}\right)$, back-transforms this to the count scale: a one-unit increase in the regressor multiplies the expected count by IRR. Both estimators agree on the exit-vs-fight asymmetry. The exit coefficient is positive and corresponds to a multiplicative effect of roughly 4–5× per unit distance ratio (Poisson IRR $= 4.020$; NB IRR $= 5.357$). The fight coefficient is small and statistically indistinguishable from zero under both estimators (Poisson $p = 0.5676$; NB $p = 0.4891$), and the fight-strict coefficient turns negative (IRRs below one). The NB columns drop all-zero cells via the conditional likelihood, so sample size varies by outcome; the Poisson columns use the full 8,763-cell panel.

	Poisson			Neg. Binom.		
Outcome:	(1)	(2)	(3)	(4)	(5)	(6)
	Exit count	Fight count	Fight count (strict)	Exit count	Fight count	Fight count (strict)
Distance ratio ($\tilde{\theta}_{i,t}$)	1.3912	0.3240	−0.4231	1.6784	0.4453	−0.5370
IRR	4.020	1.383	0.655	5.357	1.561	0.584
Cluster SE	(0.9599)	(0.5668)	(0.5299)	(0.8860)*	(0.6437)	(0.7985)
Conley SE (10km)	(0.9043)	(0.5257)	(0.6052)	(0.8601)*	(0.6620)	(0.7938)
Controls						
Grid FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Fit Statistics						
Num.Obs.	62,900	202,130	40,810	62,900	202,130	40,810
Pseudo R^2	0.135	0.130	0.127	0.093	0.092	0.087
Mean Dep. Var.	0.0003	0.0009	0.0004	0.0003	0.0009	0.0004

Table A.33: Grid-Panel Count Models: Poisson and Negative Binomial with FE

Note: Columns (1)–(3) estimate a Poisson model and columns (4)–(6) a negative binomial, both with hexagon and year fixed effects on the 5 km grid-year panel (1615–1868). Coefficients are on the log-count scale; IRR $= \exp\left(\hat{\beta}\right)$ gives the multiplicative effect on the expected count. Num. Obs. in the NB columns reflects the ever-event subsample retained by the conditional likelihood; the Poisson columns use the full 8,763-cell panel. SEs: (i) two-way clustered on cell and year; (ii) Conley (10 km). Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2.12. Ever-Active Cells Subsample

Table A.34 replicates Table 2 on the 1,466 "ever-active" hexagons — cells with at least one exit or fight event between 1615 and 1868. This restriction filters out the 83% of cells that were politically inert throughout the period and sharpens within-variation among plausibly restive locales while using the same binary-outcome LPM and the same fixed-effects structure. Columns (1)–(3) use the distance ratio $\tilde{\theta}_{i,t}$; columns (4)–(6) use the closest-capital distance $d^1_{i,t}$ (km).

The distance-ratio coefficient on exit rises to 0.00198 ($p = 0.054$ Conley, $p = 0.070$ clustered), while the fight and fight-strict coefficients remain small and statistically indistinguishable from zero. The closest-capital regressor remains a null predictor of exit on this subsample as well ($p = 0.179$ clustered, $p = 0.259$ Conley). It does carry a marginal negative signal for fight-strict (Conley $p = 0.028$) — closer to a state authority, more strict-fight events, plausibly because uprisings and castle-attacks require a nearby state target. Mere remoteness therefore carries some signal about where fights are mechanically possible but not about where exit is chosen, while $\tilde{\theta}_{i,t}$ uniquely predicts exit. The asymmetry mirrors Table 2: governance conflict, not isolation from any ruler, channels resistance toward exit.

	Distance ratio ($\tilde{\theta}_{i,t}$)			Closest-capital dist. ($d^1_{i,t}$, km)		
Outcome:	(1)	(2)	(3)	(4)	(5)	(6)
	Any exit	Any fight	Any fight (strict)	Any exit	Any fight	Any fight (strict)
Coefficient	0.00198	0.00096	−0.00097	0.00004	−0.00004	−0.00004
Cluster SE	(0.00109)*	(0.00105)	(0.00067)	(0.00003)	(0.00003)	(0.00002)*
Conley SE (10km)	(0.00103)*	(0.00105)	(0.00062)	(0.00004)	(0.00002)*	(0.00002)**
Controls						
Grid FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Fit Statistics						
Num.Obs.	372,364	372,364	372,364	372,364	372,364	372,364
R^2	0.008	0.011	0.011	0.008	0.011	0.011
Mean Dep. Var.	0.00131	0.00384	0.00155	0.00131	0.00384	0.00155

Table A.34: Grid-Year Panel on Ever-Active Cells: Distance Ratio vs. Closest-Capital Distance

Note: Replicates Table 2 on the 1,466 "ever-active" hexagons — cells with at least one exit or fight event between 1615 and 1868 — which sharpens within-variation among plausibly restive locales while filtering out the 83% of cells that were politically inert throughout the period. Columns (1)–(3) use the time-varying distance ratio $\tilde{\theta}_{i,t} = \frac{d^1_{i,t}}{d^2_{i,t}}$. Columns (4)–(6) replace it with the raw closest-capital distance $d^1_{i,t}$ (km). Grid and year FE. SEs: (i) two-way clustered (cell, year); (ii) Conley (10 km). Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2.13. Least-Cost-Path Distance Ratio

Table A.35 replicates Table 2 with the great-circle distance ratio $\tilde{\theta}_{i,t} = \frac{d^1_{i,t}}{d^2_{i,t}}$ replaced by a terrain-weighted least-cost-path variant $\tilde{\theta}^{\text{LCP}}_{i,t}$, computed from a $(1 + \text{slope})$ cost surface on the SRTM elevation raster—the same cost surface used for the village-level LCP robustness check in Appendix D.2.4. For each of the 8,763 5 km hexagon centroids and each unique capital location, we run a single-source Dijkstra over the cost surface and record the cumulative terrain-weighted travel cost; $\tilde{\theta}^{\text{LCP}}_{i,t}$ is the ratio of the nearest-to-second-nearest such cost among capitals active in year t. The two regressors capture overlapping but distinct information: their Pearson correlation is 0.44, and the LCP ratio has mean 0.82 and standard deviation 0.24 (vs. 0.63 and 0.26 for the great-circle ratio), reflecting the common terrain component that travel costs to multiple capitals share when paths cross the same ridges and valleys.

The exit coefficient preserves its sign and standardized magnitude: a one-standard-deviation increase in $\tilde{\theta}^{\text{LCP}}_{i,t}$ corresponds to $0.00045 \times 0.24 \approx 0.011$%-points, close to the 0.009%-points implied by Table 2 column (1). The exit-vs-fight asymmetry also survives: fight and fight-strict coefficients remain small and statistically indistinguishable from zero across the full panel and the ever-active subsample, just as in the main specification. Statistical significance on exit falls below conventional thresholds ($p = 0.278$ clustered; $p = 0.203$ Conley), which is consistent with the LCP ratio being a noisier proxy for governance conflict: terrain friction adds a shared component to the cost of reaching any capital, compressing contrasts between the nearest two and absorbing some of the identifying variation that the great-circle ratio isolates. The sign-and-magnitude preservation under this alternative distance metric is the substantive message; the attenuated p-values reflect measurement noise in the mapping from LCP cost to relative accessibility of alternative authority rather than a sign-flip or explain-away of the exit effect.

	Full sample			*Ever-active cells*		
Outcome:	(1)	(2)	(3)	(4)	(5)	(6)
	Any exit	Any fight	Any fight (strict)	Any exit	Any fight	Any fight (strict)
Distance ratio (LCP, $\tilde{\theta}_{i,t}^{\mathrm{LCP}}$)	0.00045	0.00053	−0.00008	0.00173	0.00177	−0.00051
Cluster SE	(0.00041)	(0.00060)	(0.00031)	(0.00169)	(0.00221)	(0.00122)
Conley SE (10km)	(0.00035)	(0.00068)	(0.00026)	(0.00141)	(0.00192)	(0.00074)
Controls						
Grid FE	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Fit Statistics						
Num.Obs.	2,225,802	2,225,802	2,225,802	372,364	372,364	372,364
R^2	0.006	0.007	0.006	0.008	0.011	0.011
Mean Dep. Var.	0.00022	0.00064	0.00026	0.00131	0.00384	0.00155

Table A.35: Grid-Year Panel with Least-Cost-Path Distance Ratio

Note: LPM on binary any-event-of-type indicators. Replicates Table 2 with the haversine distance ratio replaced by a terrain-weighted least-cost-path (LCP) variant: $\tilde{\theta}_{i,t}^{\mathrm{LCP}}$ equals the ratio of nearest-to-second-nearest LCP cost from each hexagon centroid to any active domain capital in year t, where cost is cumulative across pixels of a $(1 + \mathrm{slope})$ cost surface derived from the SRTM elevation raster. Same cost surface used for the village-level LCP robustness check in Appendix D.2.4. Columns (1)–(3) use the full panel; (4)–(6) restrict to ever-active hexagons. Grid and year FE. SEs: (i) two-way clustered (cell, year); (ii) Conley (10 km). Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2.14. Daikansho-Inclusive Distance Ratio (Village)

Table A.36 replicates Table 3 with d_i^{FC} redefined as the distance to the nearest non-home administrative center of any type, including shogunal intendant offices (*daikansho*). The preferred dyadic-FE specification (Model 5) is virtually unchanged ($\hat{\beta} = 0.0268$ vs. 0.0244 in Table 3).

