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The Impact of the Youth Housing Loan Subsidy Policy on Housing Prices: Evidence from New Taipei City

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Abstract

In August 2023, the Taiwanese government launched the New Youth Housing Loan Subsidy Policy (hereafter the New Youth Policy), which lowers the threshold for young people to buy homes through interest subsidies, higher loan ceilings, longer maturities, and extended grace periods. However, when housing supply is difficult to increase in the short run, such a demand-side subsidy may push up housing prices through a price capitalization mechanism. Taking New Taipei City as the study area, this paper uses more than one hundred thousand transaction records from the Ministry of the Interior's actual price registration system between January 2021 and September 2024, and applies a hedonic price model to construct a multiple regression. After controlling for building characteristics, location, and annual trends, the study examines the effect of the policy on the unit price of housing. The pooled specification shows that the unit price of housing was about 4.4% higher in the post-implementation period, a differential equivalent to roughly one-fifth of the conditional appreciation of the market over the same window. To test whether this differential can be attributed to the policy rather than to a market-wide trend, we exploit the NT$10 million loan ceiling to construct a difference-in-differences design in which small dwellings, whose transaction prices fall within the ceiling, are treated as the high-exposure group and large dwellings, whose prices lie far above it, serve as the low-exposure comparison group. Once the control variables are allowed to differ across the two segments, the difference-in-differences estimate is economically and statistically indistinguishable from zero, and this null result is stable across four alternative size thresholds. An event-study specification reveals a significant pre-policy divergence between the two groups, and a placebo cut-off placed in the first half of 2023 yields a significant coefficient, so the parallel-trends assumption is not supported. A split-sample estimation by unit size likewise finds no significant difference in the policy coefficient between small and large dwellings. Building age and floor area have negative effects while floor level has a positive effect, reflecting the phenomenon of higher unit prices for smaller units. We therefore conclude that housing prices in New Taipei City rose markedly over the observation window, but that the data do not provide robust evidence that this increase is attributable to price capitalization of the New Youth Policy; the 4.4% pooled differential should be read as a conditional association and most probably overstates the policy contribution. The study suggests that when the government promotes housing finance subsidies, it should simultaneously introduce supply-side measures such as increasing social housing and urban renewal, so as to fulfill the goal of housing justice.

Keywords: New Youth Housing Loan Subsidy Policy, Housing prices, Actual price registration, Hedonic price model, Multiple regression analysis

1. Introduction

In recent years, rising housing prices have made the burden of home purchase for young people an increasingly prominent social concern. In August 2023, the Taiwanese government launched the New Youth Housing Loan Subsidy Policy, hoping to assist young people in purchasing homes through preferential loan measures. Whether this policy further affects housing prices, however, remains to be tested empirically. This study therefore takes New Taipei City as its research object and investigates the impact of the New Youth Policy on housing prices.

1.1 Research Background and Motivation

In recent years, driven by a low interest rate environment, the continued concentration of population in metropolitan areas, rising housing demand, and limited land supply, housing prices in Taiwan have shown a sustained upward trend. According to the housing price index compiled by the Ministry of the Interior, the national housing price index rose from 110.61 in the first quarter of 2021 to 149.69 in the third quarter of 2024, a cumulative increase of about 35%; the New Taipei City index also rose from 108.03 to 141.11, a cumulative increase of about 31%, indicating a persistent upward trend in housing prices. The continued rise in housing prices has made the housing market a subject of intense social concern, especially for young people and first-time buyers who face greater purchasing pressure.

Rising housing prices have also gradually worsened housing affordability. According to the Ministry of the Interior's statistics on housing affordability for the third quarter of 2024, the national price-to-income ratio was 10.82, while that of New Taipei City reached 14.03—second only to Taipei City's 16.60—placing its housing affordability in the "too low" category and well above the national average. This indicates that residents of New Taipei City must devote more years of household income to purchase a home, as shown in Table 1. Against the backdrop of continuously rising housing prices, how the government can enhance purchasing capacity while maintaining housing market stability has become an important issue for housing policy.

