Paradise Premiums: Buyer Origin, Information Frictions, and Housing Affordability in Hawai'i *

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Abstract

Using transaction-level residential real estate data from Hawai'i, we study how buyer origin affects housing prices and local outcomes. Nonlocal buyers pay 4.2-4.9 percent more for comparable properties with foreign buyers paying 9.8 percent more. Preference sorting does not explain the premium, and observable financial capacity explains only part of it. The remaining evidence is consistent with information frictions and anchoring: buyers from high-price states pay 1.6 percentage points more than other domestic nonlocals, the premium has declined with online search, and it is smallest in liquid markets. At the ZIP level, nonlocal demand raises prices but not employment, business activity, or resident displacement. Buyer-composition shifts explain little median ZIP price growth, with meaningful effects concentrated in Maui and Hawai'i Island. Our findings suggest that demand-side policies may offer localized relief but are unlikely to improve statewide affordability absent supply-side reform.

Keywords: Housing Markets, Hawai'i, Information Frictions, Affordability

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1 Introduction

International residential real estate capital flows have grown in both magnitude and political salience. Governments in Canada, Australia, Singapore, New Zealand, and across the European Union have enacted restrictions on foreign residential purchases, citing affordability pressures on local residents (Li et al., 2024). A growing empirical literature documents the housing price and real economic effects of these flows (e.g., Badarinza and Ramadorai, 2018; Gorback and Keys, 2020; Li and Chau, 2024). However, the extent to which these findings generalize to geographically isolated and supply-constrained island markets remains unclear. This paper contributes to the literature by examining Hawai'i, a setting overlooked in prior research.

Hawai'i differs from prior settings in a way that economically matters. Its geographic isolation from both the continental United States and foreign markets implies that nearly all nonlocal buyers face elevated search costs, institutional unfamiliarity, and limited local market knowledge relative to resident buyers. As a result, geographic proximity provides little variation in information frictions, allowing us to use buyer origin to distinguish between mechanisms underlying nonlocal transaction premia documented in prior work.[1] Additionally, public debate over the housing affordability crisis in Hawai'i routinely centers on the role of nonresident and foreign buyers, vacancy, and short-term rentals. These concerns have translated into concrete policy proposals, including restrictions on foreign ownership of residential property, the phase-out of vacation rentals in Maui, and vacancy taxes in Honolulu (Alfonseca, 2024; Associated Press, 2024; Honolulu Civil Beat, 2024). All of these proposals rest on the premise that outside capital meaningfully drives unaffordability for local residents. Whether that premise is empirically warranted is the central question this paper addresses.

We make four primary contributions. First, we revisit the previously documented nonlocal buyer price premia in Hawai'i, building on Miller et al. (1988) with a substantially larger statewide transaction dataset, an improved identification strategy, and a contemporary sample period. Second, by combining rich spatial controls, buyer-origin classifications, and within-property comparisons, we assess competing mechanisms—search-cost frictions, anchoring, and preference heterogeneity—as drivers of the nonlocal premium. Third, we move beyond transaction prices to examine whether nonlocal demand is associated with broader local economic outcomes, including aggregate home prices, employment, business activity, and residential displacement, and ask whether any associa-

[1] See, for example, Devaney and Scofield (2017); Kandlbinder et al. (2019); Allen et al. (2023); Li and Chau (2024); Cvijanović and Spaenjers (2021); Miyakawa et al. (2024).

tions persist in the long run. Lastly, we assess the economic magnitude of nonlocal housing demand as an explanation for Hawai'i's long-run aggregate price growth.

Our analysis reveals that nonlocal buyers pay approximately 4.2–4.9 percent more than local buyers for comparable properties, a result robust to hedonic controls, a suite of interacted fixed effects, and repeat-sales specifications. We first assess whether this premium reflects preference heterogeneity or differential financial capacity before turning to informational mechanisms. The premium persists after excluding coastal properties and luxury segments where nonlocals disproportionately transact, arguing against a pure preference-based account. Splitting the sample by loan-to-value ratio and loan size relative to the local market, we find that the premium attenuates among buyers who are observably less wealthy, but does not disappear. Wealth and the discretionary nature of nonlocal purchases likely contribute at the margin, but financial capacity differences alone cannot account for the result.

The residual premium is most consistent with informational explanations. Relative to local buyers, nonlocals from California and other high-price US states both pay premia of roughly 4.7 percent, statistically larger than the 3.1 percent premium paid by buyers from other US states. This pattern is consistent with a baseline role for search-cost frictions — affecting all nonlocal buyers — augmented by anchoring. Buyers from high-HPI states pay a 1.6 percentage point higher premium than other US buyers despite plausibly similar informational environments, consistent with reference-price effects. California buyers exhibit a similar premium to high-HPI states; despite likely lower informational frictions due to stronger ties to Hawai'i, their elevated home-market prices sustain high willingness to pay. Foreign buyers pay the largest premium at 9.8 percent, reflecting steep barriers to acquiring local market expertise.

Consistent with the role of internet in facilitating information sharing (Gordon and Winkler, 2019), we find that the nonlocal premia has diminished over time from 6.7% in early 2000's to about 3.5% in recent years. Geographically, the premium concentrates in less urban markets, with Honolulu showing a statistically insignificant 0.8%, and Maui, Kauai, and Hawai'i[2] counties exhibiting a statistically significant premium. Collectively, these set of findings strengthen our interpretation of informational explanations as the primary driver of the nonlocal transaction premium.

We then turn to the real economic effects of nonlocal activity. A one-percentage-point increase in nonlocal transaction share is associated with a 0.6 percent rise in median ZIP home prices, 1.5

[2]The State of Hawai'i also has a county named Hawai'i.

percent higher employment, and 1.0 percent more business establishments in cross-sectional panel specifications, with employment effects concentrated in the non-tradable sector. We find no evidence of residential displacement across any income group. However, these cross-sectional associations reflect persistent differences across ZIP codes rather than causal spillovers: once we add ZIP fixed effects, the employment and establishment associations vanish entirely while the price effect survives.

Long-difference specifications that directly absorb time-invariant ZIP characteristics confirm this pattern. Only the price effect is robust: a one-percentage-point increase in nonlocal buyer share over two decades is associated with 0.8–1.4 percent higher cumulative price appreciation, while employment and establishment estimates are economically small and statistically indistinguishable from zero. A back-of-envelope decomposition applies this estimate ZIP-by-ZIP: the median ZIP attributes essentially none of its total price growth to shifts in nonlocal buyer composition, and the inter-quartile range of this share is narrow across all counties. The largest contributions are confined to select markets on Maui and Hawai'i Island, where nonlocal penetration grew materially over the sample period.

Our study provides a policy-relevant assessment of nonlocal demand in a high-demand, supply-constrained housing market. Although nonlocal buyers are capitalized into prices at the margin, their aggregate contribution to Hawai'i's affordability crisis is limited and geographically uneven. This distinction matters for policy. In destination markets on Maui and Hawai'i Island, where nonlocal penetration has grown materially and its contribution to price appreciation is nontrivial, targeted demand-side interventions—such as vacancy taxes or short-term rental restrictions—may provide localized relief. Statewide, however, demand-side restrictions alone are unlikely to substantially improve affordability without complementary supply-side reforms. More broadly, the results suggest that restrictions on outside buyers are most likely to have meaningful price effects in high-amenity markets where outside demand is large, spatially concentrated, and not dominated by binding supply constraints.

2 Literature Review

A substantial body of research demonstrates that housing markets are inherently heterogeneous across regions and buyer groups. House valuation and market behavior depend not only on asset characteristics but also on local socioeconomic and informational contexts. Prior studies have highlighted the influence of demographics (Poterba et al., 1991), school quality (Bayer et al., 2007), gentrification dynamics (Guerrieri et al., 2013), and information asymmetry regarding property and neighborhood characteristics (Kurlat and Stroebel, 2015). Regional variation in institutions, financing conditions, and information environments further shapes housing market outcomes (Broxterman and Zhou, 2023; Kiefer et al., 2023). Moreover, geographic constraints are a key determinant of urban development and housing market behavior in the United States (Saiz, 2010), while local fundamentals explain much of the observed variation in housing market performance across regions (Campbell et al., 2009). Collectively, this evidence underscores the need for region-specific studies to explicitly account for spatial, institutional, and informational nuances.

Hawai'i serves as a uniquely informative case study for regional market analysis. Owing to its geographic isolation and position between major national and international markets, the Hawai'i housing market exhibits a pronounced bifurcation in demand between local residents and nonlocal buyers. Hawai'i is widely recognized as a premier global destination drawing both transient tourists and long-term real estate investors (Eversole et al., 2024). The tropical weather and coastal allure incentivizes domestic and foreign buyers to acquire secondary residences or condominiums, further tightening the available housing stock (Ravago et al., 2008). Hawai'i is also the most expensive housing market in the United States (median home prices being greater than those in California by nearly 16%), the highest homelessness rate in the nation, and strict zoning laws (UHERO, 2025). The market is distinguished by the persistent presence of foreign buyers, particularly from Asia (Miller et al., 1988). Shaped by these unique geographic, institutional, and demographic forces, Hawai'i offers a compelling natural setting to examine how differences in buyer origin and information access may lead to deviations from efficient market pricing.

Extant literature explores the idea of local versus nonlocal buyers and pricing premiums in the real estate markets, though findings differ. A common finding is that nonlocal or foreign buyers tend to pay price premia, both within the United States for out-of-town, out-of-state, or international buyers (Devaney and Scofield, 2017; Kandlbinder et al., 2019; Gorback and Keys, 2020; Li and Chau, 2024) and in non-US housing markets where buyers are not domestic residents (Badarinza and Ramadorai, 2018; Cvijanović and Spaenjers, 2021; Miyakawa et al., 2024; Sá, 2025). In fact, Puy et al. (2020) find that foreign price premia in the United States increase following extreme political crises abroad, particularly in locations with strong diaspora links to affected countries. However, this pattern is not universal. Chinloy et al. (2013) find no premium for nonlocal buyers and instead show that experienced local buyers obtain better outcomes, highlighting the role of market experience rather than buyer origin per se.

Beyond just the existence of housing premiums between local and nonlocal buyers, many studies ask why these premiums exist, where three prominent theories dominate the literature. The first of these deals with information asymmetries and search costs. In the context of the real estate markets, it is hypothesized that nonlocal buyers likely have less exposure to the local housing market conditions and distributions, leading to inaccurate perceptions of housing prices, and are subject to higher search costs when transacting at a distance (Lambson et al., 2004; Ihlanfeldt and Mayock, 2012; Kandlbinder et al., 2019). Empirical evidence is mixed. Early studies find little support for systematic information-based premia (Turnbull and Sirmans, 1993; Watkins, 1998). In contrast, Lambson et al. (2004) show that out-of-state buyers pay significant premiums for apartments in the Phoenix area where these buyers lack the information to value distant properties. Ihlanfeldt and Mayock (2012) find that nonlocal buyers have less bargaining power, higher search costs, and less intricate knowledge of the local market relative to local buyers. Kandlbinder et al. (2019) document a consistent out-of-town premium as opposed to in-town counterparts, although the premium has diminished across time with the a of online platforms. Li and Chau (2024) use Hong Kong housing market data to show that nonlocal buyers pay a 2.8% premium in the second-hand housing market and prefer to purchase in the first-hand market when the problem of information asymmetry is serious.

The second prominent explanation deals with the anchoring effect, noted as a cognitive bias stemming from judgment heuristic reliance (Slovic and Lichtenstein, 1971; Tversky and Kahneman, 1974). In other words, individuals tend to rely on an initial value (or anchor point) and update their beliefs and estimates based on that point, where adjustments are insufficient and biased towards the initial value. Many real estate studies empirically confirm the presence of anchoring. Buyers coming from higher-priced markets are willing to pay more for homes as they have inflated expectations on house prices (Ihlanfeldt and Mayock, 2012; Zhou et al., 2015; Qiu et al., 2020).[3] Applied to our setting, anchoring does not straightforwardly explain why all nonlocal buyers may overpay relative to local Hawai'i residents. Hawai'i is one of the most expensive housing market in the US, and most mainland buyers arrive from comparatively lower-priced markets and would anchor downward rather than upward relative to locals. Instead, we hypothesize that a baseline premium over

[3]Anchoring has been documented in multiple real estate contexts beyond origin-market pricing. Northcraft and Neale (1987) show that even in information-rich environments the list price of a home significantly biased the pricing estimates of both non-professionals and experts in real estate alike. Diaz III and Wolverton (1998) agree with the notion of experts being biased by anchoring as they are influenced by their own previous estimates.

locals arises from search-cost frictions that all nonlocal buyers share regardless of origin, and that anchoring operates as a second, distinct channel that generates a gradient within the nonlocal population: buyers from higher priced states where home markets most closely approach Hawai'i price levels, anchor to higher reference prices and therefore exhibit greater willingness to pay than buyers arriving from lower-priced states. This decomposition — frictions as the foundation, anchoring as a within-nonlocal amplifier — guides our empirical strategy in Section 5.4.

A third hypothesis emphasizes preference heterogeneity, positing that persistent price differentials arise from distinct tastes for amenities, neighborhood characteristics, and housing quality rather than informational gaps alone (Bayer et al., 2007; Ferreira, 2010). For instance, Siebert and Seiler (2022) demonstrate that while nonlocals in Indiana pay a 12% premium, this gap is largely explained by a higher willingness to pay for specific attributes such as school quality and modern construction. In Hawai'i, nonlocals disproportionately target coastal markets (Khan et al., 2025). Since shoreline properties command significant premiums (Tarui et al., 2023), observed price differentials may simply reflect a systematic valuation of environmental amenities. This also extends to risk perception; Khan et al. (2025) find that nonlocals apply smaller discounts to properties exposed to sea level rise, attributed to a divergence in climate change beliefs and local knowledge. We extend the coastal-focused analysis of Khan et al. (2025) to the entire market to examine if premiums arise from divergent preferences for a wider range of characteristics.