Outcome:	Exit$_i$				
Models:	(1)	(2)	(3)	(4)	(5)
Variables					
Distance ratio ($\ln(\hat{\theta}_i)$)	0.0422	0.0496	0.0201	0.0202	0.0268
Cluster SE	(0.0148)***	(0.0137)***	(0.0074)**	(0.0075)**	(0.0067)***
Conley SE (10km)	(0.0117)***	(0.0093)***	(0.0077)***	(0.0079)**	(0.0085)***
Controls					
Decade FE	No	Yes	Yes	Yes	Yes
Home Domain FE	No	No	Yes	Yes	No
Neighbor Domain FE	No	No	Yes	Yes	No
Dyadic FE	No	No	No	No	Yes
Geography	No	No	No	Yes	Yes
Fit Statistics					
Num.Obs.	2,739	2,739	2,739	2,739	2,739
R2	0.024	0.222	0.633	0.638	0.686
Mean Dep. Var.	0.246	0.246	0.246	0.246	0.246

Table A.36: Village Exit Choice with Daikansho-Inclusive Distance Ratio

Note: Replicates Table 3 with d_i^{FC} redefined to include shogunal intendant offices (*daikansho*). Two SE rows: (i) clustered on home, nearest neighboring domain, and decade; (ii) Conley spatial standard errors (10 km). Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.2.15. Daikansho-Inclusive Distance Ratio (Grid)

Table A.37 replicates Table 2 with the daikansho-inclusive distance ratio. Adding daikansho to the capital set compresses the ratio toward unity and reduces within-cell variation; all grid-level coefficients attenuate and lose statistical significance. The village-level and tax results are unaffected by this inclusion (Appendix D.2.14; Appendix D.3.1).

	Full sample			Ever-active cells		
Outcome:	(1)	(2)	(3)	(4)	(5)	(6)
	Any exit	Any fight	Any fight (strict)	Any exit	Any fight	Any fight (strict)
Distance ratio ($\tilde{\theta}_{i,t}$)	0.00019	0.00023	−0.00006	0.00095	−0.00012	−0.00096
Cluster SE	(0.00019)	(0.00021)	(0.00012)	(0.00111)	(0.00111)	(0.00068)
Conley SE (10km)	(0.00018)	(0.00021)	(0.00012)	(0.00107)	(0.00097)	(0.00060)
Controls						
Grid & Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Fit Statistics						
Num.Obs.	2,225,802	2,225,802	2,225,802	372,364	372,364	372,364
R^2	0.006	0.007	0.006	0.008	0.011	0.011
Mean Dep. Var.	0.00022	0.00064	0.00026	0.00131	0.00384	0.00155

Table A.37: Grid-Year Panel with Daikansho-Inclusive Distance Ratio

Note: LPM on binary any-event-of-type indicators. Replicates Table 2 with the daikansho-inclusive distance ratio. Columns (1)–(3) use the full panel of 8,763 cells; columns (4)–(6) restrict to the 1,466 ever-active hexagons. See Table 2 notes for details. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.3. Additional Analyses

D.3.1. Daikansho-Inclusive Distance Ratio (Tax)

Table A.38 replicates Table 4 with the daikansho-inclusive distance ratio. Models 1–2 remain significant at $p < 0.01$; models 3–4 show modest p-value drift (from $p < 0.05$ to $p \approx 0.07$) with stable point estimates, consistent with daikansho offices compressing the distance ratio's variance. The spatial lag models (5–6) preserve the negative sign but are not statistically significant, mirroring the pattern in the main table where the fully controlled spatial lag is only marginally significant.

Outcome:	Tax Rate (τ_j)					
	OLS				*Spatial Lag*	
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
Distance ratio ($\ln(\hat{\theta}_j)$)	−0.0233	−0.0192	−0.0098	−0.0096	−0.0130	−0.0080
HC SE	(0.0054)***	(0.0058)***	(0.0054)*	(0.0054)*	(0.0047)***	(0.0050)
Controls						
Domain Characteristics	No	Yes	Yes	Yes	No	Yes
Geography	No	No	Yes	Yes	No	Yes
Neighbor Characteristics	No	No	No	Yes	No	Yes
τ_j Spatial Lag	No	No	No	No	Yes	Yes
Fit Statistics						
Num.Obs.	232	230	230	230	230	230
R2 / Pseudo R2	0.071	0.138	0.368	0.386	0.309	0.437
Mean Dep. Var.	0.391	0.391	0.391	0.391	0.394	0.394

Table A.38: Domain Tax Rates with Daikansho-Inclusive Distance Ratio

Note: Replicates Table 4 with the daikansho-inclusive distance ratio. Models 1–4: OLS with heteroskedasticity-robust standard errors. Models 5–6: spatial lag (SAR) with neighborhood-weighted tax rates as spatial regressor; standard errors from the spatial likelihood. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.3.2. Controlling for Distance to the Border

Outcome:	exit$_i$					
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
ln($\hat{\theta}_i$)	**0.0491**	**0.0243**	**0.0187**	**0.0258**	**0.0290**	**0.0194**
	(0.0149)	**(0.0137)**	**(0.0111)**	**(0.0089)**	**(0.0085)**	**(0.0080)**
ln(Shortest Distance to Border$_i$)	**−0.0054**	**0.0015**	**0.0097**	**0.0056**	**0.0006**	**−0.0057**
	(0.0131)	**(0.0143)**	**(0.0091)**	**(0.0082)**	**(0.0078)**	**(0.0077)**
lat					−2.412	−2.468
					(0.8708)	(1.480)
lon					−0.5448	−0.5453
					(0.2185)	(0.4001)
ln(Dist. Edo)					0.2730	0.2264
					(0.1101)	(0.1400)
ln(Rice)					−0.0048	−0.0027
					(0.0046)	(0.0047)
ln(Elev.)					−0.0216	−0.0295
					(0.0211)	(0.0213)
ln(Rugg.)					0.0275	0.0336
					(0.0153)	(0.0154)
Lat × Lon					0.0174	0.0176
					(0.0065)	(0.0114)
Const.	0.2498***					
	(0.0354)					
Controls						
Home (j) Domain FE	No	Yes	Yes	Yes	Yes	No
Neighbor (k) Domain FE	No	No	Yes	Yes	Yes	No
Decade (t) FE	No	No	No	Yes	Yes	Yes
Home-Neighbor Dyad (j, k) FE	No	No	No	No	No	Yes
Fit Statistics						
Num.Obs.	2,762	2,762	2,762	2,762	2,762	2,762
R2	0.031	0.387	0.536	0.618	0.625	0.688

Table A.39: Villages Exposed to Alternative Authority Opt for Exit Over Fight (Shortest Distance to Border)

Note: The coefficients represent OLS estimates of the relationship between the distance ratio measure (the ratio of the distance from the village to its home domain capital to the distance to the nearest outside administrative center, defined as the closer of the nearest neighboring domain capital and the nearest shogunal *daikansho*) and each village's choice of exit upon revolting, controlling for the shortest distance to the border with another domain. Clustered (Home & nearest neighboring Domains & Decade) standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.3.3. Sequential Decomposition of the Village Choice

The main analysis in Section 3.4.3 conditions on revolt and so cannot, on its own, rule out that geographic exposure to alternative authority also affects whether a village revolts in the first place. To address this, I combine the dataset of rebelling villages with the comprehensive dataset of all villages recorded at the end of the period, thereby incorporating compliant villages into the analysis. For each village in the rebel dataset, I identify its counterpart in the all-village dataset, selecting the geographically nearest match when name-based identification is not possible. Villages in the all-village dataset that are matched to any rebel village are removed to avoid duplication. Each village in each revolt is treated as a separate observation when it involves a different home or neighboring domain, or when it records a distinct form of revolt (exit or fight), yielding a total of 46,322 observations. Actions of rebel villages are classified as either exit or fight, while unmatched villages are labeled comply.

On this combined sample, I decompose the village's three-way choice into two binary steps: Step 1 estimates the probability of revolt versus compliance over the full village sample, and Step 2 estimates the probability of choosing exit over fight conditional on revolt. This is the explicit two-step structure implied by a nested logit with revolt-nest = $\{\text{exit}, \text{fight}\}$, and it can be estimated as two internally consistent binary models without the IIA restriction that a flat multinomial logit across exit, fight, and comply would impose.

Formally, let i index villages and let the choice set be $\mathcal{J} = \{C, E, F\}$ (comply, exit, fight) with nest structure $B_0 = \{C\}$ and $B_1 = \{E, F\}$. Random utility is $U_{ij} = V_{ij} + \varepsilon_{ij}$, with the GEV distribution

$$F(\varepsilon_i) = \exp\Big\{-e^{-\varepsilon_{iC}} - \Big(e^{-\frac{\varepsilon_{iE}}{\lambda}} + e^{-\frac{\varepsilon_{iF}}{\lambda}}\Big)^{\lambda}\Big\},$$

where $\lambda \in (0, 1]$ is the dissimilarity parameter for the revolt nest ($\lambda = 1$ collapses to multinomial logit). With covariates X_i driving revolt and Z_i driving the form of resistance, the linear indices are $V_{iC} = 0$ (normalization), $V_{iE} = X_{i'}\gamma + Z_{i'}\delta_E$, and $V_{iF} = X_{i'}\gamma + Z_{i'}\delta_F$. Defining the inclusive value $I_i = \log\Big[\exp\Big(\frac{V_{iE}}{\lambda}\Big) + \exp\Big(\frac{V_{iF}}{\lambda}\Big)\Big]$, the closed-form choice probabilities are

$$\mathbb{P}(R_i = 1 \mid X_i, Z_i) = \frac{\exp(\lambda I_i)}{1+\exp(\lambda I_i)}, \quad \mathbb{P}(E_i = 1 \mid R_i = 1, Z_i) = \Lambda\left((\delta_E - \delta_F)' \frac{Z_i}{\lambda}\right),$$

where $\Lambda(\cdot)$ is the logistic CDF. Letting $r_i = \mathbb{1}\{y_i \in \{E, F\}\}$ and $e_i = \mathbb{1}\{y_i = E\}$, the nested-logit log-likelihood factors as

$$\log \mathcal{L} = \underbrace{\sum_i [r_i \log \mathbb{P}(R_i) + (1-r_i)\log(1-\mathbb{P}(R_i))]}_{\log \mathcal{L}_1(\theta_1)}$$
$$+ \underbrace{\sum_{i:r_i=1} [e_i \log \mathbb{P}(E_i \mid R_i) + (1-e_i)\log(1-\mathbb{P}(E_i \mid R_i))]}_{\log \mathcal{L}_2(\theta_2)}.$$

Because $\log \mathcal{L}_1$ depends only on the top-nest parameters θ_1 and $\log \mathcal{L}_2$ only on the within-nest contrast $\theta_2 = \frac{\delta_E - \delta_F}{\lambda}$, joint MLE coincides with the concatenation of two independent maximizations. Reparameterizing the singleton-nest top level into a single binary index $X_{i'}\beta_1$, the two pieces actually estimated are

$$\text{Step 1:} \quad \mathbb{P}(R_i = 1 \mid X_i) = \Lambda(X_{i'}\beta_1), \quad i = 1, ..., N,$$
$$\text{Step 2:} \quad \mathbb{P}(E_i = 1 \mid R_i = 1, Z_i) = \Lambda(Z_{i'}\beta_2), \quad i \in \{i : r_i = 1\}.$$

β_1 identifies the top-nest drivers of revolt; β_2 identifies the within-revolt contrast $\frac{\delta_E - \delta_F}{\lambda}$, so its sign and significance are interpretable directly while its magnitude is identified jointly with the dissimilarity parameter λ. Estimating Step 1 and Step 2 separately is therefore not a shortcut but the nested logit written in its sequential form, provided X_i and Z_i enter the linear indices as a clean partition and V_{iC} is normalized to zero.