Table 1. Price-to-income ratios for the whole nation and the six special municipalities, Q3 2024

Region	Price-to-income ratio (times)	Housing affordability level
Nationwide	10.82	Low
New Taipei City	14.03	Too low
Taipei City	16.60	Too low
Taoyuan City	9.39	Low
Taichung City	12.99	Too low
Tainan City	10.35	Low
Kaohsiung City	10.59	Low

To lower the threshold for young people to purchase homes, the government launched the New Youth Housing Loan Subsidy Policy on August 1, 2023, raising the loan ceiling to NT$10 million, extending the loan term to 40 years, and providing a grace period of up to five years together with interest subsidies, in the hope of enhancing young people's purchasing capacity by lowering financing costs. However, because housing supply is rigid in the short run, when demand increases in response to policy stimulus, the market may reflect the subsidy benefit in housing prices through a price capitalization mechanism, so that a policy originally intended to reduce the burden of home purchase may instead push up housing prices, forming a phenomenon of "the more subsidized, the more expensive." Whether the New Youth Policy indeed causes housing prices to rise has therefore become an important issue of concern to the government, academia, and society.

This study selects New Taipei City as its research object for the following reasons. First, New Taipei City is the most populous special municipality in the country, and its housing transaction volume has long ranked among the highest nationwide, providing a sufficient and stable transaction sample. Second, housing prices in New Taipei City continue to rise and its price-to-income ratio is higher than the national average, indicating strong demand and a heavy purchasing burden. Third, New Taipei City and Taipei City together form the Greater Taipei living area; affected by population inflow, transportation infrastructure, and commuting demand, its housing market is highly sensitive to changes in government policy, making it a suitable object for observing the impact of policy on housing prices. Moreover, previous studies have mostly focused on policy design or overall market trends, and few have used transaction-level actual price registration data, controlling for building characteristics and location, to examine the impact of the New Youth Policy on individual housing prices. This study therefore uses transaction data from New Taipei City and, through multiple regression analysis, tests whether the New Youth Policy has a significant effect on housing prices, so as to fill the gap in the existing literature and provide a reference for future housing policy planning and evaluation.

1.2 Research Purposes

This study aims to investigate the impact of the New Youth Policy on housing prices in New Taipei City. The research questions are as follows: (1) After the implementation of the New Youth Policy, do housing prices in New Taipei City change significantly, and what are the direction and magnitude of the effect? (2) After controlling for building characteristics and locational conditions, does the policy effect remain robust?

Accordingly, using transaction-level actual price registration data from the Ministry of the Interior, this study builds a multiple regression model and, after controlling for building characteristics, locational conditions, and annual factors, tests whether the New Youth Policy has a significant effect on housing prices and analyzes the direction and degree of that effect, providing a reference for future housing policy planning and evaluation.

1.3 Research Process

This study first formulates its research questions and motivation based on the background of the New Youth Policy and the current situation of the New Taipei City housing market, and conducts a literature review as its foundation. Second, it establishes the research framework and methodology according to the literature, setting up a policy dummy variable for the New Youth Policy and incorporating building characteristic variables and district dummy variables, with multiple regression analysis as the primary research tool. Third, it collects and organizes transaction-level actual price registration data for New Taipei City from January 2021 to September 2024 for empirical analysis, so as to examine the direction and significance of the policy's impact on housing prices. Finally, it draws conclusions and offers recommendations based on the empirical results, providing a reference for policy and market participants.

2. Literature Review

This chapter reviews the mechanism of housing price formation, government housing policy, and related empirical studies as the theoretical basis for the model construction and empirical analysis of this study. It comprises four parts: housing price formation theory and the hedonic price model; housing subsidy policy and price capitalization theory; government policy and the housing market; and a synthesis of the literature with the positioning of this study.

2.1 Housing Price Formation Theory and the Hedonic Price Model

Housing prices are formed mainly through the interaction of supply and demand in the housing market, representing the determination of price under market equilibrium (Malpezzi, 1999). However, compared with ordinary goods, housing possesses characteristics such as immovability, heterogeneity, and durability, which make its price formation highly individualized. On the demand side, consumers value the living quality and use value provided by individual dwellings, including floor area, layout, building age, and living amenities, and buyers' preferences over housing conditions are directly reflected in the price formation process. On the supply side, housing is not a single homogeneous good but is composed of dwellings with various characteristics, so that market prices form a price distribution according to different combinations of characteristics. Empirical studies also show that variables such as floor area, building age, housing type, and layout have significant explanatory power for housing prices, and that floor level, building structure, and additional facilities such as parking spaces and elevators significantly affect housing prices.