An important historical precedent for examining nonlocal buyer premiums in Hawai'i is Miller et al. (1988), who analyzed Japanese purchases in Honolulu's Waialae-Kahala and Waikiki submarkets during the 1986–1988 asset bubble. Based on 421 transactions, they documented that Japanese buyers paid approximately 19–21 percent more than local buyers, attributing this to exchange rate advantages, speculative motives, and information asymmetries. However, the external validity of these findings is limited by the specific institutional context of the late 1980s and the small sample size. Our study leverages a comprehensive dataset comprising thousands of transactions across the entire state to update this literature. The modern timeframe (2000-2023) captures a mature market where the digitization of listings has reduced the search frictions characterizing the pre-internet era, and where the macroeconomic dynamics driving foreign capital have shifted.

Methodologically, we employ a hedonic interacted fixed effects model following recent literature (Giglio et al., 2015; Stroebel, 2016; Bernstein et al., 2019), and separately, a repeat sales specification with property fixed effects to address omitted variable bias (Tyndall, 2023; Khan et al., 2025). This approach allows us to absorb unobserved heterogeneity more rigorously than earlier specifications. Our analysis provides a crucial region-specific analysis of Hawai'i, a market where nonlocals account for approximately 25% of transactions, a dynamic often sensationalized in the media but rarely rigorously quantified across the entire housing stock.[4]

3 Data

This study utilizes residential real estate transaction records and property attribute data sourced from Black Knight, a financial services company. Our dataset combines two primary information sources: deed records and property assessment files. The deed records capture essential transaction details, such as sale dates and property use classifications (residential versus commercial). The assessment files contain physical property attributes including room counts, square footage, property age, and structure type (such as single-family versus multi-family units). These datasets are linked through the Black Knight Distinct Property ID (BKDPID), which serves as a unique property identifier. Our analysis concentrates solely on residential properties, examining transactions that occurred from 2000 through 2023. We exclude cash purchases and inter-family transfers. To further ensure that our sample reflects genuine market transactions and to minimize the impact of extreme values, we excluded sales priced under $50,000 or exceeding $50,000,000 each year. We rely on shapefiles from the Hawai'i Statewide GIS Program's Geospatial Data Portal, for parcel geometries, coastline boundaries, and elevation profiles across the major Hawaiian islands. All spatial attributes (distance to the nearest coastline and elevation) are calculated at the parcel centroid. Our spatial analyses uses the WGS-1984 coordinate reference system.

Table 1 reports the summary statistics for residential real estate transactions across Hawaii's four counties and for our full State sample. Median sales prices reveal substantial variation, with Kaua'i ($575,500) and Maui ($556,800) commanding premiums over Honolulu ($471,700) and Hawai'i Island ($349,400). Property composition varies markedly: multi-family units dominate all markets except Hawai'i Island where single-family homes constitute 72.8% of transactions compared to just 26.4% in Honolulu, 36.9% in Maui, and 42.2% in Kaua'i. Nonlocal buyer penetration follows a clear pattern, with Maui (47.7%) and Kaua'i (49.1%) attracting substantially more outside capital than Honolulu (15.3%) and Hawai'i Island (38.3%), suggesting these counties function primarily as destination markets for second homes and investment properties. Properties in the sample averaged 5.0 rooms and

[4] Recent popular press coverage has highlighted debates surrounding foreign and domestic ownership in Hawai'i; see, for example, *Dwell Hawai'i* and *ABC News*.

Table 1: Summary statistics by County

	State	**Honolulu**	**Maui**	**Kaua'i**	**Hawai'i**
Sales Price ($2020)					
Mean	696,300	630,000	1,087,700	802,400	556,500
Median	464,000	471,700	556,800	575,500	349,400
Sd.	1,553,200	879,000	3,266,600	1,150,100	1,049,900
25th Percentile	287,600	302,200	333,000	367,200	215,800
75th Percentile	732,900	726,400	885,000	867,800	575,000
Transaction Characteristics					
Median square feet	1,122	1,044	1,008	1,298	1,336
Single-family (%)	37.5	26.4	36.9	42.2	72.8
Nonlocal (%)	26.4	15.3	47.7	49.1	38.3
Avg. Coastal Distance (m)	2,542	2,150	1,767	1,621	4,741
Avg. Rooms	5.0	5.0	4.7	5.2	5.1
Avg. Bedrooms	2.6	2.6	2.4	2.6	2.8
Age (Years)	26.8	28.6	24.9	25.2	23.3
Elevation (m)	92.1	49.9	90.2	60.4	240.6
Count	341,681	206,491	52,365	19,379	63,446

Note: This table presents county-level summary statistics. Hawai'i refers to the County of Hawai'i (the Big Island).

2.6 bedrooms, with a mean age of 26.8 years, and approximately 2,500m (1.56 miles) from the coast.

Table 2 reveals sharp differences in transaction prices, property types, and spatial characteristics across buyer groups. Median transaction prices rise monotonically from local buyers ($449,300) to mainland buyers ($516,000) and foreign buyers ($685,400), with substantially greater dispersion among nonlocal transactions, particularly in the upper tail of the price distribution. These price differentials coincide with pronounced differences in property composition: local buyers are far more likely to purchase single-family homes (39.2%) than mainland buyers (33.2%), while foreign buyers overwhelmingly transact in non-single-family units, with only 14% of purchases involving single-family properties. Nonlocal buyers also concentrate disproportionately in coastal areas, purchasing properties that are, on average, significantly closer to the shoreline and at lower elevations than those acquired by local residents. Compared to local buyers, both mainland and foreign buyers purchase homes with fewer rooms. Square footage is smaller for foreign buyers, but comparable between local and mainland buyers. Overall, nonlocal demand in Hawai'i is characterized not only by higher prices but also by systematic sorting into denser, coastal, and predominantly multifamily segments of the housing stock—features that must be explicitly controlled for.

Table 2: Summary statistics by buyer type

	Local buyers	Mainland buyers	Foreign buyers
Sales price ($2020)			
Mean	593,700	979,900	1,099,400
Median	449,300	516,000	685,400
Sd.	819,800	2,692,300	1,500,900
25th percentile	284,900	294,300	421,900
75th percentile	693,500	873,600	1,223,100
Transaction characteristics			
Median square feet	1,122	1,125	1,025
Single-family (%)	39.2	33.2	14.0
Avg. coastal distance (m)	2,770	1,920	1,267
Avg. rooms	5.1	4.7	4.4
Avg. bedrooms	2.7	2.4	2.1
Age (years)	27.3	25.5	26.1
Elevation (m)	94.5	86.6	36.7
Count	251,437	88,456	1,788

Note: This table reports summary statistics by buyer type. Local buyers are residents of Hawai'i. Mainland buyers reside in the continental United States, while foreign buyers reside outside the United States.

Restricting the sample to repeat-sale properties with exactly two transactions reduces the number of observations by approximately 70%. Figure A1 plots annual median sale prices for the full sample and the resulting repeat-sales subsample. The two series track each other closely throughout the sample period, with a correlation of 0.99, indicating that even the repeat sales subsample is representative of the overall market trend.

Next, we collapse transaction-level data to unique properties, classifying each based on its complete transaction history into one of three mutually exclusive categories: properties purchased only by local buyers, only by nonlocal buyers, or by both types at different points in time. This ensures that each physical location is entered into the analysis only once, regardless of trading frequency or holding period. Using these property-level classifications, we construct hexagonal spatial bins with a 1500 meter (0.93 miles) flat-to-flat distance and compute, for each bin, the share of properties that have ever been purchased by a nonlocal buyer. The numerator includes properties purchased exclusively by nonlocals or by both locals and nonlocals, and the denominator encompasses all properties in the bin. We restrict attention to bins containing at least five properties to minimize noise

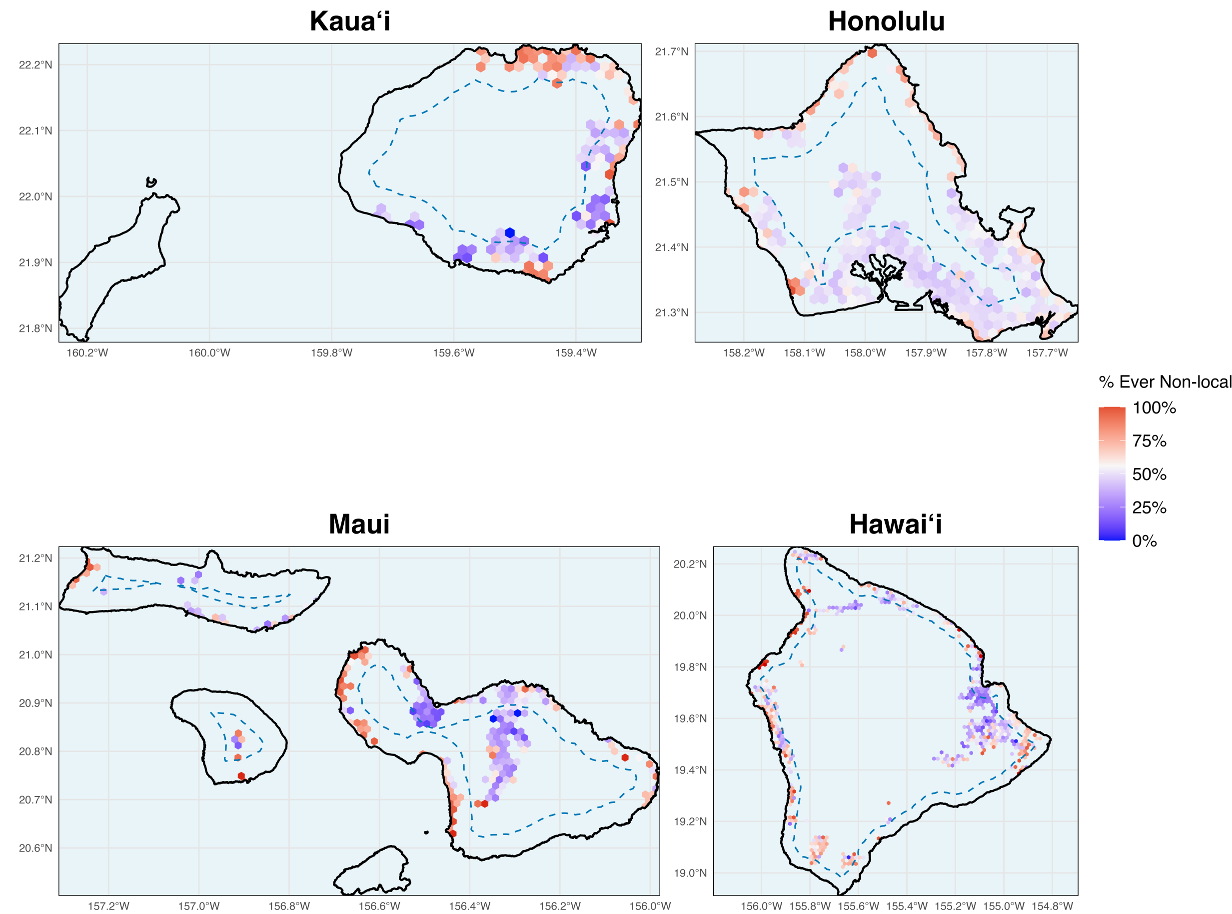

Figure 1: This figure plots the spatial distribution of properties ever purchased by nonlocal buyers over 2000–2023. Hexagonal bins with a 1,500-meter flat-to-flat distance report, for each bin, the share of properties that have ever been purchased by nonlocal buyers, including properties purchased exclusively by nonlocals or by both local and nonlocal buyers at different points in time. Only bins containing at least five properties are shown. The blue dotted line denotes a 3-mile inland coastal buffer.

in sparsely populated areas. Figure 1 plots the spatial distribution of nonlocal buyer participation in the housing market. Consistent with other studies using same dataset, we find that nonlocals are disproportionately present in the coastal segment (Khan et al., 2025).

4 Methods

To identify price differentials between local and nonlocal buyers, we compare observably equivalent properties that vary only by buyer type. Our baseline hedonic regression takes the following functional form:

$$\log(P_{it}) = \beta NonLocal_{it} + \gamma X_{it} + \lambda_{zybpc} + \epsilon_{it} \quad (1)$$

The dependent variable $\log(P_{it})$ represents the natural logarithm of the transaction price for property i in year t. $NonLocal_{it}$ is a binary variable that equals one if the property is purchased by an out-of-state buyer. X_{it} flexibly controls for property characteristics including age, square footage, and elevation profile (i.e. hundred equal sized buckets for each) following Stroebel (2016) and Bernstein et al. (2019). λ_{zybpc} represents interacted fixed effects between ZIP (z), year-month (y), number of bedrooms (b), property type (i.e. single family or multifamily) (p), and coastal distance (c), which we partition into hundred percentile bins based on its empirical distribution. The use of such interacted fixed effects is a common approach in real estate valuation literature (Giglio et al., 2015; Stroebel, 2016; Bernstein et al., 2019) and is critical to our identification strategy as it absorbs variation in house prices related to the interaction between location, time of sale, and property characteristics. ϵ_{it} is the residual idiosyncratic variation in sales prices. The main coefficient of interest, β, quantifies how much more (less) a nonlocal buyer pays over (under) that of a local buyer, for an observably equivalent property. We cluster standard errors at the ZIP level to account for spatial correlation in the error term. Additional robustness checks using two-way clustering and Conley (1999) standard errors are reported in the Appendix (Table A2).

In one specification, we use a categorical variable to replace the binary indicator $NonLocal_{it}$. Similar to Lambson et al. (2004), we define four mutually exclusive groups: (1) California buyers, (2) buyers from high-priced states, (3) buyers from other states, and (4) foreign buyers. We use this specification to examine how the price differential differs between various nonlocal buyer-types relative to in-state buyers. We also examine heterogeneity across geographical areas and time by interacting $NonLocal_{it}$ with relevant categorical buckets. Following Khan et al. (2025), this approach employs level means coding rather than traditional reference level coding, allowing each interaction coefficient to capture the effect of nonlocal buyers within a specific category. The corresponding hypothesis tests assess whether these effects differ significantly from zero.[5] The key advantage of level means coding is that it enables us to identify spatial and temporal nuances in price differentials between nonlocals and locals. Importantly, our identification still relies on comparing properties transacting in the same ZIP and year-month, with similar observable characteristics, differing only in buyer type (local vs.

[5] Alternative reference level coding (relative to some base category) would instead test whether the effect in one county differs from the reference category.

nonlocal).