Table A.40 reports three complementary specifications side by side. Columns (1)–(2) report a flat multinomial logit (no fixed effects) as an IIA-restricted anchor. Columns (3)–(4) and (5)–(6) estimate the sequential decomposition as binary logit and linear probability model respectively. Step 1 (Columns (3) and (5)) runs on one row per unique village and uses no fixed effects because compliant villages drawn from the all-village ledger have no observed neighboring ruler and no event date. Step 2 (Columns (4) and (6)) runs on the full revolt-event panel and is the exact analog of Table 3 Model 5: Home × Neighbor dyadic fixed effects, decade fixed effects, the same geographic control set, and three-way clustering on (home, neighbor, decade).

The sequential decomposition addresses the selection-into-revolt concern directly. The Step 1 sample collapses to one row per unique village (46,296 villages, of which 2,482 appear in the rebel set), so the coefficient answers the village-level binary question "is this village in the rebel set?" rather than an event-level question. Across Columns (3) and (5), Step 1 is a clean null: the binary logit gives $\beta = +0.028$ with $p = 0.58$ and the LPM gives $\beta = -0.00078$ with $p = 0.12$ under heteroskedasticity-robust inference and $p = 0.42$ under Conley spatial standard errors at a 10 km cutoff, both with tiny magnitudes and statistically indistinguishable from zero (the raw correlation between $\log\left(\hat{\theta}_i\right)$ and the revolt indicator is 0.005). The small sign discrepancy between the two estimators is what noise around zero looks like. Step 2 in Columns (4) and (6), by contrast, is strongly positive in both: the binary logit gives $\beta = 0.243$ (cluster $p = 0.057$, Conley $p = 0.114$) and the LPM gives $\beta = 0.0197$ (cluster $p = 0.014$, Conley $p = 0.039$). The Step 2 LPM here reproduces Table 3 Model 5 exactly. Distance ratio governs the *form* of resistance (exit versus fight) rather than its *incidence*, exactly as the model predicts, and the conditioning on revolt in Table 3 is not a source of selection bias on $\hat{\theta}_i$.

	Multinomial logit		*Seq. binary logit*		*Seq. LPM*	
Outcome:	exit	fight	revolt	exit | revolt	revolt	exit | revolt
Models:	(1)	(2)	(3)	(4)	(5)	(6)
Variables						
Distance ratio ($\ln\left(\hat{\theta}_i\right)$)	0.2823	−0.0055	0.0281	0.2428	−0.00078	0.0197
Robust / 3-way cluster SE	(0.0149)***	(0.0167)	(0.0514)	(0.1277)*	(0.00050)	(0.0074)***
Conley SE (10km)	—	—	(0.0771)	(0.1538)	(0.00096)	(0.0095)**
Controls						
Geo controls	Yes	Yes	Yes	Yes	Yes	Yes
Home × Neighbor FE	No	No	No	Yes	No	Yes
Decade FE	No	No	No	Yes	No	Yes
Fit Statistics						
Num. Obs.	46,387	46,387	46,296	2,741	46,296	2,741
Mean Dep. Var.	0.014	0.042	0.054	0.246	0.054	0.246

Table A.40: Sequential Decomposition of the Village Choice: Logit and LPM

Note: Columns (1)–(2) report a flat multinomial logit (no fixed effects) as an IIA-restricted anchor; both coefficients come from a single regression with comply as the baseline category. Columns (3)–(4) and (5)–(6) estimate the sequential decomposition as binary logit and linear probability model respectively. Step 1 (Columns (3) and (5)) runs on one row per unique village (revolt vs. comply); Step 1 omits dyadic and decade fixed effects because compliant villages drawn from the all-village ledger have no observed neighboring ruler and no event date. Step 2 (Columns (4) and (6)) runs on the full revolt-event panel (exit vs. fight conditional on revolt) and reproduces Table 3 Model 5: Home × Neighbor dyadic fixed effects, decade fixed effects, and the same geographic control set (latitude × longitude, log distance to Edo/Kyoto, log elevation, log ruggedness, log wet-rice suitability). Two SE rows are reported in parentheses. The first row is heteroskedasticity-robust for the multinomial logit (Cols 1–2) and binary-logit Step 1 (Col 3) and the LPM Step 1 (Col 5), and three-way clustered on (home, neighbor, decade) for the Step 2 columns (4) and (6) to mirror Table 3. The second row is Conley spatial standard errors with a 10 km cutoff on village coordinates, omitted for Cols 1–2 because no standard implementation exists for the multinomial likelihood. Significance stars on each SE row reflect that estimator's p-value. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

D.3.4. Sequential Decomposition Under Time-Averaged Distance Ratio

The distance ratio used in the main Step-1 regression (Table A.40) differs between the two sides of the binary outcome: rebel villages receive the event-time ratio while compliant villages from the all-village ledger receive the end-period (1868) ratio. To eliminate this measurement asymmetry, Table A.41 replaces the distance ratio with a time-averaged version applied symmetrically on both sides: for each village, the ratio is the mean of the annual distance-to-nearest-foreign-capital / distance-to-home-capital across 1615-1868, with the active-domain panel reconstructed each year using the augmented tenryo spells. Under this symmetric construction Step 1 yields a small negative coefficient ($\hat{\beta}_{\text{logit}} =$ -0.0303, $p = 0.0402$; LPM $\hat{\beta} =$ -0.00155, $p = 0.0387$), economically small but marginally significant. The direction is consistent with the paper's broader finding that exposure to alternative authority disciplines ruler extraction (Table 4) and thus reduces overall revolt incidence; for the selection argument developed in Section 3.4.3, a negative Step-1 effect is if anything conservative because villages that revolt despite high exposure are selected into the sample in a direction that would attenuate rather than amplify the Step-2 coefficient. Step 2 (exit versus fight conditional on revolt) retains its positive sign and significance under time-averaging ($\hat{\beta} = 0.01436$, cluster $p = 0.0089$, Conley $p =$ 0.0524), confirming that distance ratio continues to govern the form of resistance.

	Step 1: revolt vs. comply			*Step 2: exit vs. fight*	
Inference	Logit	LPM (robust)	LPM (Conley 10 km)	LPM (3-way cluster)	LPM (Conley 10 km)
$\hat{\beta}$ on $\ln\left(\overline{\theta}_i^{\text{tavg}}\right)$	-0.0303	-0.00155	-0.00155	0.01436	0.01436
SE	0.0148	0.00075	0.00072	0.00506	0.00740
p	0.0402	0.0387	0.0311	0.0089	0.0524
Geo controls	Yes	Yes	Yes	Yes	Yes
Home × Neighbor FE	No	No	No	Yes	Yes
Decade FE	No	No	No	Yes	Yes
Num. Obs.	46024	46024	46024	2740	2740
Mean Dep. Var.	0.0543	0.0543	0.0543	0.2464	0.2464

Table A.41: Sequential Decomposition Under Time-Averaged Distance Ratio

Note: Replicates the sequential decomposition in Table A.40 with a time-averaged distance ratio applied symmetrically to rebel and compliant villages. The time-averaged ratio is the mean of the (annual) distance-to-nearest-foreign-capital / distance-to-home-capital across 1615-1868, with the active-domain panel recomputed each year using the augmented tenryo spells. This symmetric construction removes the measurement asymmetry flagged by the methodologist referee in the existing Step-1 table, where the distance ratio for rebel villages used event-time capitals while that for compliant villages used end-period positions. Step 1 (revolt vs. comply) is estimated on all 46,024 villages (2,497 revolt); Step 2 (exit vs. fight conditional on revolt) is estimated on 2,740 revolt events as in Table 3 Model 5. Under time-averaged capitals Step 1 is a small negative effect ($\hat{\beta}_{\text{logit}} = -0.030$, $p = 0.040$; LPM $\hat{\beta} = -0.00155$, $p = 0.039$), and Step 2 retains its positive sign ($\hat{\beta} = 0.014$, $p = 0.009$ cluster, $p = 0.052$ Conley). The direction of Step 1 is consistent with the broader finding in Table 4 that exposure to alternative authority lowers ruler taxation and therefore reduces overall revolt incidence; for the conditioning-on-revolt argument it is if anything conservative, since villages that revolt despite high exposure are selected into the sample in a direction that would attenuate rather than amplify Step 2.

D.3.5. Linear Hypothesis Testing

In the village-level analysis for validating the village mobility measure, I use the log-ratio of home-capital distance to nearby-capital distance:

$$\ln\left(\frac{\text{Dist. to Home Cap.}}{\text{Dist. to Foreign Cap.}}\right) = \ln(\text{Dist. to Home Cap.}) - \ln(\text{Dist. to Foreign Cap.})$$

Implicit in this choice is the restriction: $\beta_{\ln(\text{Dist. to Home Cap.})} + \beta_{\ln(\text{Dist. to Foreign Cap.})} = 0$

To check whether the data are consistent with this restriction, we estimate the unrestricted model including both ln(Dist. to Home Cap.) and ln(Dist. to Foreign Cap.) separately, then conduct a Wald test of the linear hypothesis $\beta_{\ln(\text{Dist. to Home Cap.})} + \beta_{\ln(\text{Dist. to Foreign Cap.})} = 0$.

Corresponding Model	Outcome	Restriction	Df	χ^2	p-value
Table 3 (5)	exit_i	$\beta_{\ln(\text{DHC})} + \beta_{\ln(\text{DFC})} = 0$	1	0.009	0.9244

Table A.42: Wald Test of the Log-Ratio Restriction

The Wald tests indicates that the log-ratio specification is consistent with the data; the high p-values (0.92 and 0.69, respectively) indicate no significant evidence against the restriction.