From a theoretical perspective, Lancaster's (1966) theory of characteristic demand argues that consumer utility derives not from the good itself but from the combination of characteristics it embodies; Rosen (1974) further developed this concept into hedonic pricing theory, holding that market prices can be decomposed into the implicit prices of individual attributes, so that housing prices are essentially the sum of the values of multiple characteristics. In empirical applications, the hedonic price model is usually estimated through regression analysis, taking housing prices as the dependent variable and building characteristics and locational variables as explanatory variables, so as to estimate the marginal effect of each characteristic on price. The model often adopts a semi-log form to address heteroscedasticity and facilitate the interpretation of marginal effects, and incorporates regional dummy variables to control for unobservable differences across regions. Since this study is based on individual transaction data and focuses on the effect of building characteristics on housing prices, the subsequent empirical analysis takes the hedonic price model as its core framework.

2.2 Housing Subsidy Policy and Price Capitalization Theory

Housing demand is highly income-elastic and interest-rate-sensitive. When the financial environment becomes loose due to policy subsidies—such as the interest subsidies, higher loan ceilings, extended grace periods, and longer maturities provided by the New Youth Policy the purchasing demand of potential owner-occupier first-time buyers and marginal buyers is released intensively in the short run. However, demand-side subsidies also carry significant side effects: when short-run supply is inelastic, a rapid rightward shift of the demand curve often fails to substantially increase housing consumption welfare, and instead tends to convert the fiscal resources of the subsidy into capitalization leverage that pushes up housing prices this is the price capitalization mechanism in housing economics.

Price capitalization theory originated in public finance research on local public goods and property tax systems (Oates, 1969), and was later extended to assess the pass-through effect of housing subsidy policies on market prices (Rosen, 1983): when market participants expect to continuously receive some financial benefit in the future, these future income flows are discounted through the market transaction mechanism and reflected at once in the current purchase price of the asset (Mieszkowski & Zodrow, 1989). The degree of policy-induced capitalization essentially depends on the elasticity of housing supply (Hilber & Turner, 2014; Saiz, 2010): in the short run or in markets with inelastic supply, the housing supply curve is nearly vertical (Malpezzi & Mayo, 1987), and when demand expands, buyers use the subsidy as bargaining leverage to bid up prices, so that the subsidy benefit is fully capitalized into housing prices, with the policy dividend effectively transferred to developers or existing owners (Fack, 2006); in the long run or in markets with elastic supply, demand expansion induces developers to increase construction, so that the increased supply absorbs demand and minimizes the capitalization effect (Sinai & Waldfogel, 2005). This capitalization logic continues to receive empirical support in recent work: Chen, Ma, and Orazem (2025) document the capitalization of state tax rate differentials into housing values along state borders, and Crispino and Loberto (2025) show that a large Italian renovation subsidy programme was substantially absorbed into property values rather than translating fully into the intended renovation activity.

The most notable empirical comparison is the United Kingdom's Help to Buy scheme introduced in 2013, whose logic is highly similar to that of the New Youth Policy. Carozzi et al. (2024) exploited two spatial discontinuities in the scheme within a difference-in-discontinuities design and found that, in the severely supply-constrained and already unaffordable Greater London Authority area, the scheme raised house prices by more than the expected present value of the implied interest-rate subsidy while producing no detectable effect on construction volumes; conversely, near the English/Welsh border, where supply constraints are less binding, the scheme increased construction numbers without a noticeable effect on prices. Quasi-experimental evidence from other jurisdictions points in the same direction: using difference-in-differences and regression-kink designs, Berger et al. (2020) show that the United States First-Time Homebuyer Credit raised both home sales and median house prices, and Best & Kleven (2018), using tax notches and stimulus episodes in the United Kingdom, find that transaction-tax relief was substantially capitalized into prices rather than sustained in transaction volumes. Taken together, these studies indicate that the price response to a demand-side subsidy is strongly conditional on the elasticity of local housing supply (Saiz, 2010), and that demand-side subsidy policies therefore exhibit pronounced regional heterogeneity. They also illustrate the identification standard that credible policy evaluation in this literature requires, a point returned to in Section 3.4. The closest institutional analogue to the New Youth Policy, however, is the Portuguese tax exemption for first-home purchasers under the age of 35, introduced in 2024. Clemente-Casinhas & Vale (2026) evaluate that scheme using a synthetic difference-in-differences design with high-frequency municipal asking-price data and Spain as a comparison country, and find that sellers adjusted prices upward in anticipation of and immediately following the announcement, so that a substantial part of the intended benefit was absorbed into higher asking prices within months. Their design also illustrates the point that credible evaluation of a youth-targeted demand-side subsidy requires an explicit comparison group and a test of the parallel-trends assumption, which is the standard we adopt in Section 3.4.