To address any remaining concerns about omitted variable bias, we implement a repeat sales approach with property fixed effects (Tyndall, 2023; Khan et al., 2025). We restrict our sample to properties that sold exactly twice to ensure results are not driven by properties with unusually high transaction frequencies. For a property transaction i, the following equation is estimated:

$$\log(P_i) = \beta_1 NonLocal_i + \beta_2(Z_i \times T_i) + H_i + Y_i + \epsilon_i \quad (2)$$

The dependent variable $\log(P_i)$ is the log of the sale price. $NonLocal$ is a dummy variable indicating whether a nonlocal purchased the property. Z is the ZIP where property is located at. T is a continuous variable generated from the sale date.[6] H is a property level fixed effect and Y is a fixed effect for year-month of sale. This design effectively compares how the same physical unit sells in different time periods, net of (i) any time invariant property attributes (e.g. coastal distance) absorbed via H_i, (ii) common market wide price trends due to the inclusion of year-month fixed effects Y_i, and (iii) local area time trends ($Z_i \times T_i$) to account for potentially changing preferences across different areas. The coefficient of interest β_1 reflects the within-property premium/discount (or average % price difference) when it sells to a nonlocal rather than a local buyer. Note that while non-switching properties provide no identifying variation for our main coefficient, retaining them helps estimate aggregate time fixed effects and local trends with greater precision.

Two caveats merit discussion. First, property fixed effects cannot account for characteristic changes occurring between sales. For example, owners might make improvements that affect the value of the property, but these enhancements would not be captured by the fixed effects. Consequently, we leverage permit data when available to exclude properties that underwent substantial renovations (including fence upgrades, construction, electrical work, and plumbing). Minor renovations that we cannot observe are unlikely to bias results in a repeat sales framework (Billings, 2015). Second, restricting to properties selling only twice may yield a subsample unrepresentative of the overall real estate market (Melser, 2023). Figure A1 shows this is not the case.

[6] A sale occurring in the middle of 2020 takes a value of 2020.5.

5 Transaction Level Premiums

5.1 Main Findings

Table 3 displays our main results, where point estimates of the price differential paid by nonlocal buyers relative to locals are positive and significant across all specifications. Column 1 presents results controlling only for property characteristics including square footage, age, and elevation. In this specification, nonlocal buyers pay 15.6% more than locals. In Column 4, our preferred specification, we see a significantly positive (yet reduced) 4.2% higher price paid by the nonlocal buyer. Importantly, this premium is unlikely to be driven primarily by short-horizon speculative or investment motives. In Column 6, we re-estimate the preferred specification restricting the sample to properties that transact only once during the sample period and continue to find a comparable nonlocal premium.[7] Our findings are further corroborated by the repeat-sales specification, which exploits within-property variation and yields a 4.1% nonlocal premium (Column 7).

5.2 Preference Based Explanations

One intuitive explanation for the nonlocal premium in Hawaiian real estate is that buyers from outside Hawai'i simply have different preferences than local purchasers. The descriptive evidence supports this notion: Both Figure 1 and Table 2 demonstrate that nonlocal buyers systematically gravitate toward properties closer to the coast. If nonlocals possess a stronger preference for coastal proximity and are willing to pay more for such locations, they would naturally outbid other buyers for these desirable properties. To test whether coastal allure alone drives the observed premium, we systematically exclude properties at increasing distances from the coast and re-estimate the hedonic model. Table 4 presents the results of this robustness check. Column (1) excludes properties within 0.25 miles of the coast, while subsequent columns progressively expand this exclusion zone to 0.5, 1, 2, 3, and ultimately 6 miles from the shoreline. Remarkably, the nonlocal premium persists across all specifications. Even when excluding all properties within six miles of the coast, nonlocal buyers pay a statistically significant premium of 4.0 percent. This finding strongly suggests that coastal proximity alone cannot fully account for the price differential.

A related preference-based explanation concerns property quality and luxury characteristics. The

[7] Restricting the sample to properties that transact only once during the sample period eliminates repeat sales and short-horizon resales. Since housing speculation is typically proxied by rapid turnover or short holding periods (Mills, 1983; Malpezzi and Wachter, 2005; Wang et al., 2024; Khan et al., 2025), the persistence of a positive nonlocal price premium in this restricted sample suggests that short-run speculative resale activity is unlikely to drive the observed price differential.

Table 3: Nonlocal Premium

This table presents ordinary least squares estimates where the dependent variable is log(Price) for Equations 1 and 2. The explanatory variable of interest is NonLocal, a binary variable equal to one if the property is purchased by an out-of-state buyer. Column 1 includes controls for square footage, house age, and elevation. Column 2 adds ZIP × year-month fixed effects ($Z \times Y$). Column 3 further interacts these fixed effects with bedrooms and property type ($Z \times Y \times B \times P$). Column 4, our preferred specification, adds coastal distance interactions ($Z \times Y \times B \times P \times C$). Column 5 replaces year-month fixed effects with year-quarter interactions ($Z \times YQ \times B \times P \times C$). Column 6 re-estimates our preferred specification restricting the sample to properties that transact only once during the sample period (i.e., single-transaction properties). Column 7 estimates a repeat-sales specification with property and year-month fixed effects. Standard errors clustered at the ZIP level are reported in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
NonLocal	0.156***	0.121***	0.109***	0.042***	0.046***	0.049***	0.041***
	(0.052)	(0.019)	(0.015)	(0.010)	(0.010)	(0.017)	(0.005)
Square Footage	Y	Y	Y	Y	Y	Y	Y
House Age	Y	Y	Y	Y	Y	Y	Y
Elevation	Y	Y	Y	Y	Y	Y	Y
Z × Y	N	Y	N	N	N	N	N
Z × Y × B × P	N	N	Y	N	N	N	N
Z × Y × B × P × C	N	N	N	Y	N	Y	N
Z × YQ × B × P × C	N	N	N	N	Y	N	N
Property, Year-Month, ZIP Trend	N	N	N	N	N	N	Y
R^2	0.352	0.710	0.805	0.957	0.927	0.985	0.946
Adjusted R^2	0.352	0.692	0.744	0.852	0.841	0.907	0.846
N	341,681	341,681	341,681	341,681	341,681	92,787	102,544

summary statistics in Table 2 reveal that nonlocal buyers disproportionately purchase higher-priced properties. To address this concern, Table A3 restricts the analysis to progressively lower segments of the price distribution. Column (1) limits the sample to properties priced below the 90th percentile, effectively excluding the most expensive luxury homes, while subsequent columns impose increasingly stringent restrictions. The nonlocal premium remains positive and statistically significant across all price segments. Even when restricting the analysis to properties priced below the median (Column 5), nonlocal buyers pay approximately 2.3 percent more than local buyers for comparable homes. This pattern indicates that the premium is not merely an artifact of nonlocals concentrating their purchases in the luxury market.

Taken together, these results substantially weaken preference-based explanations grounded in coastal proximity or luxury segmentation. The premium persists across increasingly restrictive samples and, if anything, is slightly larger among the most inland properties (4.0 percent beyond six miles versus 3.1 percent within 0.25 miles) — the reverse of what preference-based coastal sorting would predict. A pure preference-based account would also imply a steeper premium in high-end segments, yet the estimate is stable down to the median of the price distribution. While differences in willingness to pay cannot be ruled out entirely, preferences alone cannot account for the pattern.

Table 4: Nonlocal Premium: Excluding Coastal Sample

This table presents estimates for Equation 1 where the dependent variable is log(Price). Each column sequentially excludes properties within increasing coastal buffer zones: Column (1) excludes those within 0.25 miles of the coast; Column (2) within 0.5 miles; Column (3) within 1 miles; Column (4) within 2 mile; Column (5) within 3 miles; and Column (6) within 6 miles. Standard errors, clustered at the ZIP level, are reported in parentheses. * p < 0.10, ** p < 0.05, *** p < 0.01.

	(1) $\geq$ 0.25 mi	(2) $\geq$ 0.5 mi	(3) $\geq$ 1 mi	(4) $\geq$ 2 mi	(5) $\geq$ 3 mi	(6) $\geq$ 6 mi
NonLocal	0.031***	0.024***	0.023***	0.030***	0.032***	0.040**
	(0.009)	(0.008)	(0.007)	(0.010)	(0.011)	(0.018)
Square Footage	Y	Y	Y	Y	Y	Y
House Age	Y	Y	Y	Y	Y	Y
Elevation	Y	Y	Y	Y	Y	Y
Z $\times$ Y $\times$ B $\times$ P $\times$ C	Y	Y	Y	Y	Y	Y
R^2	0.956	0.964	0.966	0.959	0.953	0.951
Adjusted R^2	0.844	0.862	0.874	0.862	0.855	0.847
N	269,436	214,569	151,295	79,007	54,363	19,587

5.3 Buyer Financial Capacity

Nonlocal buyers may differ from local buyers in wealth, income, or financial capacity, and these differences may contribute to the estimated premium. Wealthier buyers may have greater ability or willingness to pay, and nonlocal buyers may be drawn disproportionately from higher points of the wealth distribution. A related consideration is that the purchase serves different roles across buyer types. For local buyers, the transaction is more likely to reflect primary housing demand, whereas for nonlocal buyers it may more often be discretionary. Both channels could generate a price premium independently of information frictions, and we do not claim to fully disentangle them.

Our sample already excludes cash transactions, which removes the segment of the market where wealth differences are likely to be most pronounced. To further assess the role of financial capacity, we estimate the baseline specification on subsamples defined by observable mortgage characteristics that proxy for buyer resources. We split the sample by loan-to-value ratio, separating high-LTV transactions (LTV $\geq$ 0.90) from low-to-moderate LTV transactions, and by loan amount relative to the ZIP-year median loan amount among local buyers. If wealth differences are the primary driver of the premium, we would expect the premium to be substantially attenuated among high-LTV buyers and among borrowers at or below the local benchmark.

Table 5 reports the results. Among high-LTV buyers, the premium is approximately 2.2 percent,

Table 5: Nonlocal Premium by Loan-to-Value and Local Loan Size

This table presents estimates for Equation 1 where the dependent variable is log(Price). Columns (1) and (2) split the sample by loan-to-value ratio. High-LTV transactions are defined as those with LTV ≥ 0.90, while low-to-moderate LTV transactions have LTV below 0.90. Columns (3) and (4) split the sample by whether the transaction loan amount falls at or below versus above the ZIP-year median loan amount among local buyers, proxying for whether the nonlocal buyer is borrowing at a scale comparable to the typical local buyer in that market. Standard errors, clustered at the ZIP level, are reported in parentheses. * p < 0.10, ** p < 0.05, *** p < 0.01.

	(1) High LTV	(2) Low LTV	(3) Below Median	(4) Above Median
NonLocal	0.022**	0.046***	0.026***	0.058***
	(0.010)	(0.008)	(0.008)	(0.011)
Square Footage	Y	Y	Y	Y
House Age	Y	Y	Y	Y
Elevation	Y	Y	Y	Y
$Z \times Y \times B \times P \times C$	Y	Y	Y	Y
R^2	0.997	0.989	0.984	0.976
Adjusted R^2	0.965	0.934	0.916	0.853
N	38,291	108,607	53,057	114,138

roughly half the magnitude estimated among low-to-moderate LTV buyers (4.6 percent). A similar pattern holds for loan size: the premium is 2.6 percent among nonlocal buyers borrowing at or below the local median, compared with 5.8 percent above it. Overall, differences in financial capacity appear to explain part, but not all, of the estimated nonlocal premium.

5.4 Informational Explanations

Having ruled out preference-based sorting and bounded the contribution of differential financial capacity, we now examine the informational mechanisms that account for the residual premium. We present evidence consistent with two primary informational channels: search-cost frictions and behavioral anchoring. Canonical real estate search models imply that buyers facing higher search costs conduct fewer searches and, on average, transact at higher prices (Turnbull and Sirmans, 1993; Lambson et al., 2004). Intuitively, information frictions arise because property values depend not only on observable physical and financial characteristics, but also on locally-specific factors such as school quality, zoning regulations, taxes, and neighborhood conditions that are costly to acquire and interpret. In-state buyers accumulate this information through repeated exposure to their environment, whereas out-of-state buyers must incur explicit costs to obtain comparable knowledge, including time, travel, and reliance on intermediaries (Lambson et al., 2004). A second mechanism involves biased prior beliefs, specifically anchoring. Given that many nonlocal buyers in Hawai'i originate from high-cost markets (such as California), they may rely on home-market prices as arbitrary reference points when forming value estimates. This leads to incomplete adjustments when faced with local Hawai'i market data (Tversky and Kahneman, 1974). In our setting, anchoring suggests that the nonlocal premium is not merely a result of missing information, but a systematic bias where nonlocal buyers perceive Hawai'i's high prices as reasonable relative to their origin markets.

Prior work typically proxies information asymmetry using geographic proximity, the rationale being distance is directly related to search costs (Lambson et al., 2004). Given Hawaii's geographic isolation from both foreign markets and the continental United States, proximity-based measures are less informative in our setting. Consequently, we classify buyers by origin (In-State, Out-of-State, Foreign) rather than a continuous distance metric to better capture shifts in search costs and institutional familiarity.

We divide nonlocal buyers into four mutually exclusive groups. We isolate California buyers as this group occupies a distinct theoretical position. Their origin market—typically more expensive than Hawai'i—provides a strong anchoring signal, while their deep institutional and social ties to Hawai'i imply lower informational frictions relative to other out-of-state groups. This is exemplified by high market participation, with California buyers representing the largest nonlocal cohort, accounting for 47% of all nonlocal transactions (Figure A2) and strong transoceanic historical ties (Edinger-Marshall, 2000). Even today, modern ties remain strong through travel, with Los Angeles being Hawai'i's largest visitor source market, accounting for 9.1% of all statewide visitor arrivals in 2024 and 30.2% of all visitors from California (DBEDT, 2025).

Second, we consider buyers from other high-priced US states, defined as the ten states with the highest average FHFA house price index over our sample period.[8] While these markets are expensive relative to the US distribution, they remain less costly than Hawai'i. As such, anchoring still predicts a gradient within the nonlocal population: buyers from higher-priced origin markets carry higher reference price levels than those from lower-priced states, and thus exhibit a greater willingness to pay, even when all buyers are moving into a more expensive destination. At the same time, their weaker institutional ties imply greater informational frictions than for Californians. Third, buyers from the remaining US states plausibly face informational frictions but lack a clear anchoring channel, as their home-market prices are substantially below Hawai'i levels. Finally, we distinguish foreign buyers,

[8]We report these states in Table A1. Washington, DC, is treated as a separate state throughout the paper. Hawai'i and California are excluded.

who are likely to face the greatest search costs as well as additional institutional and informational barriers.