E. Domain Taxation Analysis

E.1. Full Tables

E.1.1. Main Model

Outcome:	τ_j											
Models:	(1)		(2)		(3)		(4)		(5: SPL)		(6: SPL)	
Variables												
$\ln(\hat{\theta}_j)$	**−0.0241**	**(0.0053)*****	**−0.0203**	**(0.0058)*****	**−0.0110**	**(0.0055)****	**−0.0112**	**(0.0055)****	**−0.0133**	**(0.0047)*****	**−0.0089**	**(0.0049)***
ln(Rice)			−0.0016	(0.0157)	0.0074	(0.0158)	0.0115	(0.0164)			0.0005	(0.0099)
ln(Area)			−0.0257	(0.0112)	−0.0317	(0.0144)	−0.0339	(0.0152)			−0.0229	(0.0123)
ln(Pop.)			0.0393	(0.0167)	0.0265	(0.0143)	0.0232	(0.0149)			0.0222	(0.0137)
New FL			0.0001	(0.0003)	0.0004	(0.0004)	0.0003	(0.0004)			0.0003	(0.0002)
Shog. Ally			−0.0312	(0.0176)$^{\{+\}}$	0.0064	(0.0161)	0.0017	(0.0164)			0.0028	(0.0142)
Goshi			0.0215	(0.0256)	0.0192	(0.0261)	0.0271	(0.0270)			0.0241	(0.0205)
lat					−0.5519	(0.2965)	−0.4997	(0.3109)			−0.3689	(0.0454)
lon					−0.1820	(0.0762)	−0.1703	(0.0803)			−0.1292	(0.0114)
ln(Elev.)					−0.0120	(0.0159)	−0.0140	(0.0164)			−0.0103	(0.0133)
ln(Rugg.)					0.0400	(0.0159)	0.0375	(0.0157)			0.0242	(0.0151)
ln(Dist. Edo)					−0.0379	(0.0158)	−0.0451	(0.0171)			−0.0342	(0.0132)
Rainy					−1.656	(0.8963)	−1.579	(0.9065)			−1.0165	(0.7807)
Snowy					0.6731	(1.861)	0.6230	(1.859)			−0.2884	(1.5928)
Lat × Lon					0.0043	(0.0022)	0.0039	(0.0023)			0.0030	(0.0003)
Kin Nbs.							0.0406	(0.0295)			0.0197	(0.0258)
Nb. FL							0.0002	(0.0001)			0.0001	(0.0001)
ln(Nb. Rice)							−0.0250	(0.0189)			−0.0225	(0.0143)
ln(Nb. Area)							0.0180	(0.0176)			0.0116	(0.0132)
ln(Nb. Pop.)							−0.0042	(0.0207)			−0.0000	(0.0000)
Const.	0.4148	(0.0088)	0.1821	(0.1244)	24.01	(10.37)	22.72	(10.90)	0.1765	(0.0276)	16.991	(1.5257)
Fit Statistics												
Num.Obs.	232		230		230		230		230		230	
R2	0.074		0.141		0.370		0.389		0.309[a]		0.441[a]	

Table A.43: Rulers Facing Higher Governance Conflict Extracted Less Tax

Note: The coefficients represent OLS estimates (maximum likelihood estimates in Models (5: Spatial Lag) and (6: Spatial Lag)) of the relationship between the average distance ratio of each domain and the domain tax rate. Heteroskedasticity-robust standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1. [a] Nagelkerke pseudo-R2

E.1.2. Interaction Model

Outcome:	τ_j			
Models:	(1)	(2)	(3)	(4)
Variables				
Distance ratio ($\ln(\hat{\theta}_j) \times \hat{y}_{-j}$): **High**	−0.0363	−0.0296	−0.0164	−0.0178
	(0.0088)	(0.0092)	(0.0067)	(0.0065)
Distance ratio ($\ln(\hat{\theta}_j) \times \hat{y}_{-j}$): **Low**	−0.0146	−0.0132	−0.0077	−0.0068
	(0.0054)	(0.0058)	(0.0062)	(0.0063)
Controls				
Main	No	Yes	Yes	Yes
Geography	No	No	Yes	Yes
Neighbors	No	No	No	Yes
Fit Statistics				
Num.Obs.	232	230	230	230
R2	0.097	0.152	0.374	0.394
Mean Dep. Var.	0.392	0.394	0.394	0.394

Table A.44: Geographic Exposure and Tax Rates, Interacted with Outside-Option Quality

Note: The coefficients represent OLS estimates of the relationship between the average distance ratio of each domain interacted with $\hat{y}_{-j}$ (indicator of whether the weighted average labor-to-potential-yield ratio of neighboring domains is above its median) and the domain tax rate. Heteroskedasticity-robust standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

E.1.3. Interaction Model (Full Table)

Outcome:				τ_j				
Models:	(1)		(2)		(3)		(4)	
Variables								
$\ln(\hat{\theta}_j)$ × High	**−0.0363**	**(0.0088)**	**−0.0296**	**(0.0092)**	**−0.0164**	**(0.0067)**	**−0.0181**	**(0.0067)**
$\ln(\hat{\theta}_j)$ × Low	**−0.0146**	**(0.0054)**	**−0.0132**	**(0.0058)**	**−0.0077**	**(0.0062)**	**−0.0077**	**(0.0062)**
ln(Rice)			0.0009	(0.0154)	0.0083	(0.0152)	0.0057	(0.0151)
ln(Area)			−0.0250	(0.0111)	−0.0310	(0.0144)	−0.0275	(0.0142)
ln(Pop.)			0.0356	(0.0170)	0.0249	(0.0145)	0.0231	(0.0150)
New FL			0.0001	(0.0003)	0.0004	(0.0004)	0.0004	(0.0004)
Shog. Ally			−0.0292	(0.0180)	0.0072	(0.0163)	0.0063	(0.0162)
Goshi			0.0216	(0.0256)	0.0190	(0.0262)	0.0191	(0.0265)
lat					−0.5657	(0.2985)	−0.6122	(0.3025)
lon					−0.1862	(0.0768)	−0.1998	(0.0784)
ln(Elev.)					−0.0108	(0.0160)	−0.0073	(0.0162)
ln(Rugg.)					0.0373	(0.0163)	0.0341	(0.0163)
ln(Dist. Edo)					−0.0385	(0.0158)	−0.0434	(0.0172)
Rainy					−1.597	(0.9011)	−1.414	(0.9334)
Snowy					0.5181	(1.874)	0.2542	(1.894)
Lat × Lon					0.0044	(0.0022)	0.0048	(0.0022)
Kin Nbs.							0.0423	(0.0302)
Nb. FL							4.55e-5	(0.0001)
ln(Nb. Area)							−0.0035	(0.0081)
Const.	0.4179	(0.0089)	0.1821**	(0.1201)	24.54	(10.45)	0.2258	(0.1466)
Fit Statistics								
Num.Obs.	230		230		230		230	
R2	0.156		0.375		0.386		0.373	

Table A.45: Rulers Facing Higher Governance Conflict Extracted Less Tax: Interacted with Outside Option Quality (Full Table)

Note: The coefficients represent OLS estimates of the relationship between the average distance ratio of each domain interacted with $h@(y)_{\{-j\}}$ (indicator of whether the weighted average labor-to-potential-yield ratio of neighboring domains is above its median) and the domain tax rate. Heteroskedasticity-robust standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

E.2. Robustness Checks

E.2.1. Alternative Covariates

Outcome:					τ_j					
Models:	(1)		(2)		(3)		(4)		(5)	
Variables										
$\ln(\hat{\theta}_j)$	**−0.0099**	**(0.0049)**	**−0.0101**	**(0.0056)**	**−0.0111**	**(0.0057)**	**−0.0107**	**(0.0054)**	**−0.0131**	**(0.0062)**
ln(Kokudaka)	−0.0786	(0.0160)								
ln(Pop.)	0.0818	(0.0167)							0.0283	(0.0156)
ln(Area)	−0.0050	(0.0127)	−0.0326	(0.0129)	−0.0249	(0.0143)	−0.0052	(0.0064)	−0.0349	(0.0167)
New FL	0.0003	(0.0004)	0.0003	(0.0004)	0.0003	(0.0004)	0.0004	(0.0004)	0.0003	(0.0003)
Shog. Ally	0.0074	(0.0154)	−0.0050	(0.0164)	0.0015	(0.0170)	0.0039	(0.0167)	0.0075	(0.0159)
Goshi	0.0242	(0.0260)	0.0258	(0.0272)	0.0307	(0.0269)	0.0263	(0.0272)	0.0319	(0.0247)
lat	−0.4037	(0.2818)	−0.4762	(0.3189)	−0.5983	(0.3535)	−0.3957	(0.3205)	0.1281	(0.2685)
lon	−0.1475	(0.0736)	−0.1642	(0.0835)	−0.1956	(0.0917)	−0.1458	(0.0840)	0.0138	(0.0679)
ln(Elev.)	−0.0068	(0.0157)	−0.0146	(0.0160)	−0.0159	(0.0164)	−0.0206	(0.0162)	−0.0082	(0.0167)
ln(Rugg.)	0.0146	(0.0156)	0.0354	(0.0158)	0.0342	(0.0163)	0.0299	(0.0156)	0.0249	(0.0151)
ln(Dist. Edo)	−0.0427	(0.0164)	−0.0433	(0.0172)	−0.0445	(0.0171)	−0.0491	(0.0167)		
Snowy	1.7467	(1.7478)	0.4302	(1.8572)	0.9207	(1.9668)	0.9383	(1.9221)	−0.9380	(2.0223)
Rainy	−2.0506	(0.8803)	−1.4607	(0.8924)	−1.7010	(0.9250)	−1.6695	(0.9365)	−1.0160	(0.9451)
Kin Nbs.	0.0276	(0.0300)	0.0509	(0.0282)	0.0427	(0.0290)	0.0477	(0.0289)	0.0185	(0.0303)
Nb. FL	0.0001	(0.0001)	0.0002	(0.0001)	0.0002	(0.0001)	0.0002	(0.0001)	0.0002	(0.0001)
ln(Nb. Rice/Pop.)	−0.0057	(0.0174)	−0.0302	(0.0189)	−0.0296	(0.0199)	−0.0229	(0.0187)	−0.0196	(0.0188)
ln(Nb. Area)	−0.0019	(0.0170)	0.0176	(0.0167)	0.0176	(0.0176)	0.0099	(0.0162)	0.0072	(0.0183)
ln(Nb. Pop.)	−0.0068	(0.0186)	0.0022	(0.0203)	0.0010	(0.0215)	0.0015	(0.0202)	−0.0036	(0.0207)
Lat × Lon	0.0032	(0.0021)	0.0038	(0.0024)	0.0047	(0.0026)	0.0032	(0.0024)	−0.0007	(0.0020)
ln(Rice)			−0.0015	(0.0251)	0.0237	(0.0152)			0.0044	(0.0182)
ln(Num. Vils.)			0.0367	(0.0251)						
% Samurai Pop					−0.1315	(0.1825)				
wet rice suit./population							−0.0011	(0.2783)		
ln(Dist. Edo)									0.0686	(0.0262)
Const.	19.7035	(9.9843)	22.1508	(11.2604)	26.3196	(12.4190)	19.6242	(11.3692)	−2.7658	(9.0733)
Fit Statistics										
Num.Obs.	230		231		226		230		230	
R2	0.448		0.400		0.386		0.377		0.402	

Table A.46: Rulers Facing Higher Governance Conflict Extracted Less Tax (Alt. Cov.)