With regard to the mechanism of influence, the mortgage interest rate, loan-to-value ratio, repayment term, and grace period are the four key levers determining a buyer's affordability (Hendershott & Slemrod, 1982): interest subsidies directly lower the real cost of funds for buyers, and a loose funding environment has been empirically shown to significantly push up housing prices (Favilukis et al., 2017; Taylor, 2007); extending the loan term substantially reduces the monthly repayment amount, effectively raising the maximum total price a buyer can afford; and extending the grace period lowers the initial cash-flow pressure of entry, attracting marginal buyers to cross the purchase threshold earlier. When the purchasing power of first-time buyers is amplified in the short run, buyers tend to use up their loan quota to chase properties, driving up overall unit transaction prices (Taltavull de La Paz & White, 2016); the chain move-up effect triggered by first-time buyers transmits the upward price momentum from low-total-price products to the mid-to-high-price market (Acolin et al., 2016); and policy announcement effects further reinforce price increases through the self-fulfillment of market expectations (Case & Shiller, 2003; Himmelberg et al., 2005).

2.3 Government Policy and the Housing Market

The New Youth Policy, launched in August 2023, is an important government program that supports young people's home purchases through housing loan terms; its content is shown in Table 2. Existing studies note that the policy improves young buyers' financing conditions by providing higher loan-to-value ratios and interest subsidies, and that housing market transaction activity increased after its implementation, suggesting that the policy stimulus may strengthen market demand. From the perspective of financial engineering, when the loan term is extended from 30 to 40 years, holding the monthly disposable repayment amount constant, the total loanable amount increases substantially, causing buyers to lose price sensitivity and bid up prices when making offers; and the relaxation of the grace period substantially changes the liquidity constraint on housing assets, greatly lowering the entry threshold for marginal first-time buyers and generating speculative expectations that strongly boost unit transaction prices.

Table 2. Content of the New Youth Housing Loan Subsidy Policy

Item	Description
Implementation period	From August 1, 2023 to July 31, 2026
Loan ceiling	Up to NT$10 million
Interest rate	1.775% after government subsidy
Loan term	Up to 40 years
Grace period	Up to 5 years
Eligibility	(1) The property must be owner-occupied and may not be rented out, used for speculation, or refinanced for arbitrage; (2) the adult applicant, spouse, and minor children must own no other self-use residence.

In addition to macroeconomic conditions, government regulatory measures are also an important institutional factor affecting housing market prices. Since 2010, the central bank has successively implemented selective credit controls, curbing excessive leverage and speculative demand by adjusting loan-to-value ratios, grace periods, and lending standards; subsequent adjustments have gradually strengthened the policy's influence, accompanied by a slowdown in market transaction activity. The regulation of the real estate loan exposure ratio under Article 72-2 of the Banking Act also constrains banks' lending behavior. The data period of this study ends before the implementation of the seventh wave of credit controls in September 2024, in order to avoid confounding the sample estimation with institutional changes and to reduce other sources of policy interference when evaluating the impact of the New Youth Policy on housing prices. Evidence on the Taiwanese transmission mechanism is consistent with this institutional picture: using city-level data on bank mortgage rates, Hsieh & Yang (2024) find that monetary tightening raises mortgage rates and that regional price and credit risks amplify the resulting shock, so that the interest-rate channel operates unevenly across local housing markets. This heterogeneity is one reason why the average price response estimated for a single municipality need not generalize, and it motivates the sub-market analysis reported in Section 4.5.

2.4 Synthesis of the Literature and Research Positioning

The above literature yields the following points. First, housing price formation is highly heterogeneous, and the hedonic price model is the primary tool for analyzing individual housing prices. Second, price capitalization theory indicates that in markets with inelastic supply, demand-side subsidies are highly likely to translate into housing price increases, a conclusion supported by both domestic and international evidence. Third, existing studies on the market impact of the New Youth Policy have provided preliminary findings, but most focus on policy background, institutional content, and market response, and few provide regional, quantitative analysis based on transaction-level data on whether the policy intervention has a statistically significant effect on housing prices. This study therefore takes New Taipei City as an example, uses transaction-level actual price registration data, and controls for building characteristics and locational factors to empirically test the impact of the New Youth Policy on housing prices, so as to fill the gap in the existing literature. This also constitutes the theoretical basis for the expectation that the coefficient of the policy dummy variable on housing prices is significantly positive in the empirical model of this study.