Figure 2 presents estimated price differentials relative to in-state buyers. We conduct *F*-tests to formally assess whether premia differ significantly across subgroups. All nonlocal buyers pay statistically significant premia relative to in-state buyers, consistent with informational disadvantages in geographically isolated housing markets. Buyers from both California and High-HPI states pay premia of approximately 4.7%, and the two estimates are statistically indistinguishable. One interpretation for this is that Californians' stronger anchoring—arising from higher origin-market prices—offsets their lower informational frictions, yielding premia comparable to those of high-HPI buyers with weaker ties to Hawai'i.

Both groups, however, pay significantly more than buyers from other US states, who exhibit a smaller but still positive premium of 3.1%. This gradient is consistent with anchoring operating within the nonlocal population: even when all buyers move into a more expensive market, those from higher-priced origin markets carry higher reference prices and thus exhibit a greater willingness to pay. Buyers from the remaining states plausibly face similar informational frictions but lack a strong price anchor relative to Hawai'i. We interpret the difference in premia across these groups as evidence that anchoring operates on top of a baseline information premium. Even more extreme is the case of foreign buyers, where the information asymmetry based on both distance and other barriers may result in even stronger results. Upon testing this, we find they pay a 9.8% premium. This estimate is statistically distinct from all other nonlocal buyer types.

We next examine how the nonlocal premium evolves over time. If informational frictions drive overpayment, the expansion of internet-based platforms (e.g., Zillow, Redfin, Trulia) should reduce the premium by lowering search costs and narrowing information gaps. For instance, Gordon and Winkler (2019) find that with the internet revolution, new-home premiums decreased over time, consistent with the role of internet in facilitating information sharing and reputation feedback. In Figure 3, we plot the premium for 4-year time periods. The premium peaked at 6.7% during 2000–2003 and then stabilized between 3.2% and 3.7% in the post-crisis period through 2023. The premium appears to vanish during the Global Financial Crisis (2008–2011), though the confidence intervals are wide and the estimate is statistically indistinguishable from both zero and later periods. The imprecision likely stems from the sharp contraction in nonlocal transaction volume during the GFC. A joint F-test confirms that the early 2000s premium is significantly different from both the late-period and

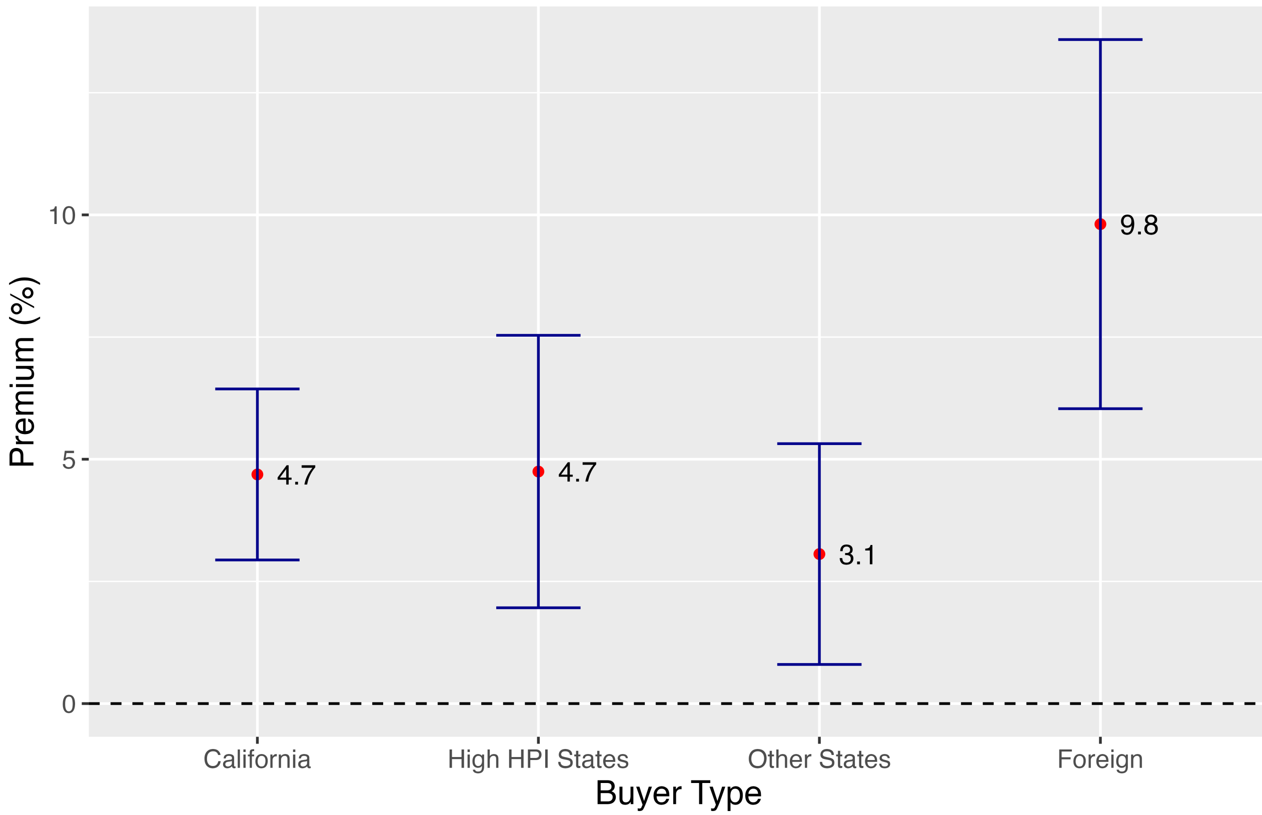

Figure 2: This figure displays the estimated coefficients from Equation 1, where *NonLocal* is a categorical variable with in-state buyers as the reference category. Each coefficient reflects the percent price premium paid by buyers from different origins, conditional on property characteristics and fixed effects. Estimates where the zero line falls outside the confidence bands are statistically significant at the $p < 0.05$ level. All nonlocal buyer groups pay statistically higher prices than in-state buyers at the $p < 0.05$ level. Pairwise F-tests indicate that the premiums paid by buyers from *California* and *High HPI States* are statistically indistinguishable from one another, while both differ significantly from the premium paid by buyers from *Other States* ($p < 0.05$). The premium paid by *Foreign* buyers is statistically distinct from all other buyer categories ($p < 0.05$).

crisis-era values. Overall, the evidence is consistent with a gradual convergence toward a lower but persistent steady state.

Higher transaction volumes improve information quality in real estate markets by facilitating price discovery, as each transaction reveals new information about property values and local market conditions. Markets with greater trading activity therefore exhibit lower information costs (Capozza et al., 2004; Clapp et al., 1995). Consistent with this mechanism, Honolulu County accounts for the largest share of property transactions in our data (Table 1) and serves as Hawaii's dominant economic center, with substantially higher real GDP, employment, and labor productivity than the neighbor-island counties (DBEDT, 2025). We next test the implication that information asymmetries are sub-

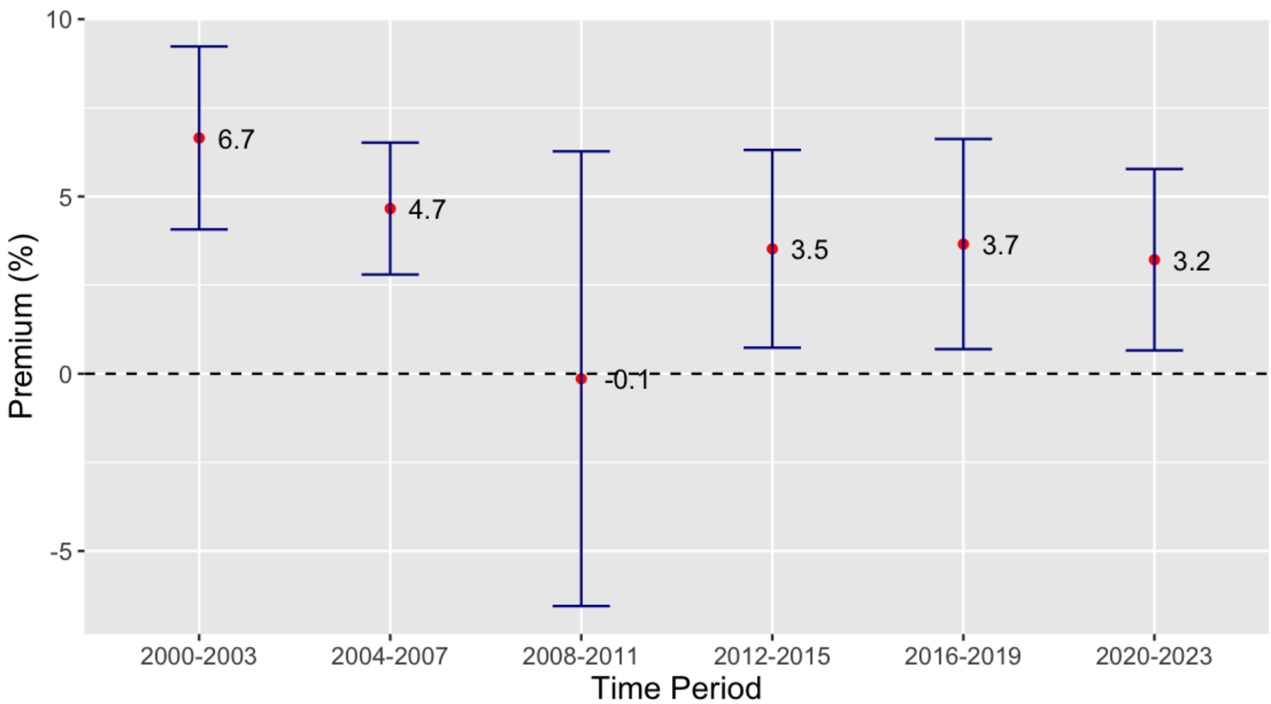

Figure 3: This figure plots the estimated coefficients from Equation 1, where the independent variable of interest is the interaction between *NonLocal* and *Period*, with periods defined in four-year intervals spanning 2000 to 2023. The estimation uses *mean-level coding*, such that each coefficient captures the percent price premium paid by a nonlocal buyer within that specific period, conditional on property characteristics and fixed effects. Under this coding scheme, statistical significance of a coefficient indicates that the nonlocal premium in that period is significantly different from zero. Confidence intervals are shown at the 95% level; estimates for which the zero line lies outside the bands are statistically significant at $p < 0.05$. A joint F-test rejects the null hypothesis that the nonlocal premium is zero across all time periods ($F = 7.95$, $p < 0.01$). Additionally, the nonlocal premium in 2000-2003 differs significantly from both 2020-2023 ($F = 8.34$, $p < 0.01$) and 2008-2011 ($F = 6.3$, $p < 0.05$). We fail to reject the null that any period from 2012 onward is statistically equal to the 2008-2011 baseline.

stantially attenuated in more urban, liquid real estate markets by examining geographic heterogeneity across Hawaiian counties. [9]

Figure 4 shows that Honolulu county exhibits a minimal nonlocal premium of just 0.8%, which is not statistically different from zero. In contrast, Maui commands the highest premium at 11.5%, almost twice the premium in Hawaii County (6.2%) and Kauai (6.4%). A joint F-test strongly rejects the null that the nonlocal premium is zero across all counties ($F = 24.8$, $p < 0.01$). Pairwise F-tests show that the premium in Honolulu is statistically different from those in Hawaii, Kauai, and Maui (all $p < 0.01$). Notably, wealth differences cannot easily explain this pattern: Honolulu attracts

[9] Our conclusions are unchanged when Census-defined urban areas are used as an alternative measure of urbanization. County-level differences are of primary interest given their economic and policy relevance in Hawaii.

high-income nonlocal buyers yet exhibits the smallest premium, consistent with market liquidity and information quality rather than buyer resources determining the gradient.

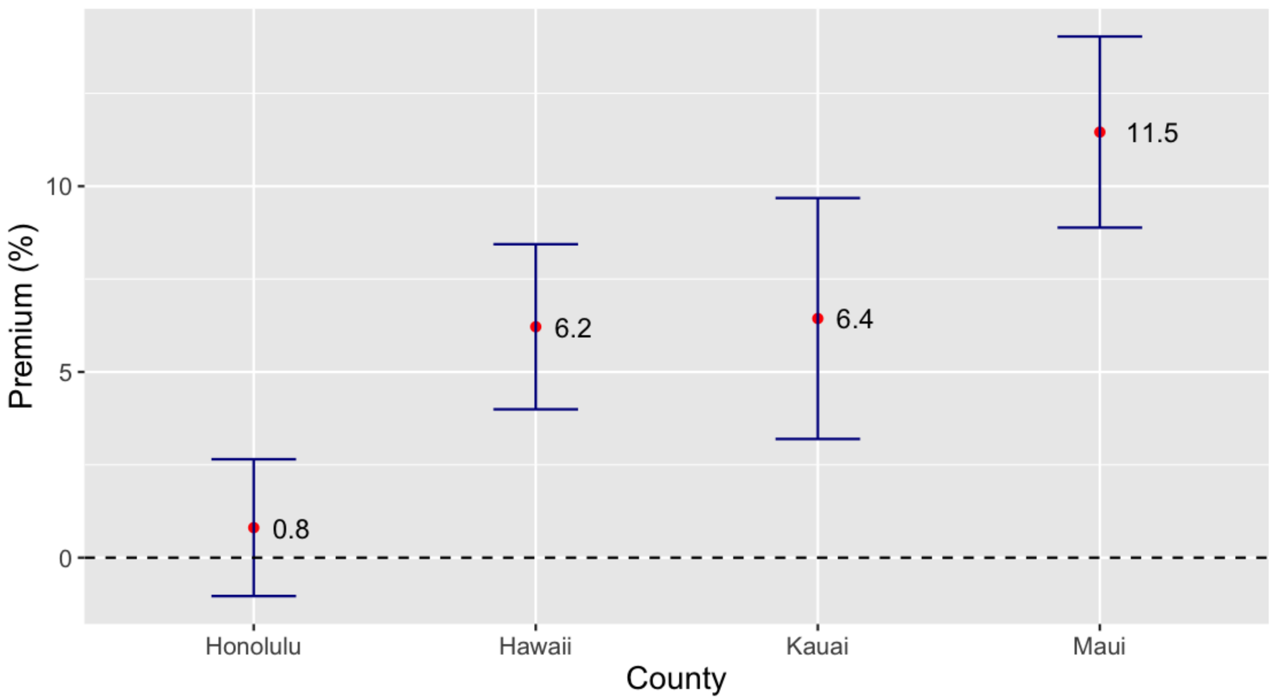

Figure 4: This figure plots the estimated coefficients from Equation 1, where the independent variable of interest is the interaction between *NonLocal* and *County*, where $County = \{Honolulu, Hawaii, Kauai, Maui\}$. The estimation uses *mean-level coding*, such that each coefficient captures the percent price premium paid by a nonlocal buyer within that specific county, conditional on property characteristics and fixed effects. Under this coding scheme, statistical significance of a coefficient indicates that the nonlocal premium in that county is significantly different from zero. Confidence intervals are shown at the 95% level; estimates for which the zero line lies outside the bands are statistically significant at $p < 0.05$. A joint F-test rejects the null hypothesis that the nonlocal premium is zero across all counties($F = 24.8$, $p < 0.01$). Pairwise F-tests show that the nonlocal premium in Honolulu and Maui is statistically different from all other counties. We fail to reject the null that the nonlocal premium in Hawaii County and Kauai are statistically equal.