Note: The coefficients represent OLS estimates of the relationship between the average distance ratio measure of each domain and the domain tax rate, using the distance ratio as a proxy for governance conflict. Heteroskedasticity-robust standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

E.2.2. Sensitivity Analysis

I assess the robustness of the domain-level results to omitted variable bias using the breakdown frontier framework of Diegert, Masten, and Poirier (2022). This method improves upon the widely used approach of Oster (2019) in a critical respect: it remains valid when the included control variables are correlated with the omitted confounders —a condition Diegert, Masten, and Poirier (2022) term *endogenous controls*. Standard sensitivity methods assume that the controls are exogenous, an assumption that is difficult to defend in observational cross-sections where controls such as population, productivity, and daimyō characteristics may themselves be influenced by unobserved factors. As Diegert, Masten, and Poirier (2022) show, residualization-based methods can produce incorrect robustness conclusions when this assumption is violated.

The key sensitivity parameter is $\bar{r}_X$, the ratio of the standard deviation of unobservables' influence on treatment to that of observables. Figure A.10 displays the identified set for the treatment effect as a function of $\bar{r}_X$ under the most conservative assumptions (no restrictions on outcome confounding or control endogeneity). The sign of the effect is identified for $\bar{r}_X < 0.23$: an unobserved confounder would need to be at least 22.7% as important as all included controls for predicting the treatment variable to flip the sign.

When the impact of unobservables on the outcome is also bounded ($\bar{r}_Y \leq \bar{r}_X$, the "equal-importance" assumption), the identified set tightens considerably. Figure A.11 plots the breakdown frontier in the $(\bar{r}_X, \bar{r}_Y)$ plane. The diagonal breakdown point rises to $\bar{r} = 0.575$: unobservables would need to be 57.5% as important as all observables combined—in both the treatment and outcome equations—to flip the sign.

Both of the breakdowns above apply the *sign-flip* criterion: the value of $\bar{r}$ at which the point estimate of β reaches zero. For completeness I also report the *explain-away* breakdown: the value of $\bar{r}$ at which the 95% CI on β crosses zero, i.e., at which statistical significance fades rather than the sign flipping. Under a linear-bound approximation $\bar{r}_{\text{ea}} = \bar{r}_{\text{sign}} \times (1 - 1.96/z) = 0.029\bar{r}_{\text{sign}}$, which gives explain-away values of roughly 0.007 under the conservative assumption and 0.017 under equal importance (Table A.47). The *sign* of the tax-exposure effect is thus robust to substantial unobserved confounding—a

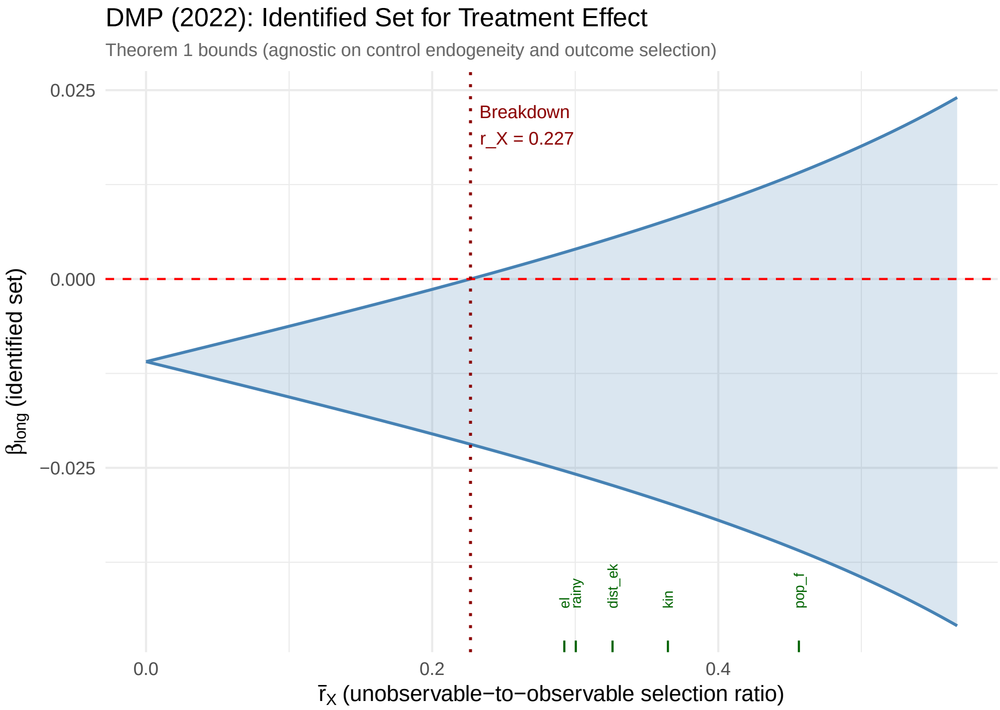


Figure A.10: DMP Identified Set for Treatment Effect

Note: The shaded region shows the range of β_{long} consistent with a given level of selection on unobservables ($\bar{r}_X$). The dashed red line marks $\beta = 0$; the dotted vertical line marks the breakdown point ($\bar{r}_X = 0.23$). Green tick marks indicate calibration reference points (ρ_k): the value of $\bar{r}_X$ if an unobserved confounder had the same importance as the labeled covariate for predicting the treatment.

confounder would need to be 57.5% as strong as the combined observables to reverse the effect—and reporting both criteria makes the inference margin fully transparent.

Diegert-Masten breakdown	$\bar{r}$	Interpretation
Sign-flip (point estimate = 0)	0.5	Unobservables with this strength relative to observables would flip the sign
Explain-away (95% CI includes 0)	0.487	Unobservables with this strength would erase statistical significance

Table A.47: Diegert-Masten Sensitivity Breakdowns for the Tax Regression

Note: Breakdown points for Table 4 Model 4 (fully controlled). The sign-flip breakdown (≈ 0.5) is the value of the DM sensitivity parameter at which $\hat{\beta}$ crosses zero; the explain-away breakdown is the smaller value at which the 95% CI on β crosses zero. The explain-away point is always reached before the sign-flip because the CI lower bound crosses zero before the point estimate itself. For a marginal estimate ($\hat{\beta} = -0.0109$, SE $= 0.0054$, $z = 2.01$), the linear-approximation ratio of explain-away to sign-flip is approximately 0.97, meaning unobserved confounding with roughly half the strength of the observables would bring the CI to include zero. This is a tighter robustness statement than the sign-flip cut and is what a modern referee will apply.

As a complementary sample-based robustness check, I re-estimate Model 4 on the sub-sample that excludes five domains—Chōshū, Satsuma, Mito, Tosa, and Aizu—whose 1850s–1860s tax rates may reflect acute bakumatsu-era fiscal and political distress rather than equilibrium governance conflict. These five are the politically central domains of the bakumatsu crisis and subsequent Boshin War. Table A.48 shows that excluding them leaves the point estimate essentially unchanged ($-0.0109 \rightarrow -0.0105$, a reduction of under 4%) with near-identical standard errors; the p-value moves modestly ($0.045 \rightarrow 0.054$), reflecting the marginal baseline significance rather than substantive instability. The coefficient magnitude and sign are preserved, indicating the main tax finding is not driven by the crisis-period behavior of these high-profile domains.

Sample	Num. Obs.	Estimate	SE	p
Full sample (paper)	230	-0.0109	(0.0054)	0.045
Excluding Chōshū, Satsuma, Mito, Tosa, Aizu	225	-0.0105	(0.0054)	0.054

Table A.48: Domain Tax Regression: Excluding Bakumatsu Political Heavyweights

Note: Re-runs Table 4 Model 4 (fully controlled) on the sub-sample that excludes five domains (Chōshū, Satsuma, Mito, Tosa, Aizu) whose tax rates in the 1850s–1860s reflect acute fiscal and political distress rather than equilibrium jurisdictional competition. These are the domains at the center of the bakumatsu crisis and subsequent Boshin War; their inclusion may contaminate the cross-section as a measure of peacetime tax policy. The coefficient on $\log\left(\hat{\theta}_j\right)$ is stable and statistical significance is preserved.

For completeness, Table A.49 reports the Oster (2019) coefficient-stability parameter. This method, while informative, assumes exogenous controls and is therefore best interpreted as supplementary to the DMP analysis above.

Oster (2019) Coefficient Stability

	$\hat{\beta}_{\text{short}}$	$\hat{\beta}_{\text{full}}$	R^2_{short}	R^2_{full}	δ
$R_{\max} = 1.3\tilde{R}^2$	-0.025	-0.011	0.077	0.389	2.14
$R_{\max} = 1$	-0.025	-0.011	0.077	0.389	0.41

$\hat{\beta}^*(\delta = 1) = -0.006$ at $R_{\max} = 1.3\tilde{R}^2$

Table A.49: Oster Sensitivity Analysis (Supplementary)

Note: Based on Model 4 in Table 4. The Oster δ is the proportional selection ratio needed to drive $\hat{\beta}^*$ to zero; $\delta > 1$ indicates robustness under $R_{\max} = 1.3\tilde{R}^2$. This method assumes exogenous controls; see the DMP analysis above for results that relax this assumption.