3. Research Methodology

This chapter describes the empirical design of the study. By constructing a multiple regression model, it examines the change in housing prices before and after the implementation of the policy, and incorporates building characteristics and locational variables to control for the effect of housing heterogeneity on prices and improve the accuracy of the policy effect estimate. The chapter comprises four parts: the sample and data source; the research variables; the research method and empirical model specification; and the identification strategy together with the threats to causal inference.

3.1 Sample and Data Source

This study is based on the Ministry of the Interior's real estate transaction actual price registration data, collecting transaction-level records of the New Taipei City housing market as the empirical sample. The data period is set from January 2021 to September 2024, mainly because the central bank successively implemented selective credit control measures during this period; to capture the effect before and after the New Youth Policy and to avoid estimation bias from mixing different policy periods, the end point of the study is set before the seventh wave of credit controls.

In terms of sample selection, following the treatment of existing studies, the following filters were applied: (1) transaction target selection: residential products were taken as the analytical object, excluding commercial, industrial, and other non-residential real estate; (2) building type selection: detached houses, whose large land-share differences make unit prices prone to bias, were removed, retaining only collective dwellings such as apartments, walk-up condominiums, and residential towers; (3) regional scope selection: districts with too few transactions, such as Pingxi, Shiding, Pinglin, Shuangxi, and Wulai, were excluded to ensure sample representativeness and statistical stability; (4) exclusion of special transactions: non-market-price samples such as transactions among relatives, pre-sale, distress sales, debt settlement, and court auctions were removed; (5) data completeness treatment: samples with missing key variables such as total price, building area, building age, or floor level were removed; (6) outlier treatment: abnormal price samples were excluded, and housing prices were log-transformed to reduce distributional skewness. After the above filtering, the original data of 112,733 records were reduced to a final valid sample of 108,671 records, a removal ratio of about 3.6%.

3.2 Research Variables

Following the concept of the hedonic price model, this study treats housing prices as a function of building characteristics, locational conditions, and policy factors. The dependent variable is the natural logarithm of the housing transaction unit price (price per ping), ln(Price), which reduces heteroscedasticity and gives the coefficients an elastic interpretation. The main explanatory variable is the New Youth Policy dummy (Policy), which uses the August 2023 implementation date as the dividing point, set to 0 before implementation and 1 after; since the policy may raise market purchase demand and thereby drive up housing prices by lowering the loan threshold and cost of funds, its effect is expected to be positive.

For the control variables: building age (AGE) reflects depreciation and is expected to be negative; for floor area (AREA), because smaller units have a lower total-price threshold and are favored by first-time buyers and investment demand—the “small unit, high unit price” phenomenon—it is expected to be negatively related to the unit price; floor level (FLOOR) reflects lighting, ventilation, and view conditions and is expected to be positive; building type (TYPE) distinguishes elevator buildings (including residential towers and walk-up condominiums) from apartments, with apartments as the reference; district (DISTRICT) is divided, according to the average transaction unit price of each district during the study period, into high-price areas (Banqiao, Yonghe, Zhonghe, Sanchong, Xindian, Xinzhuang, Tucheng, Luzhou), mid-price areas (Xizhi, Linkou, Taishan, Sanxia, Shulin, Wugu, Shenkeng, Tamsui), and low-price areas (Yingge, Bali, Jinshan, Sanzhi, Ruifang, Wanli, Gongliao, Shimen), with the low-price area as the reference; parking (PARKING) and management organization (MANAGEMENT) both use “none” as the reference and are expected to be positive; and the year variable (YEAR) uses 2021 as the reference to control for differences in the market environment across years and avoid mistaking overall market price changes for the policy effect. The variables are summarized in Table 3.