6 Nonlocals and Economic Outcomes

We now address whether changes in the composition of marginal buyers affect local economic outcomes. For instance, there is a widely held view that outside buyers have contributed to rising housing costs in Hawai'i (Grassroot Institute of Hawaii, 2022), yet there has been little systematic empirical evidence documenting this relationship. Figure A3 shows that nonlocal buyer participation at the county level shows no persistent long-run secular trend across any county, though cyclical variation is evident. Importantly, no county exhibits a sustained structural upward or downward shift over the two-decade period. Over the same period, statewide home prices rose sharply (Figure A1), with several ZIP codes exhibiting appreciation rates greater than 100% (Figure A5). This apparent disconnect motivates our subsequent ZIP–level analysis.

Specifically, we are interested in the following outcomes: sale prices, employment, and migration. Li et al. (2024) develop a theoretical framework linking foreign real estate capital inflows to local labor market outcomes through two opposing channels. The first is a housing net worth channel: rising foreign demand pushes up house prices, which in turn stimulates household spending on locally supplied non-tradable goods and raises employment. The second is a displacement channel: higher housing costs crowd out lower-income residents, contracting local consumption demand and depressing non-tradable employment. In their empirical analysis of Chinese real estate purchases in California, Li et al. (2024) find that the net worth channel dominates on average—foreign capital inflows raise house prices and expand local employment, particularly in the non-tradable sector—though these gains are accompanied by displacement of lower-income residents.

6.1 Cross-Sectional Evidence

Following Li et al. (2024), our estimating equation is given by:

$$\log(Y_{zt}) = \beta NonLocal_{zt} + \gamma X_{z,early} + \lambda_{ct} + \epsilon_{zt}, \tag{3}$$

where Y_{zt} denotes the outcome variable of interest in ZIP z in year t. The key explanatory variable, $NonLocal_{zt}$, measures the share of non-local demand in a ZIP–year, constructed in two ways: (i) the percentage of transactions involving non-local buyers, and (ii) the percentage of total transaction value accounted for by non-local buyers. The vector $X_{z,early}$ includes ZIP-level controls obtained from the U.S. Census Bureau and Hawai'i statewide GIS program. Controls include population, population density, education (the share of residents with a bachelor's degree), and median household income, all measured at the start of the sample period (year 2000). We also include an indicator for post-secondary institution presence within two miles of the ZIP centroid, as well as pre-sample changes in the dependent variable to account for differential pre-existing trends. Our identification strategy exploits within-county, cross-ZIP variation in exposure to non-local demand while controlling for time-varying county-level economic conditions using county-by-year fixed effects, λ_{ct}.

Table 6 estimates Equation 3 for three outcomes: median ZIP home sale prices, total employment, and the number of business establishments. We find strong support for the housing net worth channel. A one percentage point increase in nonlocal transaction share is associated with a 0.6 percent

Table 6: Nonlocal Share, Home Price, and Employment Outcomes

This table presents estimates for Equation 3 at the ZIP–year level where the dependent variables are the natural logs of median sale price (Columns 1–2), total number of paid employees (Columns 3–4), and total number of business establishments (Columns 5–6). Employment and establishment data is sourced from the Census Bureau's County Business Patterns (CBP) ZIP Business Patterns files. The explanatory variables of interest are nonlocal transaction share (odd columns), measured as the percentage of residential transactions in a ZIP–year made by out-of-state buyers, and nonlocal value share (even columns), measured as the percentage of total dollar volume attributable to out-of-state buyers. All specifications include county × year fixed effects and ZIP-level controls for population, population density, median household income, bachelor's degree share, college presence, and pre-period trends in outcome variables. The sample includes ZIP-year observations with at least five yearly transactions. Standard errors clustered at the ZIP level are reported in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

	Home Price		Employment		Establishments	
	(1)	(2)	(3)	(4)	(5)	(6)
NonLocal (%)	0.006***	0.006***	0.015**	0.013**	0.010**	0.009**
	(0.002)	(0.001)	(0.006)	(0.005)	(0.005)	(0.004)
Controls	Y	Y	Y	Y	Y	Y
County × Year FE	Y	Y	Y	Y	Y	Y
R^2	0.745	0.753	0.670	0.672	0.739	0.740
Adjusted R^2	0.728	0.737	0.648	0.651	0.722	0.723
N	1,672	1,672	1,632	1,632	1,672	1,672

increase in ZIP home prices, 1.5 percent higher employment, and 1.0 percent more business establishments. Table A4 shows that non-tradable sector employment drives these results. A one percentage point increase in nonlocal demand is associated with a 2.0–2.4 percent increase in non-tradable sector employment. This result is robust to the exclusion of the construction sector employment numbers. The estimates for tradable sector employment are statistically significant and negative. All results are comparable using the transaction value measure.

Next, we test whether nonlocal demand displaces local residents using IRS Statistics of Income data. Table A5 reports regression estimates for tax return outcomes over 2005–2022. We find no meaningful relationship between nonlocal buyer presence and total returns filed, low-income returns, high-income returns, or average adjusted gross income. If displacement were operative, we would expect declining total filings—particularly among low-income households—in ZIP codes with greater nonlocal penetration. The absence of such effects provides no support for the displacement channel formalized in Li et al. (2024).

6.2 Long Difference Evidence

Housing is a long-lived asset whose price reflects expectations about future fundamentals and resale values (Poterba, 1984). To examine whether nonlocal demand is capitalized over longer horizons, we thus follow Burke and Emerick (2016); Acemoglu et al. (2020); Babina et al. (2024) and estimate a long-difference specification. We collapse transaction-level data to the ZIP code level, compute outcome of interest and nonlocal buyer shares within two multi-year endpoint windows, and regress the change in outcomes on the change in nonlocal presence:

$$\Delta \log(Y_z) = \beta \, \Delta NL_z + \gamma X_{z,\text{early}} + \lambda_c + \varepsilon_z, \tag{4}$$

where $\Delta \log(Y_z)$ is the change in the logarithm of outcome Y in ZIP code z between the early and late endpoint periods, and ΔNL_z is the corresponding change in nonlocal buyer participation, measured as either the change in the transaction share or the value share of nonlocal purchases (in percentage points).[10] The vector $X_{z,\text{early}}$ includes controls measured around the initial endpoint: population, population density, median household income, the share of residents holding a bachelor's degree, vacancy rate, and pre-sample trends in the outcome variable.

The key advantage of the long-difference design is that differencing eliminates time-invariant ZIP code characteristics — including unobserved amenities, geography, and baseline desirability — that may jointly determine both price levels and nonlocal buyer sorting in the cross-section. County fixed effects λ_c account for any unobserved county-level trends. Identification only comes from within county variation, eliminating any concerns of time-trending unobservables at the county level. One potential drawback of using county fixed effects is that it could absorb most of the variation of interest in our buyer composition variable. Figure A6 visualizes the variation that is used in our identification strategy. Within county, some ZIP codes have had increased nonlocal presence while others have had shrinking nonlocal presence.

Table 7 shows that a one percentage point increase in the nonlocal transaction share over roughly two decades is associated with approximately 0.8–1.4 percent higher growth in median ZIP home prices. Tables A6 and A7 show that the corresponding estimates for employment, establishments, and income measures are statistically insignificant.

6.3 Reconciling the Two Designs

The two empirical designs yield a consistent finding on housing prices but diverge on employment and establishment outcomes, and the source of that divergence is informative about what each design identifies.

[10] Our baseline uses 2000–2005 and 2018–2023 as endpoints, adjusted for some outcomes depending on data availability as noted in the regression tables. Results are robust to alternative endpoint choices.

Table 7: Nonlocal Demand and Home Prices: Long Differencing

This table presents estimates from Equation 4, where the dependent variable is the change in log median sale price (inflation adjusted) between 2000–2005 and 2018–2023 at the ZIP level. The independent variable of interest is the corresponding change in nonlocal buyer participation, measured either as the percentage of transactions involving nonlocal buyers (transaction share) or as the percentage of total transaction value accounted for by nonlocal buyers (value share), expressed in percentage points. All specifications include county fixed effects and baseline ZIP-level controls measured around the early period: population, population density, education, median income,and pre-sample changes in outcome variable (1995-2000). Columns (1) and (2) use transaction share, while columns (3) and (4) use transaction value share. Columns (2) and (4) weight observations by the square root of the number of transactions in 2000–2005. Heteroskedasticity-robust standard errors are reported in parentheses. * p < 0.10, ** p < 0.05, *** p < 0.01.

	Transaction Share		Value Share	
	(1)	(2)	(3)	(4)
Δ NonLocal (%)	0.008***	0.014***	0.008***	0.004
	(0.002)	(0.004)	(0.002)	(0.004)
Controls	Y	Y	Y	Y
County FE	Y	Y	Y	Y
Weighted	N	Y	N	Y
R^2	0.381	0.460	0.378	0.362
Adjusted R^2	0.289	0.380	0.286	0.266
N	78	78	78	78

In the cross-sectional specifications of Equation 3, higher nonlocal demand is associated with higher house prices, greater total employment, and more business establishments, with the employment effects concentrated in the non-tradable sector. We also find a negative association with tradable sector employment. On the surface, these associations are superficially consistent with the housing net worth channel of Li et al. (2024). However, Equation 3 relies on cross-ZIP variation conditional on county-by-year fixed effects and baseline controls, it does not absorb time-invariant ZIP characteristics. We therefore interpret these estimates as reflecting the fact that ZIP codes with persistently higher nonlocal buyer presence tend to have higher employment, of which more is in nontradable sectors and less in tradable sectors — a cross-sectional sorting artifact. This interpretation is supported by the robustness check in Table A8, which adds ZIP fixed effects to Equation 3: the price association survives, while the employment and establishment associations become statistically insignificant.

The long-difference design of Equation 4 addresses this directly. By construction, differencing eliminates time-invariant ZIP characteristics — unobserved amenities, geography, and baseline desirability — that may jointly govern nonlocal buyer sorting and employment levels in the cross-section. Under this design, the employment, non-tradable sector, and establishment associations all disappear, corroborating the view that the cross-sectional patterns are artifacts of omitted time-invariant confounders rather than evidence of genuine economic stimulus from nonlocal demand.

Across both designs, the only conclusion that survives is that nonlocal demand is reliably capitalized into housing prices. We find no support for the displacement channel in either specification: tax filings across all income groups are unresponsive to nonlocal demand, providing consistent evidence against the displacement mechanism formalized in Li et al. (2024). The cross-sectional associations with employment and non-tradable activity do not survive the introduction of ZIP fixed effects or the long-difference design, and as such, cannot be treated as establishing a robust housing net worth channel operating through local economic outcomes.

7 Economic Magnitude

Our long-difference estimates establish that nonlocal demand has a statistically significant effect on home prices. We now assess the economic magnitude of this effect and ask whether it can account for any meaningful portion of Hawai'i's housing cost increase. The following exercise should be interpreted as an illustrative lower-bound decomposition. For each ZIP code z, we take three objects from our data:

1. $\hat{\beta}_{LD}$: the estimated long-difference coefficient from Equation 4. This tells us what percentage the home prices rose over two decades, for a 1 percent point increase in nonlocal presence or vice versa. We use $\hat{\beta}_{LD} = 0.008$.

2. ΔNL_z: the actual observed change in nonlocal transaction share (in percentage points) in ZIP code z between 2000–2005 and 2018–2023.

3. $\Delta \log P_z$: the actual change in log median sale prices in ZIP code z over the same period.

The price growth that would result from the change in nonlocal buyer composition (numerator) is expressed as a fraction of the total price growth (denominator):

$$\theta_z = \frac{\hat{\beta}_{LD} \times \Delta NL_z}{\Delta \log P_z}. \tag{5}$$

A value of $\theta_z = 0.05$ means that 5 percent of the ZIP's price growth is explained by the change in its nonlocal buyer share; the remaining 95 percent reflects all other forces. Note that θ_z can be negative in ZIP codes where nonlocal participation declined.