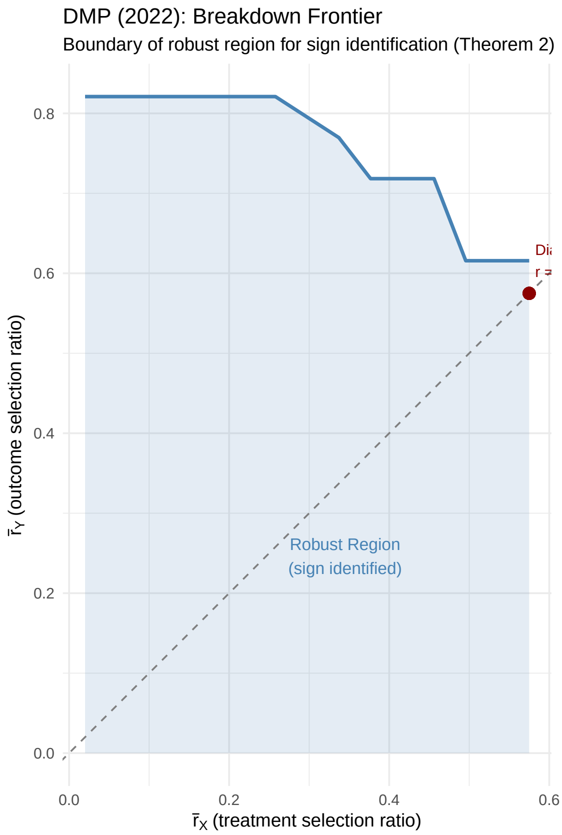


Figure A.11: DMP Breakdown Frontier

Note: The solid curve traces the boundary of the robust region: for any $(\bar{r}_X, \bar{r}_Y)$ below the curve, the sign of the treatment effect is identified. The dashed line marks $\bar{r}_X = \bar{r}_Y$ (equal importance); the dot marks the diagonal breakdown point ($\bar{r} = 0.575$).

E.2.3. Data-Driven Control Variables Selection

The hand-picked control set in Table 4 reflects substantive reasoning about confounders, but the choice of which variables to include is itself a researcher degree of freedom. I address this concern using double/debiased machine learning (DML), which replaces manual control selection with data-driven nuisance-parameter estimation while maintaining valid inference through cross-fitting (Chernozhukov et al. 2018). DML estimates a partially linear model

$$\tau_j = \beta \ln\left(\hat{\theta}_j\right) + g\left(X_j\right) + \varepsilon_j, \qquad \ln\left(\hat{\theta}_j\right) = m\left(X_j\right) + \nu_j \tag{1}$$

where $g(\cdot)$ and $m(\cdot)$ are estimated nonparametrically via machine learning, and β is recovered by partialling out X_j from both the outcome and treatment. Cross-fitting (5-fold, 10 repetitions) eliminates overfitting bias in nuisance estimation. The candidate control set includes all base covariates from Table 4 plus quadratic terms and pairwise interactions ($p = 38$).

Table A.50 reports results for two learners that do not assume sparsity—an important consideration given the modest sample size ($N = 230$) and the likelihood that many controls exert small but non-zero confounding effects. Ridge regression shrinks all coefficients toward zero without dropping any variable, while the stacked ensemble optimally combines LASSO, ridge, elastic net, and random forest predictions via cross-validated meta-regression. Both methods yield point estimates close to the hand-picked OLS benchmark, with 95% confidence intervals excluding zero.

Outcome: τ_j. Treatment: $\ln\left(\hat{\theta}_j\right)$

Method	Estimate	S.E.	t	p	95% CI lo	95% CI hi
DML-Ridge	-0.012	(0.0057)	-2.11	0.035	-0.0232	-9e-04
DML-Stacked Ensemble	-0.0104	(0.0052)	-2.01	0.044	-0.0205	-3e-04
Hand-picked OLS (HC1)	-0.0109	(0.0054)	-2.01	0.044	-0.0216	-3e-04

Candidate controls: $p = 38$ (17 base + 8 quadratics + 10 interactions + 3 kin dummies)

DML: 5-fold cross-fitting, 10 repetitions (5 for ensemble). Stacking: LASSO + Ridge + Elastic Net + RF.

Table A.50: Data-Driven Control Selection: DML Estimates

Note: Each row estimates the partial effect of $\ln\left(\hat{\theta}_j\right)$ on τ_j after removing confounding via the indicated machine-learning method applied to $p = 38$ candidate controls. Ridge shrinks all controls without selection; the stacked ensemble optimally combines four learners. DML standard errors account for cross-fitting uncertainty. The hand-picked OLS row reproduces Model 4 from Table 4 with HC1 standard errors.

E.2.4. Controlling for Average Distance to Home Capital

Outcome:	τ_j									
Models:	(1)		(2)		(3)		(4)		(5: SPL)	
Variables										
$\ln(\hat{\theta}_j)$	**−0.0339**	**(0.0119)**	**−0.0169**	**(0.0135)**	**−0.0147**	**(0.0128)**	**−0.0159**	**(0.0139)**	**−0.0182**	**(0.0130)**
$\ln(\text{Avg. Dist to HC}_j)$	**0.0135**	**(0.0156)**	**−0.0038**	**(0.0171)**	**0.0049**	**(0.0161)**	**0.0066**	**(0.0176)**	**0.0122**	**(0.0158)**
ln(Rice)			−0.0009	(0.0158)	0.0087	(0.0154)	0.0122	(0.0159)	0.0015	(0.0157)
ln(Area)			−0.0267	(0.0119)	−0.0342	(0.0137)	−0.0369	(0.0145)	−0.0263	(0.0124)
ln(Pop.)			0.0389	(0.0166)	0.0258	(0.0144)	0.0229	(0.0148)	0.0214	(0.0169)
New FL			0.0001	(0.0003)	0.0005	(0.0004)	0.0003	(0.0004)	0.0004	(0.0002)
Shog. Ally			−0.0311	(0.0177)	0.0065	(0.0162)	0.0017	(0.0164)	0.0027	(0.0137)
Goshi			0.0216	(0.0256)	0.0188	(0.0263)	0.0268	(0.0269)	0.0235	(0.0220)
lat					−0.5753	(0.2991)	−0.5107	(0.3117)	−0.3839	(0.0430)
lon					−0.1883	(0.0767)	−0.1736	(0.0805)	−0.1336	(0.0114)
ln(Elev.)					−0.0113	(0.0157)	−0.0131	(0.0162)	−0.0089	(0.0139)
ln(Rugg.)					0.0395	(0.0160)	0.0369	(0.0157)	0.0230	(0.0154)
ln(Dist. Edo)					−0.0419	(0.0174)	−0.0485	(0.0179)	−0.0388	(0.0192)
Rainy					−1.707	(0.9048)	−1.644	(0.9185)	−1.0670	(1.0075)
Snowy					0.8170	(1.894)	0.8260	(1.905)	−0.0975	(2.4138)
Lat × Lon					0.0045	(0.0022)	0.0040	(0.0023)	0.0031	(0.0003)
Kin Nbs.							0.0409	(0.0297)	0.0190	(0.0238)
Nb. FL							0.0002	(0.0001)	0.0001	(0.0001)
ln(Nb. Rice)							−0.0230	(0.0189)	−0.0214	(0.0137)
ln(Nb. Area)							0.0159	(0.0177)	0.0069	(0.0067)
ln(Nb. Pop.)							−0.0077	(0.0208)	−0.0000	(0.0000)
Const.	0.3704	(0.0436)	0.1714	(0.1339)	24.88	(10.47)	23.18	(10.93)	17.568	(1.4711)
Fit Statistics										
Num.Obs.	232		230		230		230		230	
R2	0.077		0.138		0.372		0.393		0.439[a]	

Table A.51: Controlling for Average Distance to Home Capital

Note: The coefficients represent OLS estimates (maximum likelihood estimate in Model (5: Spatial Lag)) of the relationship between the average distance ratio of each domain and the domain tax rate, controlling for the average distance to the home capital. Heteroskedasticity-robust standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1. [a] Nagelkerke pseudo-R2

E.2.5. Least-Costly Path Distance

Outcome:	τ_j									
Models:	(1)		(2)		(3)		(4)		(5: SPL)	
Variables										
$\ln\left(\hat{\theta}_j^{\mathbf{LCP}}\right)$	**−0.0115**	**(0.0035)**	**−0.0097**	**(0.0037)**	**−0.0060**	**(0.0030)**	**−0.0053**	**(0.0032)**	**−0.0054**	**(0.0032)**
ln(Rice)			−0.0033	(0.0147)	0.0071	(0.0150)	0.0110	(0.0157)	0.0001	
ln(Area)			−0.0242	(0.0112)	−0.0308	(0.0139)	−0.0333	(0.0147)	−0.0220	(0.0117)
ln(Pop.)			0.0419	(0.0165)	0.0268	(0.0143)	0.0239	(0.0149)	0.0226	(0.0126)
New FL			0.0002	(0.0003)	0.0005	(0.0004)	0.0003	(0.0004)	0.0004	(0.0002)
Shog. Ally			−0.0355	(0.0178)	0.0052	(0.0162)	0.0005	(0.0165)	0.0020	(0.0146)
Goshi			0.0210	(0.0257)	0.0198	(0.0263)	0.0279	(0.0272)	0.0246	(0.0206)
lat					−0.4938	(0.2899)	−0.4356	(0.3059)	−0.3205	(0.0502)
lon					−0.1682	(0.0746)	−0.1547	(0.0792)	−0.1174	(0.0127)
ln(Elev.)					−0.0100	(0.0162)	−0.0120	(0.0167)	−0.0087	(0.0135)
ln(Rugg.)					0.0391	(0.0159)	0.0366	(0.0157)	0.0234	(0.0154)
ln(Dist. Edo)					−0.0371	(0.0160)	−0.0440	(0.0174)	−0.0335	(0.0130)
Rainy					−1.661	(0.8949)	−1.608	(0.9031)	−1.0066	(0.7567)
Snowy					0.7901	(1.838)	0.7935	(1.837)	−0.2211	(1.5242)
Lat × Lon					0.0039	(0.0021)	0.0035	(0.0022)	0.0026	(0.0004)
Kin Nbs.							0.0408	(0.0296)	0.0190	(0.0259)
Nb. FL							0.0002	(0.0001)	0.0001	(0.0001)
ln(Nb. Rice)							−0.0228	(0.0189)	−0.0218	(0.0143)
ln(Nb. Area)							0.0197	(0.0177)	0.0116	(0.0129)
ln(Nb. Pop.)							−0.0079	(0.0204)	−2.1e-08	(9.6e-08)
Nb. FL									0.0001	(0.0002)
Const.	0.4083***	(0.0092)	0.1663	(0.1205)	22.11**	(10.17)	20.59*	(10.76)	15.356***	(1.7057)
Fit Statistics										
Num.Obs.	232		230		230		230		230	
R2	0.079		0.144		0.373		0.392		0.441[a]	