Table 3. Summary of variables

Variable category	Variable name	Symbol	Expected sign
Dependent variable	Log of housing unit price	ln(Price)	NA
Main explanatory variable	New Youth Policy	Policy	+
Control variables	Building age	AGE	−
	Floor area	AREA	−
	Floor level	FLOOR	+
	Parking space	PARKING	+
	Management organization	MANAGEMENT	+
	Building type	TYPE	NA
	District	DISTRICT	NA
	Year	YEAR	NA

3.3 Research Method and Empirical Model Specification

This study uses ordinary least squares (OLS) for multiple regression analysis. Considering the skewed distribution of housing prices, the dependent variable is log-transformed to improve estimation stability, and categorical variables are converted into dummy variables and included in the model. The main regression model is specified as follows:

$$\mathit{ln(Price)} = \beta_0 + \beta_1\text{Policy} + \beta_2\text{AGE} + \beta_3\text{AREA} + \beta_4\text{FLOOR} + \beta_5\text{PARKING} + \beta_6\text{MANAGEMENT} + \beta_7\text{TYPE} + \beta_8\text{DISTRICT} + \beta_9\text{YEAR} + \varepsilon$$

where ln(Price) is the natural logarithm of the unit price of the i-th housing transaction; Policy is the New Youth Policy dummy, whose coefficient β_1 is the core basis for testing the policy effect—if β_1 is significantly positive, it indicates that the New Youth Policy raises housing prices; AGE, AREA, FLOOR, PARKING, MANAGEMENT, TYPE, DISTRICT, and YEAR are the control variables for building age, floor area, floor level, parking, management organization, building type, district, and year, respectively; and ε is the error term. In addition, to check whether multicollinearity exists among the explanatory variables, this study uses the variance inflation factor (VIF), taking a VIF below 10 as the criterion for the absence of high collinearity.

3.4 Identification Strategy and Threats to Causal Inference

The specification in Section 3.3 identifies the policy effect from a single post-implementation dummy conditioned on building characteristics, district fixed effects, and year fixed effects. Because year dummies absorb common annual shifts in the New Taipei City market, the coefficient on Policy is estimated from the residual within-year difference between transactions completed before and after August 2023. It is a conditional association, not a treatment effect identified from an exogenous assignment mechanism. To move closer to a causal design we implement, in Section 4.5, a difference-in-differences estimator based on differential exposure to the policy, together with an event-study specification and two placebo tests. This section sets out the logic of that design and the assumptions on which it rests.

The New Youth Policy is a nationwide, eligibility-based scheme: every qualifying first-time owner-occupier buyer in Taiwan became eligible on 1 August 2023. There is consequently no geographic discontinuity of the kind exploited by Carozzi et al. (2024), and the actual price registration data do not record whether a buyer drew on the subsidized loan, so treatment status cannot be observed directly. What the scheme does provide, however, is a sharp variation in the intensity of exposure across the product space. The loan ceiling is NT$10 million, and eligibility is confined to owner-occupier first-time buyers. In our pre-policy sample the median transaction price of dwellings of 35 ping or less is approximately NT$11.2 million, so the ceiling is close to binding and the subsidy is economically relevant for this segment; for dwellings larger than 50 ping the median pre-policy transaction price exceeds NT$19.7 million, so the ceiling covers only a small fraction of the purchase and the subsidy is of limited relevance. Following the ex ante exposure logic used by Berger et al. (2020), we therefore define dwellings of 35 ping or less as the high-exposure group and dwellings larger than 50 ping as the low-exposure comparison group, and drop the intermediate band so that the contrast is not blurred at the margin. Because building area is a physical attribute recorded independently of the transaction price, this assignment does not condition on the outcome variable.

A further feature of the data makes validation possible. The policy dummy divides calendar year 2023 internally: 15,727 transactions in the sample occur between January and July 2023 and 12,248 occur between August and December 2023, a split of 43.8% that is consistent with an August cut-off. The observation window can therefore be divided into five periods—2021, 2022, the first seven months of 2023, the last five months of 2023, and 2024—of which the first three precede the policy. This permits a direct test of the parallel-trends assumption and the placement of placebo cut-off dates within the pre-policy period, which is what we report in Section 4.5. Three assumptions must nevertheless hold for the design to support a causal reading, and we assess each of them in Section 4.5. First, parallel trends: absent the policy, the price paths of the two exposure groups would have moved in parallel. Second, no differential contemporaneous shocks: no other intervention or market development affecting small and large dwellings differently coincides with the August 2023 threshold; the seventh wave of selective credit controls falls outside the sample by construction. Third, no spillovers: the low-exposure group must be genuinely untreated. This third assumption is the most vulnerable, because the chain move-up effect discussed in Section 2.2 (Acolin et al., 2016) implies that price pressure originating among first-time buyers may be transmitted to the mid-to-high-price segment. If such spillovers are present, the difference-in-differences estimator recovers only the differential effect and would understate a uniform market-wide policy effect. We return to this point when interpreting the results. In implementing and reporting these tests we follow current practice in the difference-in-differences literature, which emphasizes that the credibility of such a design rests on explicit statement of the comparison group, transparent pre-trend evidence, and inference that accounts for the limited number of effective clusters (Arkhangelsky & Imbens, 2024; Baker, Callaway, Cunningham, Goodman-Bacon, & Sant'Anna, 2026).