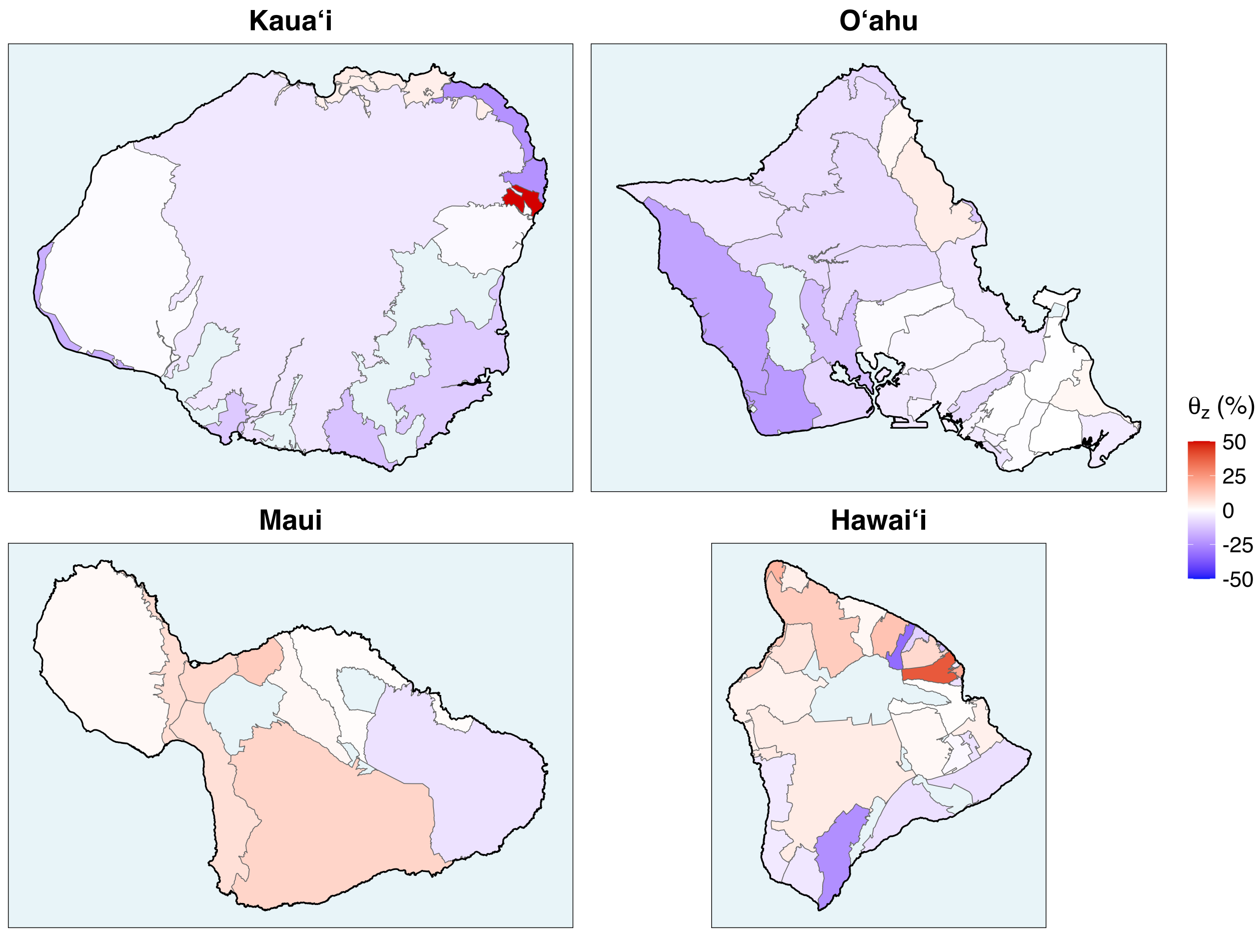

Figure 5: Share of ZIP-level price appreciation attributable to changes in nonlocal demand (θ_z) between 2000–2005 and 2018–2023 for the major islands. Red shading indicates ZIP codes with a positive predicted contribution to price growth, while blue shading indicates ZIP codes with a negative predicted contribution. White indicates near-zero contribution.

Figure 5 maps θ_z across ZIP codes, shading each area by the share of its total price growth that our estimates attribute to the shift in nonlocal buyer composition. The pattern varies across islands. Maui and Hawai'i have the largest predicted nonlocal demand contribution to price growth, with several ZIP codes exceeding 25 percent. Kaua'i and O'ahu are predominantly negative, implying that changes in nonlocal demand, if anything, should dampen price growth in these areas. Notable exceptions include the northeast Kaua'i coast and isolated windward (east) O'ahu ZIP codes, where θ_z turns positive. Figure A4 summarizes these distributions by county: the median θ_z is close to zero everywhere, the interquartile range is narrow, and outliers arise when there are large swings in nonlocal share. Figure A5 contextualizes these estimates: residential properties appreciated substantially across all islands over the period, and even in ZIP codes where nonlocal presence contracted, prices rose markedly.

Overall, this qualitative exercise suggests that nonlocals may disproportionately bear public scrutiny relative to their actual contribution to price appreciation. In specific ZIP codes — particularly on Maui and Hawai'i Island, where nonlocal penetration increased materially — the predicted price contribution is nontrivial, and localized policy concern may be warranted. However, policies targeting nonlocal ownership alone are unlikely to materially improve affordability absent supply-side reforms that address the binding constraints on new construction.

8 Discussion and Conclusion

This paper provides a systematic examination of how buyer origin influences housing prices and local economic outcomes in Hawai'i, a market uniquely defined by its geographic isolation and high exposure to external capital. Leveraging a comprehensive transaction-level dataset spanning over two decades, we document that nonlocal buyers pay a robust per-transaction price premium of 4.2–4.8 percent for properties that are otherwise observably equivalent to those purchased by local residents. This premium is significantly larger for foreign buyers, who pay 9.8 percent more than locals, reflecting the heightened informational and institutional barriers faced by international investors.

The evidence points to a joint role for information frictions and behavioral anchoring, with wealth and the discretionary nature of nonlocal purchases playing a secondary role. Among buyers who are observably less wealthy by mortgage characteristics — those financing at high loan-to-value ratios or borrowing at or below the local median — the premium attenuates but remains positive and significant, indicating that differential financial capacity explains part but not all of the result. The finding that the premium has nearly halved since the early 2000s — coinciding with the expansion of digital real estate platforms — suggests that declining search costs have moved the market closer to pricing efficiency. The gradient across buyer origins, with high-HPI and foreign buyers paying systematically larger premia than buyers from lower-priced markets, supports an anchoring mechanism in which home-market prices serve as reference points. Our findings are difficult to reconcile with a pure preference-based account, as the premium persists even when excluding high-end segments and shoreline properties.

While these effects are pronounced at the transaction level, their implications for broader economic outcomes are limited. In cross-sectional specifications, ZIP codes with higher nonlocal buyer shares exhibit higher prices, employment, and business activity. However, these relationships do not survive the inclusion of ZIP fixed effects or long-difference designs. Once time-invariant local characteristics are absorbed, only the price effect remains. Over two decades, a one-percentage-point increase in nonlocal share is associated with modest price appreciation of 0.8–1.4 percent. Back-of-the-envelope calculations indicate that, for the median ZIP code, changes in buyer composition explain little of overall price growth. Larger effects are concentrated in parts of Maui and Hawai'i Island, where nonlocal penetration increased substantially, while Kaua'i and O'ahu exhibit limited or offsetting effects.

These findings also inform the policy debate. While nonlocal buyers contribute to higher prices in specific markets, their role in driving long-run affordability is limited and spatially uneven. Targeted demand-side interventions—such as vacancy taxes or short-term rental restrictions—may provide localized relief but are unlikely to meaningfully improve housing affordability at the aggregate level. Instead, concerns about nonlocal demand, while understandable, should be interpreted in the context of broader structural constraints on housing supply. Recent state-level reforms aimed at reducing regulatory barriers to small-scale infill development and the adaptive reuse of underutilized commercial space represent welcome steps in this direction.[11] More broadly, our results highlight a distinction between micro-level pricing distortions and macro-level housing outcomes: informational frictions and behavioral biases can produce persistent transaction-level premia, yet these effects do not scale into large or persistent changes in aggregate housing dynamics.

Some limitations merit discussion. First, our analysis focuses on residential transactions and does not examine rental or commercial markets, where nonlocal ownership may have distinct implications for affordability and displacement. Second, we do not directly observe buyer intent—such as whether purchases are for investment, second homes, or primary residence—which may shape willingness to pay. Third, our classification relies on buyers' reported addresses at the time of purchase. To the extent that recent migrants are classified as local buyers despite retaining nonlocal reference prices or informational disadvantages, our estimates are likely attenuated. Finally, while our long-difference

[11] In 2024, Hawai'i enacted two supply-oriented reforms. Act 39 requires counties to permit at least two accessory dwelling units on residentially zoned lots, subject to infrastructure and short-term rental safeguards. Act 37 mandates that counties allow multi-unit residential development and adaptive reuse of existing commercial structures in business-zoned areas. While these laws expand the legal envelope for housing supply, their effectiveness will depend on county-level implementation, including parking requirements, lot-size rules, development fees, and infrastructure capacity (UHERO, 2025).

design and ZIP fixed effects absorb time-invariant confounders, we cannot fully rule out that price appreciation and nonlocal demand are jointly determined. If appreciating markets attract nonlocal buyers, our long-difference estimates may overstate the causal effect of buyer composition on prices.

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A Appendix

A.1 Additional Figures

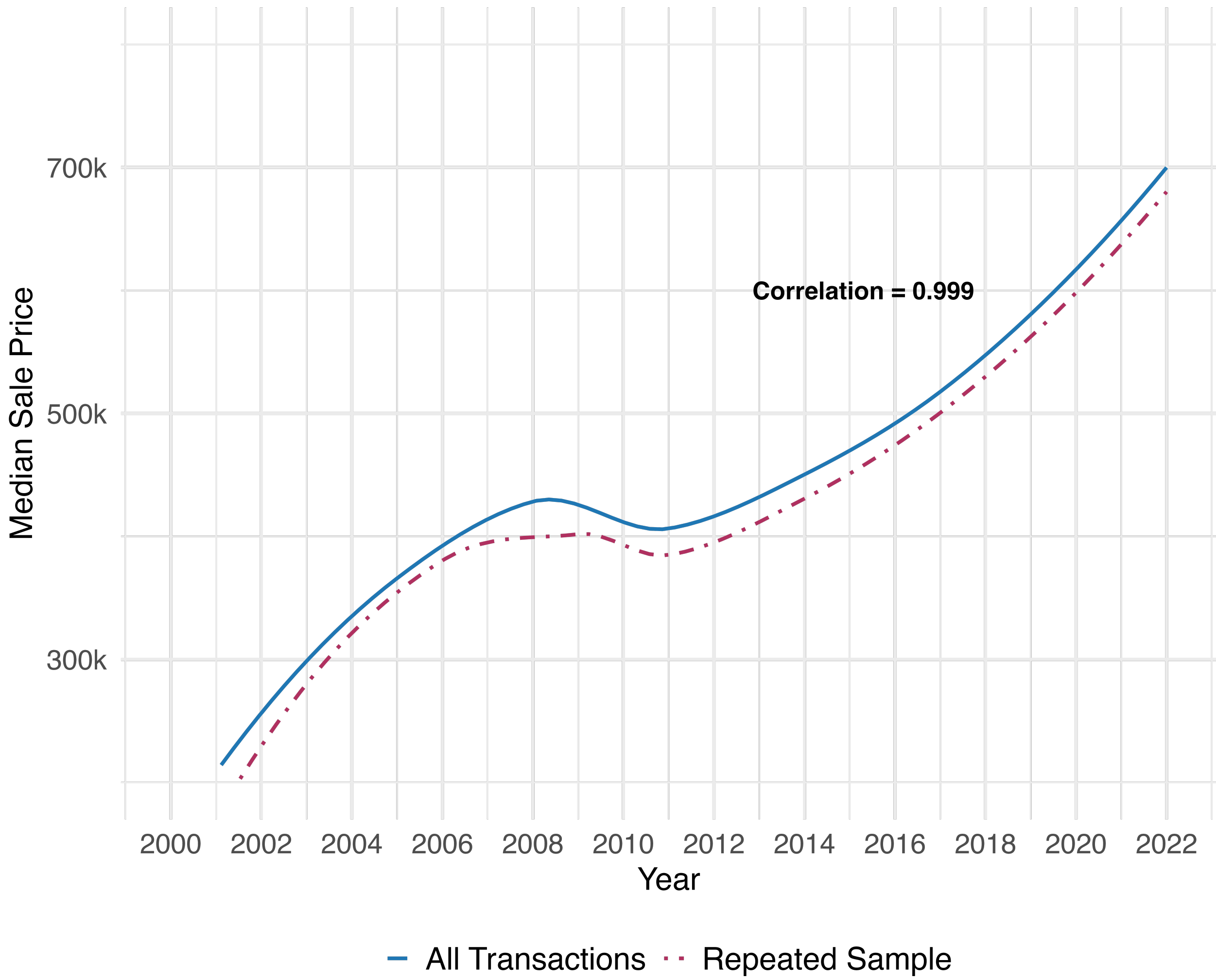

Figure A1: Median nominal price trend for the repeated sales sample from 2000 to 2022 closely resemble the full sample.