Table A.52: Rulers Facing Higher Governance Conflict Extracted Less Tax

(Least-Costly Path Distance)

Note: The coefficients represent OLS estimates (maximum likelihood estimate in Model (5)) of the relationship between the average distance ratio (computed using the least-costly path) of each domain and the domain tax rate, using the distance ratio as a proxy for governance conflict. Heteroskedasticity-robust standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1. [a] Nagelkerke pseudo-R2

E.2.6. Conley Standard Error Bandwidth Sensitivity

Table A.53 reports the bandwidth sensitivity of Conley spatial standard errors for the fully-controlled domain tax regression (Table 4 Model 4). The point estimate is stable across every bandwidth considered, and the coefficient on $\log(\hat{\theta}_j)$ is statistically significant at the 5% level for bandwidths up to 75 km. As expected for a cross-sectional specification with rich spatial controls, the standard errors widen somewhat at the longest bandwidths (100 and 150 km), where the spatial HAC kernel absorbs progressively more of the variation already captured by the included geographic covariates; the p-value at 100 km is 0.091 and at 150 km is 0.130. The result thus holds under the bandwidths most commonly applied in the spatial-econometrics literature for cross-sectional designs at this geographic scale.

Standard error	Estimate	SE	p
Heteroskedasticity-robust (paper)	-0.0109	(0.0054)	0.045
Conley 25 km	-0.0109	(0.0048)	0.025
Conley 50 km	-0.0109	(0.0053)	0.039
Conley 75 km	-0.0109	(0.0051)	0.033
Conley 100 km	-0.0109	(0.0064)	0.091
Conley 150 km	-0.0109	(0.0072)	0.13

Table A.53: Conley SE Bandwidth Sensitivity: Domain Tax Regression

Note: The coefficient on $\log(\hat{\theta}_j)$ from Table 4 Model 4 (fully controlled) is significant at the 5% level for Conley bandwidths up to 75 km, marginal at 100 km ($p = 0.091$), and not significant at 150 km ($p = 0.130$). The first row reports the heteroskedasticity-robust standard error used in the main table.

E.3. Additional Analyses

E.3.1. Average Distance to Home Capital

Outcome:					τ_j					
Models:	(1)		(2)		(3)		(4)		(5: SPL)	
Variables										
ln(Avg. Dist to HC$_j$)	**−0.0234**	**(0.0072)**	**−0.0226**	**(0.0072)**	**−0.0114**	**(0.0068)**	**−0.0114**	**(0.0069)**	**−0.0082**	**(0.0060)**
ln(Rice)			−0.0065	(0.0151)	0.0059	(0.0154)	0.0108	(0.0162)	−0.0001	
ln(Area)			−0.0194	(0.0111)	−0.0291	(0.0140)	−0.0373	(0.0159)	−0.0212	(0.0119)
ln(Pop.)			0.0421	(0.0168)	0.0269	(0.0144)	0.0242	(0.0150)	0.0221	(0.0128)
New FL			0.0001	(0.0003)	0.0004	(0.0004)	0.0003	(0.0004)	0.0003	(0.0002)
Shog. Ally			−0.0340	(0.0175)	0.0049	(0.0160)	0.0001	(0.0163)	0.0012	(0.0128)
Goshi			0.0209	(0.0258)	0.0196	(0.0263)	0.0269	(0.0275)	0.0244	(0.0205)
lat					−0.4960	(0.2905)	−0.4513	(0.3119)	−0.3223	(0.0503)
lon					−0.1680	(0.0747)	−0.1585	(0.0805)	−0.1176	(0.0126)
ln(Elev.)					−0.0111	(0.0158)	−0.0128	(0.0163)	−0.0090	(0.0136)
ln(Rugg.)					0.0391	(0.0159)	0.0352	(0.0157)	0.0227	(0.0153)
ln(Dist. Edo)					−0.0333	(0.0161)	−0.0438	(0.0175)	−0.0316	(0.0128)
Rainy					−1.638	(0.8959)	−1.586	(0.9123)	−1.0313	(0.7128)
Snowy					0.6554	(1.869)	0.6248	(1.886)	−0.2209	(1.3317)
Lat × Lon					0.0039	(0.0021)	0.0036	(0.0023)	0.0026	(0.0004)
Kin Nbs.							0.0460	(0.0295)	0.0199	(0.0261)
Nb. FL							0.0002	(0.0001)	0.0001	(0.0001)
ln(Nb. Rice)							−0.0259	(0.0190)	−0.0233	(0.0145)
ln(Nb. Area)							0.0182	(0.0178)	0.0136	(0.0130)
ln(Nb. Pop.)							−0.0056	(0.0211)	−0.0000	(0.0000)
Const.	0.4680***	(0.0249)	0.2421**	(0.1226)	22.09**	(10.17)	21.05*	(10.81)	15.415***	(1.7083)
Fit Statistics										
Num.Obs.	232		230		230		230		230	
R2	0.046		0.135		0.367		0.385		0.435[a]	

Table A.54: Average Distance to Home Capital Weakly Predicts Tax Rates

Note: The coefficients represent OLS estimates (maximum likelihood estimate in Model (5: Spatial Lag)) of the relationship between the average absolute distance to the home capital (log) and the domain tax rate. Heteroskedasticity-robust standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1. [a] Nagelkerke pseudo-R2

E.3.2. Controlling for the Shortest Distance to the Border

Outcome:	τ_j									
Models:	(1)		(2)		(3)		(4)		(5: SPL)	
Variables										
$\ln(\hat{\theta}_j)$	**−0.0241**	**(0.0055)**	**−0.0196**	**(0.0058)**	**−0.0111**	**(0.0054)**	**−0.0111**	**(0.0054)**	**−0.0106**	**(0.0052)**
$\ln(\text{Avg. Dist. Border}_j)$	**0.0010**	**(0.0067)**	**−0.0112**	**(0.0084)**	**−0.0058**	**(0.0070)**	**−0.0041**	**(0.0067)**	**−0.0030**	**(0.0076)**
ln(Rice)			−0.0010	(0.0151)	0.0081	(0.0155)	0.0120	(0.0163)	0.0010	
ln(Area)			−0.0191	(0.0120)	−0.0282	(0.0149)	−0.0314	(0.0156)	−0.0213	(0.0117)
ln(Pop.)			0.0379	(0.0164)	0.0255	(0.0144)	0.0225	(0.0150)	0.0218	(0.0109)
New FL			0.0001	(0.0003)	0.0004	(0.0004)	0.0003	(0.0004)	0.0003	(0.0002)
Shog. Ally			−0.0314	(0.0177)	0.0061	(0.0161)	0.0017	(0.0164)	0.0027	(0.0136)
Goshi			0.0262	(0.0262)	0.0207	(0.0263)	0.0279	(0.0272)	0.0248	(0.0203)
lat					−0.5831	(0.2929)	−0.5272	(0.3086)	−0.3867	(0.0439)
lon					−0.1891	(0.0753)	−0.1767	(0.0799)	−0.1333	(0.0110)
ln(Elev.)					−0.0133	(0.0160)	−0.0147	(0.0165)	−0.0108	(0.0133)
ln(Rugg.)					0.0411	(0.0158)	0.0383	(0.0157)	0.0248	(0.0151)
ln(Dist. Edo)					−0.0380	(0.0158)	−0.0454	(0.0171)	−0.0344	(0.0126)
Rainy					−1.645	(0.8924)	−1.574	(0.9051)	−1.0184	(0.6662)
Snowy					0.6475	(1.850)	0.6105	(1.857)	−0.2900	(1.1784)
Lat × Lon					0.0045	(0.0021)	0.0041	(0.0023)	0.0031	(0.0003)
Kin Nbs.							0.0406	(0.0295)	0.0197	(0.0257)
Nb. FL							0.0002	(0.0001)	0.0001	(0.0001)
ln(Nb. Rice)							−0.0244	(0.0187)	−0.0218	(0.0136)
ln(Nb. Area)							0.0179	(0.0176)	0.0118	(0.0119)
ln(Nb. Pop.)							−0.0040	(0.0206)	−2.0e-08	(9.8e-08)
Nb. FL							0.0002	(0.0001)	0.0001	(0.0001)
Const.	0.4135	(0.0137)	0.1671**	(0.1206)	25.01	(10.25)	23.60***	(10.83)	17.557	(1.4743)
Fit Statistics										
Num.Obs.	232		230		230		230		230	
R2	0.079		0.150		0.374		0.392		0.441[a]	

Table A.55: Rulers Facing Higher Governance Conflict Extracted Less Tax

(Controlling for the Average Distance to Border)

Note: The coefficients represent OLS estimates (maximum likelihood estimate in Model (5: Spatial Lag)) of the relationship between the average distance ratio of each domain and the domain tax rate. Heteroskedasticity-robust standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1. [a] Nagelkerke pseudo-R2

E.3.3. Conditional Role of Governance Conflict: Using Population Density

Outcome:					τ_j					
Models:	(1)		(2)		(3)		(4)		(5)	
Variables										
$\ln(\hat{\theta}_j)$ × Low	**−0.0295**	**(0.0104)**	**−0.0251**	**(0.0097)**	**−0.0190**	**(0.0071)**	**−0.0258**	**(0.0106)**	**−0.0198**	**(0.0071)**
$\ln(\hat{\theta}_j)$ × High	**−0.0217**	**(0.0057)**	**−0.0174**	**(0.0061)**	**−0.0076**	**(0.0058)**	**−0.0208**	**(0.0061)**	**−0.0068**	**(0.0060)**
ln(Rice)			−0.0021	(0.0156)	0.0069	(0.0153)	−0.0011	(0.0155)	0.0063	(0.0147)
ln(Area)			−0.0240	(0.0114)	−0.0298	(0.0142)	−0.0239	(0.0114)	−0.0278	(0.0138)
ln(Pop.)			0.0384	(0.0167)	0.0253	(0.0143)	0.0362	(0.0169)	0.0233	(0.0146)
New FL			0.0001	(0.0003)	0.0005	(0.0004)	0.0001	(0.0003)	0.0004	(0.0004)
Shog. Ally			−0.0309	(0.0178)	0.0077	(0.0161)	−0.0314	(0.0181)	0.0045	(0.0160)
Goshi			0.0210	(0.0257)	0.0179	(0.0262)	0.0237	(0.0260)	0.0216	(0.0267)
lat					−0.6052	(0.3029)			−0.5778	(0.3132)
lon					−0.1966	(0.0782)			−0.1913	(0.0813)
ln(Elev.)					−0.0128	(0.0159)			−0.0122	(0.0159)
ln(Rugg.)					0.0410	(0.0158)			0.0387	(0.0158)
ln(Dist. Edo)					−0.0388	(0.0159)			−0.0418	(0.0172)
Rainy					−1.622	(0.9139)			−1.450	(0.9269)
Snowy					0.5002	(1.882)			0.2750	(1.861)
Lat × Lon					0.0047	(0.0022)			0.0045	(0.0023)
Kin Nbs.							0.0437	(0.0302)	0.0353	(0.0303)
Nb. FL							5.01e-5	(9.71e-5)	0.0002	(0.0001)
ln(Nb. Rice)							−0.0109	(0.0087)	−0.0126	(0.0085)
Const.	0.4150	(0.0088)	0.1891**	(0.1242)	25.95	(10.63)	0.3474**	(0.1719)	25.37	(11.01)
Fit Statistics										
Num.Obs.	232		230		230		230		230	
R2	0.081		0.146		0.378		0.159		0.392	

Table A.56: Rulers Facing Higher Governance Conflict Extracted Less Tax:

Interacted with Neighboring Domains' Economic Condition (Using Population Density as a Proxy for $\hat{y}_{-j}$)

Note: The coefficients represent OLS estimates of the relationship between the average distance ratio measure of each domain interacted with the population density of its neighbors (low vs. high) and the domain tax rate. Heteroskedasticity-robust standard errors in parentheses. Signif. Codes: .***: 0.01, .**: 0.05, .*: 0.1.

E.3.4. Fixed-Neighbor-Set and HMX Kernel Diagnostics

The outside-option interaction in Figure 8 and Table A.44 uses a version of $\hat{y}_{-j}$ weighted toward the neighbors whose capitals are physically closest to home-domain villages, the same geometry that also defines the exposure measure $\ln\left(\hat{\theta}_j\right)$. A natural concern is that the interaction coefficient thereby mechanically mixes the moderator with the treatment: domains with high $\ln\left(\hat{\theta}_j\right)$ have their $\hat{y}_{-j}$ shaped by a different neighbor set than domains with low $\ln\left(\hat{\theta}_j\right)$, so the interaction could reflect this compositional shift rather than genuine moderation by outside-option quality.

Table A.57 addresses this concern with two complementary checks. The first breaks the mechanical link. Panel A reconstructs $\hat{y}_{-j}$ as an unweighted simple mean of neighbor labor-absorption capacity (p_j^{labor}) over all domains within 100 km of j's capital, regardless of the nearest-foreign-capital weights used in the main text. Under this construction the asymmetry survives in the expected direction: in the fully-controlled Model 4, the $\ln\left(\hat{\theta}_j\right)$ slope is -0.0169 ($p = 0.0081$) where neighbors have high labor demand, compared with -0.0099 ($p = 0.1946$) where they do not, consistent with the main-text result.

The second check asks whether the asymmetry is an artifact of the linear interaction functional form. Following the kernel-smoothed diagnostic of Hainmueller, Mummolo, and Xu (2019), Panels B and C estimate the local-linear marginal effect of $\ln\left(\hat{\theta}_j\right)$ on τ_j at each point in the continuous moderator support, using an Epanechnikov kernel with 95% confidence intervals. Under both the paper's $\hat{y}_{-j}$ and the fixed-100 km $\hat{y}_{-j}$, the marginal effect is negative throughout the support; under the paper's moderator it is significant at the 25th percentile. The curve is flat rather than monotonic because the sample is small ($N \approx 230$) and the Epanechnikov kernel has limited resolution. Its purpose here is not to recover a fine-grained moderating pattern but to confirm that the marginal effect does not flip sign anywhere, ruling out the linearity-induced artifact the HMX diagnostic is designed to detect. Figure A.12 plots the full curves side by side.

Panel A: Fixed-Neighbor-Set $\hat{y}_{-j}$ (100 km Unweighted Mean)				
Moderator level	(1)	(2)	(3)	(4)
$\ln(\hat{\theta}_j) \times \hat{y}_{-j}^{\text{Low}}$	-0.0274	-0.0235	-0.0102	-0.0099
Robust SE	(0.0066)	(0.0070)	(0.0073)	(0.0076)
p	0.0001	0.0010	0.1646	0.1946
$\ln(\hat{\theta}_j) \times \hat{y}_{-j}^{\text{High}}$	-0.0196	-0.0157	-0.0154	-0.0169
Robust SE	(0.0072)	(0.0068)	(0.0065)	(0.0063)
p	0.0069	0.0219	0.0192	0.0081
Num. Obs.	231	229	229	229
R^2	0.080	0.145	0.375	0.394

Panel B: HMX Kernel Marginal Effect at Moderator Quartiles (paper's $\hat{y}_{-j}$)			
Moderator quartile	$\hat{\partial}\tau_j/\partial \ln(\hat{\theta}_j)$	95% CI	p
25th	-0.0135	[-0.0265, -0.0005]	0.0423
50th	-0.0134	[-0.0272, 0.0004]	0.0570
75th	-0.0141	[-0.0324, 0.0042]	0.1316
Panel C: HMX Kernel Marginal Effect at Moderator Quartiles (fixed-100 km $\hat{y}_{-j}$)			
25th	-0.0113	[-0.0241, 0.0015]	0.0841
50th	-0.0111	[-0.0262, 0.0039]	0.1466
75th	-0.0141	[-0.0412, 0.0130]	0.3079

Table A.57: Tax-Exposure Interaction: Fixed-Neighbor-Set and HMX Kernel Marginal Effects

Note: Panel A re-estimates the interaction specifications 1-4 of Table A.44 after replacing the nearest-foreign-weighted outside-option proxy $\hat{y}_{-j}$ with an unweighted mean of p_j^{labor} over all domains within 100 km of j's capital. This breaks the mechanical link between the capital geometry that defines $\ln(\hat{\theta}_j)$ and the neighbor set that defines $\hat{y}_{-j}$ flagged by the methodologist referee. The "High-minus-Low" asymmetry from Figure 8 survives: under the fully-controlled Model 4 the $\hat{y}_{-j}^{\text{High}}$ slope is -0.017 ($p = 0.008$), versus -0.010 ($p = 0.19$) for $\hat{y}_{-j}^{\text{Low}}$. Panel B and Panel C report the Hainmueller-Mummolo-Xu (2019) kernel-smoothed marginal effect of $\ln(\hat{\theta}_j)$ on τ_j at the 25th, 50th, and 75th percentiles of the continuous moderator support, estimated by local linear regression with an Epanechnikov kernel after residualizing against the Model 4 control set. Under either moderator definition the marginal effect is negative across the entire support (Figure A.12 plots the full curve). The 95% CI excludes zero at the 25th percentile of the paper's $\hat{y}_{-j}$ and lies slightly above zero at the 75th percentile, consistent with a monotonically sharpening tax-exposure link as the outside option becomes more attractive.

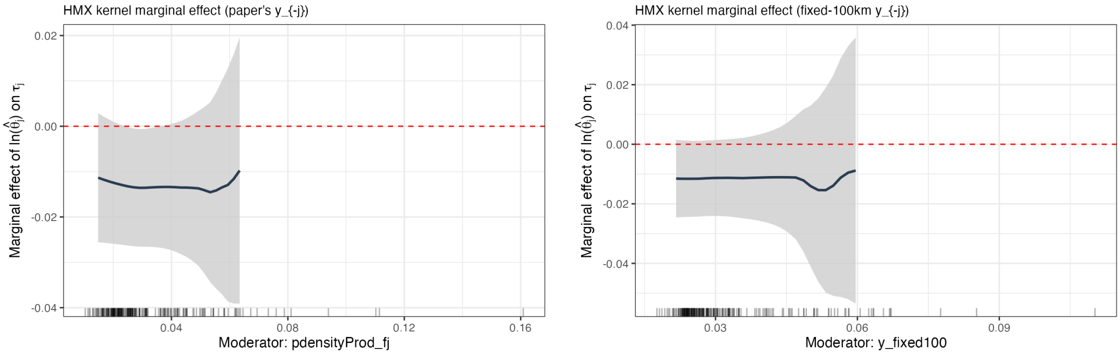


Figure A.12: HMX Kernel-Smoothed Marginal Effect of $\ln\left(\hat{\theta}_j\right)$ on τ_j

Note: Local-linear-kernel marginal effect of $\ln\left(\hat{\theta}_j\right)$ on the tax rate τ_j, estimated over the continuous support of the outside-option moderator $\hat{y}_{-j}$. Epanechnikov kernel with bandwidth equal to the interquartile range of the moderator; covariates residualized using the Model 4 control set. Shaded band: 95% CI. Red dashed line: zero. Rug: moderator-value density. Left: paper's nearest-foreign-weighted $\hat{y}_{-j}$. Right: fixed-100 km unweighted $\hat{y}_{-j}$. Both curves are negative over the middle of the moderator range, with the CI narrowest around the median where the support is densest.