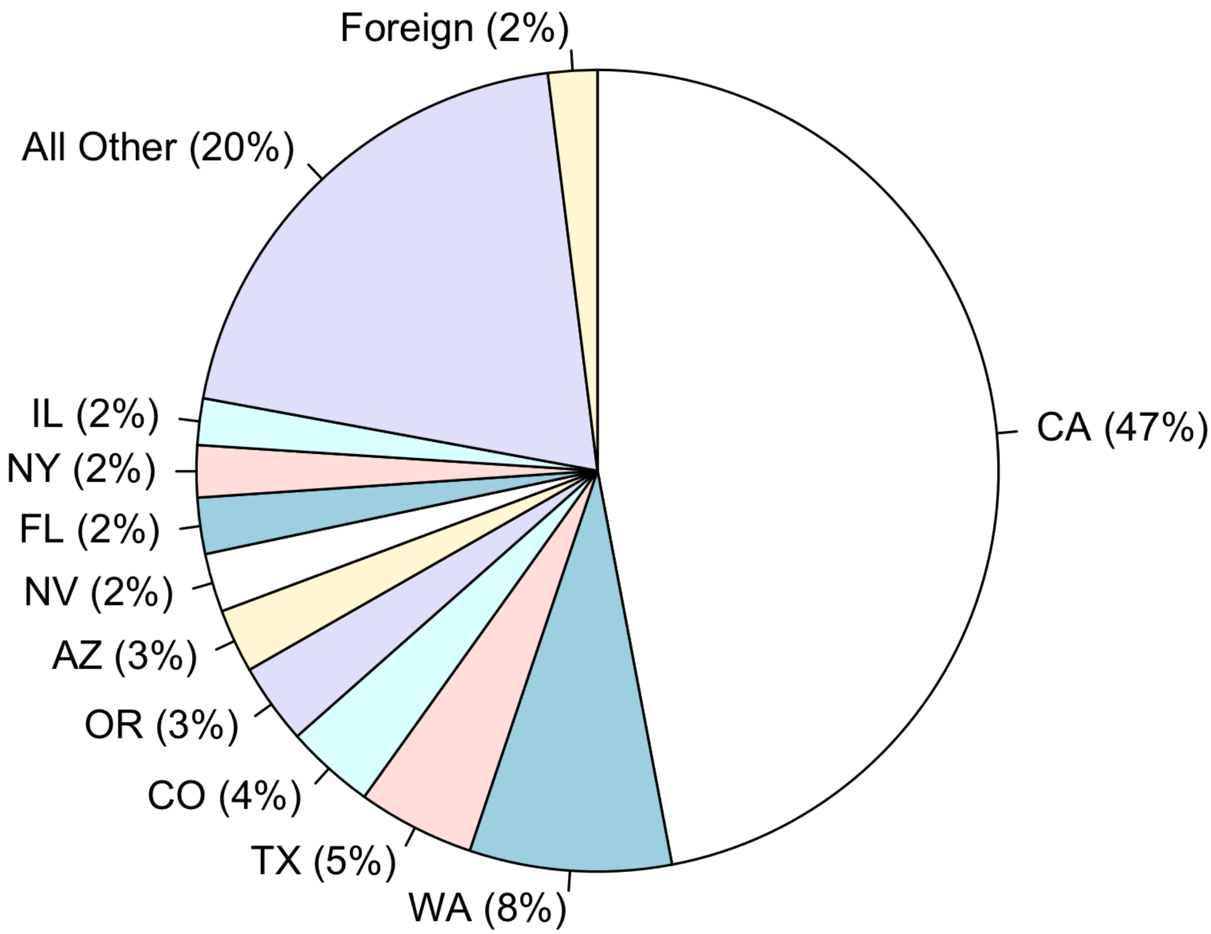

Figure A2: Composition of nonlocal homebuyers by origin. The figure reports the share of nonlocal housing transactions accounted for by the top ten mainland US states, with remaining states grouped as 'All Other' and foreign buyers shown separately.

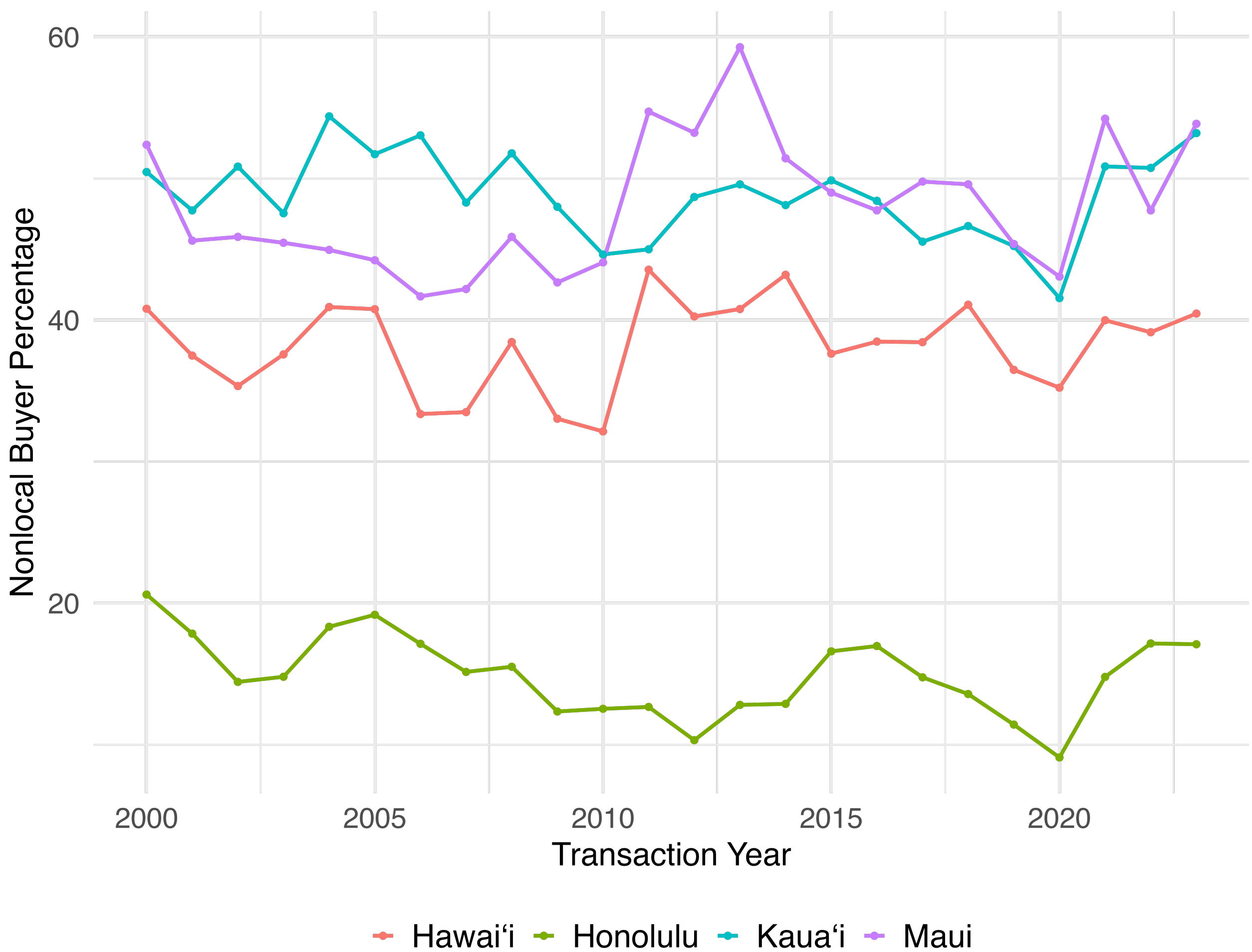

Figure A3: Trends in nonlocal buyer participation by county over time.

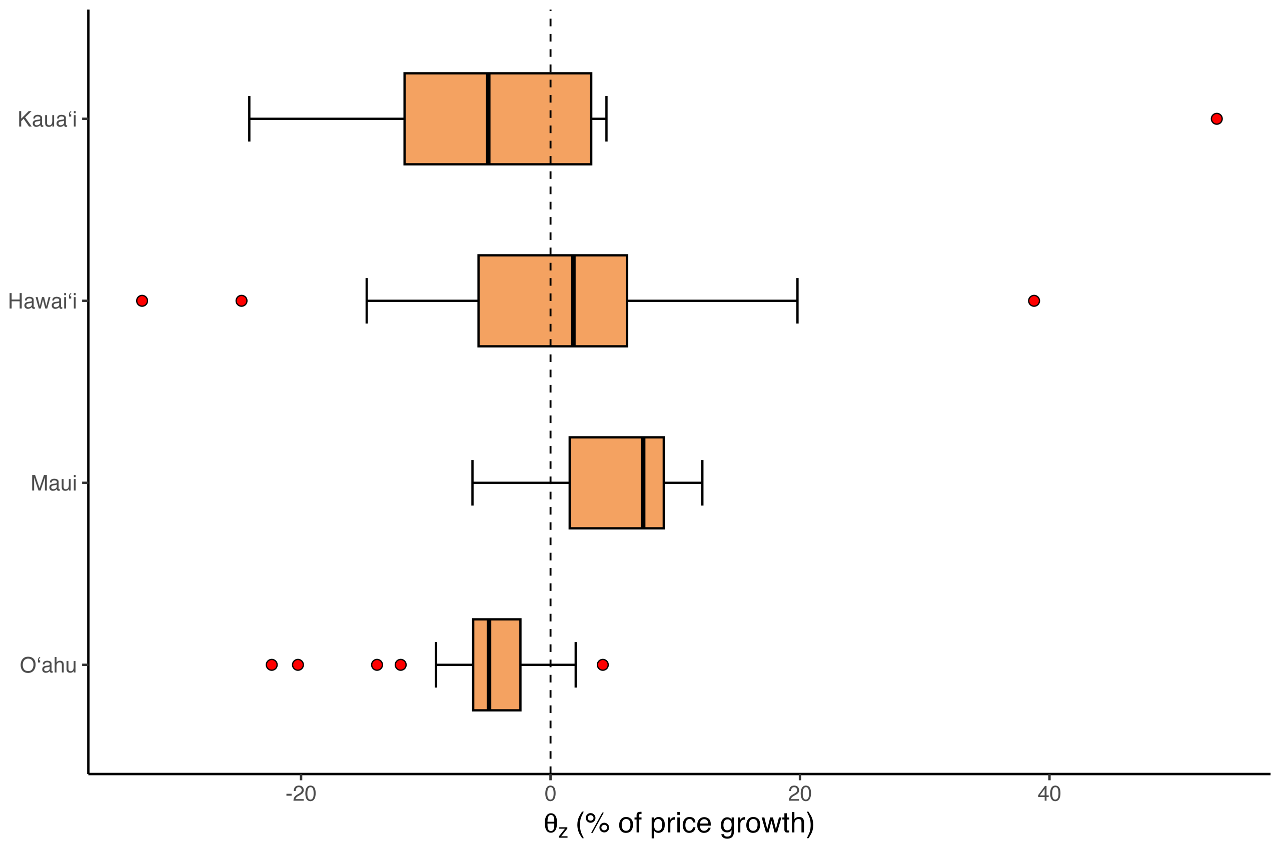

Figure A4: Distribution of θ_z by county. Each box shows the interquartile range (25th–75th percentile) of the share of ZIP-level price appreciation attributable to changes in nonlocal demand; the vertical line within each box marks the median. Whiskers extend to 1.5 times the interquartile range; red dots indicate outliers. The dashed vertical line at zero separates ZIP codes where nonlocal inflows contributed to price growth (right) from those where nonlocal outflows dampened it (left).

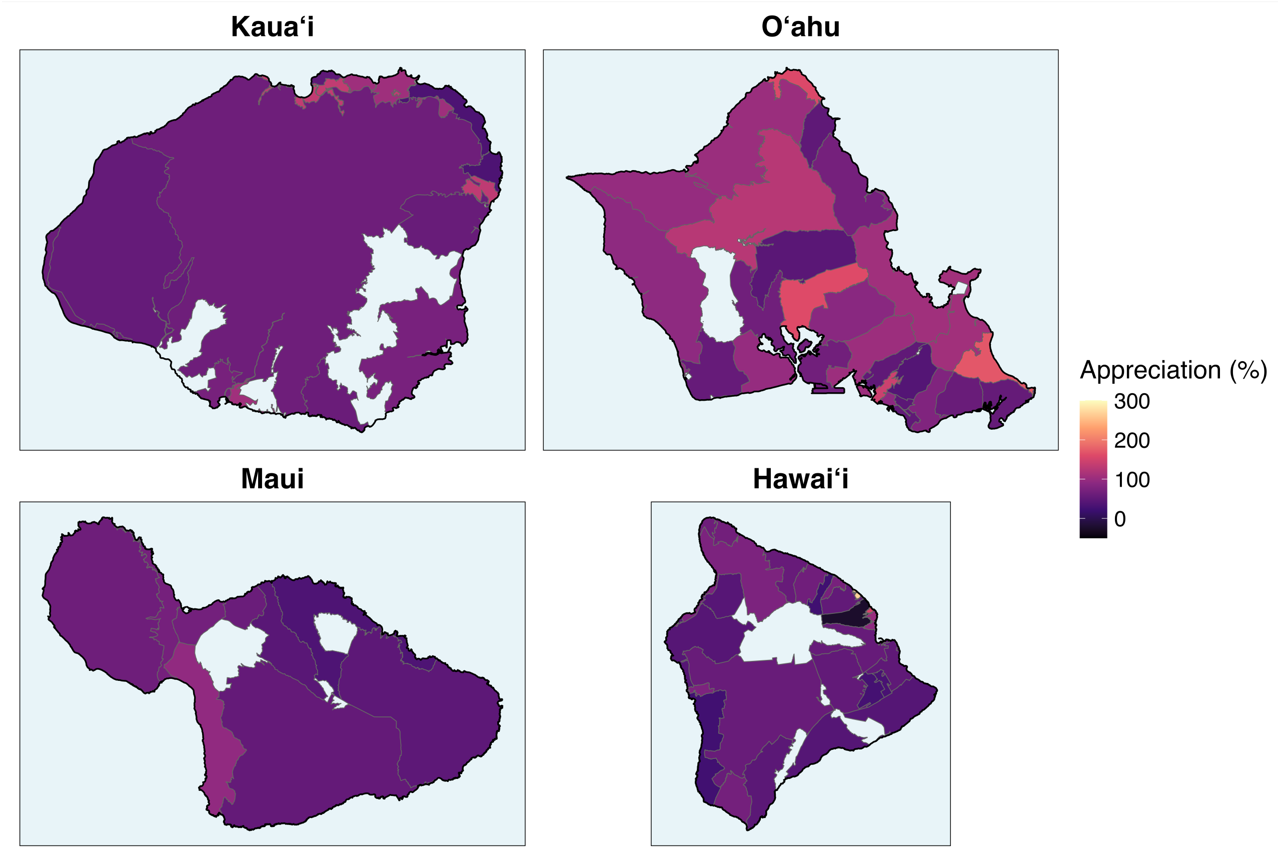

Figure A5: Inflation adjusted price changes (%) by ZIP code between 2000–2005 and 2018–2023 for the major islands. Each ZIP code is shaded by the percentage change in median residential sale price between the two endpoint periods.

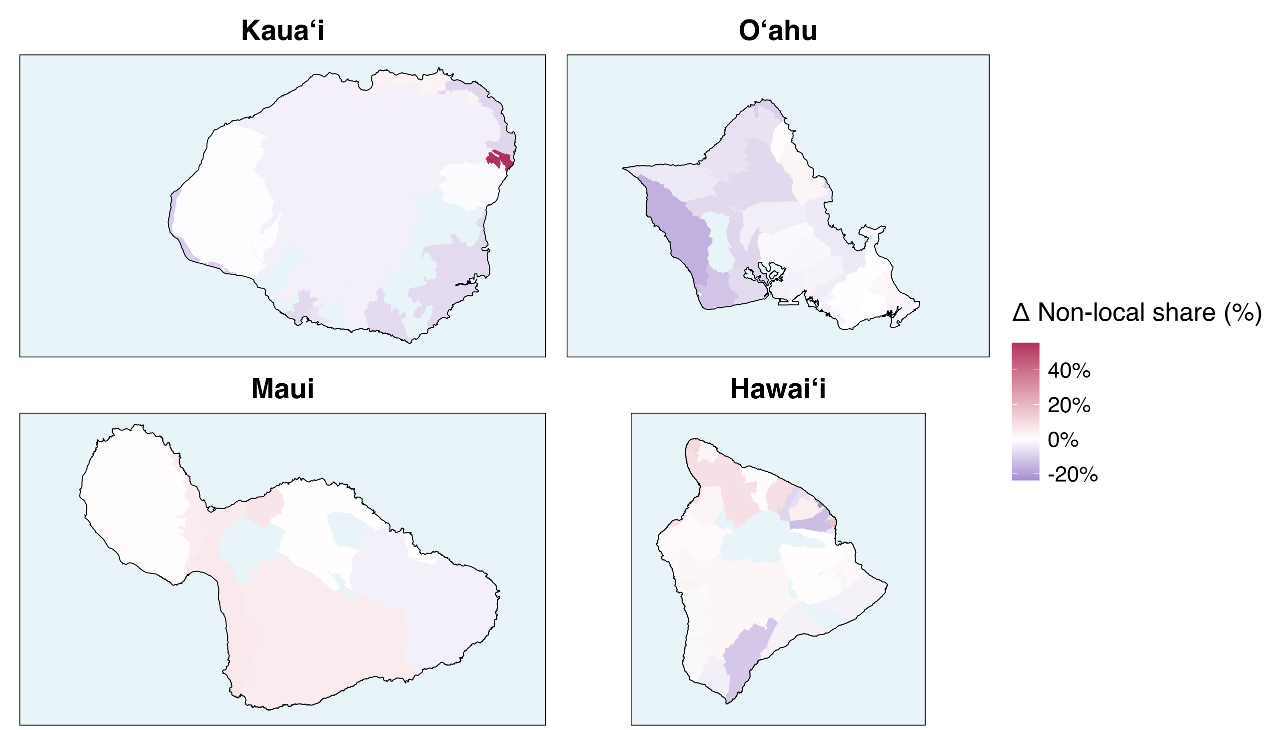

Figure A6: Geographic variation in changes in nonlocal buyer presence across ZIP codes. Colors represent the change in the share of residential transactions involving nonlocal buyers between 2000–2005 and 2018–2023.

A.2 Additional Tables

Table A1: Top ten states by home price

Rank	1	2	3	4	5	6	7	8	9	10
State	MA	DC	NY	RI	WA	NJ	ME	VT	OR	MD

Note: Top ten states by average FHFA home price index, 2000–2022. Seven are also among the top ten in 2025 median single-family home prices (Zillow), while the rest (MD, VT, and ME) rank within the top twenty. Overall, this set closely mirrors the states identified by latest (as of 2025) median house price rankings.

We report additional results using alternative standard error adjustments to account for potential spatial and temporal dependence in the regression residuals (Table A2). We begin with specifications that use a two-way clustering adjustment. Spatially, we cluster our observations at the ZIP level. Depending on the nature of serial correlation, we show our results are robust to temporal adjustments at the year, year–quarter, and year–month levels. These adjustments allow for arbitrary correlation in the error structure among transactions occurring within the same ZIP and the same time period. We are therefore accounting for localized demand or supply shocks that persist over short horizons, or common market-wide forces that may affect transactions contemporaneously. We also report standard errors corrected for spatial dependence (Conley, 1999). Similar to clustered standard errors, which allow for within-group dependence, Conley standard errors allow for arbitrary correlation in the error structure among observations in specified geographic proximity. This approach is particularly appropriate in housing markets, where unobserved neighborhood-level shocks or spatial spillovers may induce correlation across nearby transactions that does not align with administrative boundaries such as ZIP codes. Our main results remain robust to these corrections.

Table A2: Main Results Under Alternative Standard Errors

This table presents estimates for Equation 1 where the dependent variable is log(Price). Each column reports the estimated nonlocal buyer premium under alternative standard error specifications. Columns (1)–(3) use a double clustering adjustment at the ZIP and transaction time level, where time is defined at the year, year–quarter, and year–month levels, respectively. Columns (4)–(7) report Conley standard errors with distance cutoffs of 2, 4, 6, and 8 miles. Standard errors are reported in parentheses. * p < 0.10, ** p < 0.05, *** p < 0.01.

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
NonLocal	0.042***	0.042***	0.042***	0.042***	0.042**	0.042**	0.042*
	(0.011)	(0.010)	(0.010)	(0.016)	(0.019)	(0.021)	(0.023)
Square Footage	Y	Y	Y	Y	Y	Y	Y
House Age	Y	Y	Y	Y	Y	Y	Y
Elevation	Y	Y	Y	Y	Y	Y	Y
$Z \times Y \times B \times P \times C$	Y	Y	Y	Y	Y	Y	Y
SE	Z-Y	Z-YQ	Z-YM	2 mi	4 mi	6 mi	8 mi
R^2	0.957	0.957	0.957	0.957	0.957	0.957	0.957
Adjusted R^2	0.852	0.852	0.852	0.852	0.852	0.852	0.852
N	341,681	341,681	341,681	341,681	341,681	341,681	341,681

Table A3: Nonlocal Premium: Excluding Upper Price Tails

This table presents estimates for Equation 1 where the dependent variable is log(Price). Each column restricts the sample to properties with sales prices less than or equal to the indicated percentile of the statewide price distribution. Column (1) includes properties priced below the 90th percentile; Column (2) below the 80th percentile; Column (3) below the 70th percentile; Column (4) below the 60th percentile; and Column (5) below the 50th percentile. Standard errors, clustered at the ZIP level, are reported in parentheses. * p < 0.10, ** p < 0.05, *** p < 0.01.

	(1)	(2)	(3)	(4)	(5)
	$\leq$ 90th pct.	$\leq$ 80th pct.	$\leq$ 70th pct.	$\leq$ 60th pct.	$\leq$ 50th pct.
NonLocal	0.033***	0.029***	0.026***	0.023***	0.023***
	(0.008)	(0.008)	(0.008)	(0.008)	(0.007)
Square Footage	Y	Y	Y	Y	Y
House Age	Y	Y	Y	Y	Y
Elevation	Y	Y	Y	Y	Y
$Z \times Y \times B \times P \times C$	Y	Y	Y	Y	Y
R^2	0.945	0.935	0.925	0.918	0.911
Adjusted R^2	0.815	0.788	0.759	0.734	0.712
N	307,548	273,690	239,177	205,920	170,861

Table A4: Nonlocal Share and Sectoral Employment Outcomes

This table presents estimates for Equation 3 at the ZIP–year level where the dependent variables are the natural logs of total employees in the nontradable sector (Columns 1–2), the tradable sector (Columns 3–4), and the nontradable sector excluding construction (Columns 5–6). Employment counts are constructed from the Census Bureau's County Business Patterns (CBP) ZIP Business Patterns files by converting establishment size bins into employment using midpoint approximations and aggregating to the ZIP–year level. Industries are classified into tradable and nontradable sectors based on four-digit NAICS codes following the methodology in Mian et al. (2013). The explanatory variables of interest are nonlocal transaction share (odd columns), measured as the percentage of residential transactions in a ZIP–year made by out-of-state buyers, and nonlocal value share (even columns), measured as the percentage of total dollar volume attributable to out-of-state buyers. All specifications include county × year fixed effects and ZIP-level controls for population, population density, median household income, bachelor's degree share, college presence, and pre-period trends in outcome variables. The sample includes ZIP–year observations with at least five yearly transactions. Standard errors clustered at the ZIP level are reported in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

	Nontradable		Tradable		NT excl. Const.	
	(1)	(2)	(3)	(4)	(5)	(6)
NonLocal (%)	0.024***	0.020***	−0.020**	−0.013*	0.024***	0.020***
	(0.007)	(0.006)	(0.008)	(0.007)	(0.007)	(0.006)
Controls	Y	Y	Y	Y	Y	Y
County × Year FE	Y	Y	Y	Y	Y	Y
R^2	0.665	0.665	0.391	0.383	0.665	0.665
Adjusted R^2	0.643	0.643	0.339	0.329	0.643	0.643
N	1,637	1,637	1,280	1,280	1,637	1,637

Table A5: Nonlocal Share and IRS Outcomes

This table presents estimates for Equation 3 at the ZIP–year level where the dependent variables are the natural logs of total tax returns (Columns 1–2), low-income tax returns (Columns 3–4), high-income tax returns (Columns 5–6), and average adjusted gross income (Columns 7–8). Total returns denote the number of individual income tax returns filed in a ZIP–year. Low-income returns correspond to filings with adjusted gross income below \$50,000, while high-income returns correspond to filings with adjusted gross income of at least \$50,000. Average adjusted gross income is constructed as total adjusted gross income divided by total returns. All tax return data are drawn from the IRS Statistics of Income (SOI) Individual Income Tax ZIP Code Data for 2005–2022. The explanatory variables of interest are nonlocal transaction share (odd columns), measured as the percentage of residential transactions in a ZIP–year made by out-of-state buyers, and nonlocal value share (even columns), measured as the percentage of total dollar volume attributable to out-of-state buyers. All specifications include county × year fixed effects and ZIP-level controls for population, population density, median household income, bachelor's degree share, college presence, and pre-period trends in outcome variables. The sample includes ZIP–year observations with at least five yearly transactions. Standard errors clustered at the ZIP level are reported in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

	Total Returns		Low Income Returns		High Income Returns		Avg. Income	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
NonLocal (%)	−0.002	−0.001	−0.002	−0.001	−0.002	−0.002	0.001	0.001*
	(0.004)	(0.003)	(0.004)	(0.003)	(0.004)	(0.003)	(0.001)	(0.001)
Controls	Y	Y	Y	Y	Y	Y	Y	Y
County × Year FE	Y	Y	Y	Y	Y	Y	Y	Y
R^2	0.836	0.836	0.831	0.830	0.842	0.842	0.996	0.996
Adjusted R^2	0.824	0.824	0.818	0.818	0.830	0.830	0.996	0.996
N	1,037	1,037	1,037	1,037	1,037	1,037	1,037	1,037

Table A6: Nonlocal Share and Employment Outcomes: Long Differencing

This table presents estimates from Equation 4, where the dependent variables are the change in log average annual total employment (Columns 1–2) and the change in log average annual number of business establishments (Columns 3–4), between 2000–2005 and 2018–2023 at the ZIP code level. Employment is the total number of paid employees and establishments is the total number of business establishments, both drawn from the Census Bureau's County Business Patterns. The independent variable of interest is the change in nonlocal transaction share (in percentage points) over the same period. All specifications include County fixed effects and baseline ZIP-level controls: population, population density, median income, and bachelor's degree share (2000 Census). Odd columns report unweighted estimates; even columns weight observations by the square root of the number of transactions in 2000–2005. Heteroskedasticity-robust standard errors are reported in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

	Employment		Establishments	
	(1)	(2)	(3)	(4)
Δ NonLocal (%)	0.004	0.000	−0.002	−0.005
	(0.006)	(0.008)	(0.003)	(0.006)
Controls	Y	Y	Y	Y
County FE	Y	Y	Y	Y
Weighted	N	Y	N	Y
R^2	0.124	0.130	0.122	0.231
Adjusted R^2	0.025	0.032	0.023	0.144
N	80	80	80	80

Table A7: Nonlocal Share and IRS Outcomes: Long Differencing

This table presents estimates from Equation 4, where the dependent variables are the change in log average annual total tax returns (Columns 1–2), low-income returns with AGI below $50,000 (Columns 3–4), high-income returns with AGI at or above $50,000 (Columns 5–6), and average adjusted gross income (Columns 7–8), between 2005–2007 and 2018–2022 at the ZIP code level. Tax return data are drawn from the IRS Statistics of Income Individual Income Tax ZIP Code Data. The independent variable of interest is the change in nonlocal transaction share (in percentage points) between 2000–2005 and 2018–2023. All specifications include County fixed effects and baseline ZIP-level controls: population, population density, median income, and bachelor's degree share (2000 Census). Odd columns report unweighted estimates; even columns weight observations by the square root of the number of transactions in 2000–2005. Heteroskedasticity-robust standard errors, are reported in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

	Total Returns		Low Income Returns		High Income Returns		Avg. Income	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Δ NonLocal (%)	−0.003	−0.008	−0.003	−0.006	−0.003	−0.008	0.006*	0.005
	(0.003)	(0.005)	(0.004)	(0.006)	(0.004)	(0.005)	(0.004)	(0.005)
Controls	Y	Y	Y	Y	Y	Y	Y	Y
County FE	Y	Y	Y	Y	Y	Y	Y	Y
Weighted	N	Y	N	Y	N	Y	N	Y
R^2	0.176	0.256	0.287	0.346	0.240	0.205	0.542	0.569
Adjusted R^2	0.041	0.134	0.170	0.240	0.116	0.075	0.467	0.498
N	58	58	58	58	58	58	58	58

Table A8: Nonlocal Share, Home Price, and Employment Outcomes

This table presents estimates for augmented Equation 3 with ZIP fixed effects. The dependent variables are the natural logs of median sale price (Columns 1–2), total number of paid employees (Columns 3–4), and total number of business establishments (Columns 5–6). Employment and establishment data is sourced from the Census Bureau's County Business Patterns (CBP) ZIP Business Patterns files. The explanatory variables of interest are nonlocal transaction share (odd columns), measured as the percentage of residential transactions in a ZIP–year made by out-of-state buyers, and nonlocal value share (even columns), measured as the percentage of total dollar volume attributable to out-of-state buyers. All specifications include county × year fixed effects. The sample includes ZIP-year observations with at least five yearly transactions. Standard errors clustered at the ZIP level are reported in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

	Home Price		Employment		Establishments	
	(1)	(2)	(3)	(4)	(5)	(6)
NonLocal (%)	0.003***	0.002***	0.000	0.000	0.000	0.000
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.000)
ZIP FE	Y	Y	Y	Y	Y	Y
County × Year FE	Y	Y	Y	Y	Y	Y
R^2	0.911	0.910	0.988	0.988	0.994	0.994
Adjusted R^2	0.901	0.900	0.986	0.986	0.993	0.993
N	1,709	1,709	1,659	1,659	1,707	1,707