4. Empirical Results and Analysis

This chapter analyzes the impact of the New Youth Policy on housing prices in New Taipei City. It first uses descriptive statistics to understand the basic characteristics of the sample; then conducts correlation analysis and a multicollinearity test to ensure the reliability of the model estimates; next applies OLS for regression analysis; then verifies the stability of the results through a robustness check; and finally examines sub-market heterogeneity by housing unit size. The chapter comprises five parts: descriptive statistical analysis; correlation and multicollinearity tests; regression results; the robustness check; and the sub-market heterogeneity analysis.

4.1 Descriptive Statistical Analysis

This study collected 108,671 housing transaction samples in New Taipei City. The dependent variable, the log of the housing unit price, has a mean of 3.61 and a median of 3.65; the price distribution is slightly skewed, and log transformation helps reduce skewness and improve model stability. The policy dummy has a mean of 0.31, indicating that about 31% of the samples fall in the post-implementation period, covering transactions before and after the policy. For the control variables, building age averages 22.59 years with a median of 24, indicating that the sample is dominated by second-hand housing; floor area averages 37.28 ping with a median of 33.09, showing a right-skewed distribution; floor level averages 7.26, reflecting a sample that includes both low-rise apartments and high-rise towers; about 45% of the dwellings have parking spaces, 72% have management organizations, and 72% are elevator buildings; in terms of regional distribution, high-price areas account for 58% and mid-price areas for 37%, consistent with the concentration of New Taipei City housing transactions in densely populated and conveniently connected areas; and the annual sample proportions are evenly distributed, effectively reflecting market changes across years, as shown in Table 4.

Table 4. Descriptive statistics

Variable	Mean	SD	Min	Median	Max
Log of housing unit price	3.61	0.41	0.39	3.65	5.49
New Youth Policy	0.31	0.464	0	0	1
Building age (years)	22.59	14.64	1	24	73
Floor area (ping)	37.28	20.13	0.33	33.09	1172.84
Floor level	7.26	5.67	1	5	43
Has parking	0.45	0.50	0	0	1
Has management	0.72	0.45	0	1	1
Elevator building	0.72	0.45	0	1	1
High-price area	0.58	0.49	0	1	1
Mid-price area	0.37	0.48	0	0	1
2022	0.24	0.43	0	0	1
2023	0.26	0.44	0	0	1
2024	0.20	0.40	0	0	1

Note: N = 108,671; the data period is January 2021 to September 2024.

4.2 Correlation and Multicollinearity Tests

This study first uses correlation analysis to examine the degree of association among the variables; the results show that the direction of correlation between each explanatory variable and the log of the housing unit price is broadly consistent with theoretical expectations, and no abnormally high correlation appears among the explanatory variables. A further multicollinearity test using the VIF, shown in Table 5, indicates that the VIF values of the variables range from 1.317 to 5.601, all below the threshold of 10, showing no high collinearity among the explanatory variables and confirming the reliability of the regression estimates. The slightly higher VIF values for elevator buildings, high-price areas, and management organizations reflect a certain degree of association among building type, management mechanism, and location, but remain within an acceptable range.

Table 5. Multicollinearity test

Variable	VIF	Tolerance
New Youth Policy	3.404	0.294
Building age	3.112	0.321
Floor area	1.562	0.640
Floor level	1.317	0.759
Has parking	2.222	0.450
Has management	4.750	0.211
Elevator building	5.601	0.179
High-price area	5.156	0.194
Mid-price area	5.138	0.195
2022	1.379	0.725
2023	1.979	0.505
2024	3.894	0.257

4.3 Regression Results

The regression results are shown in Table 6. The model's R^2 and adjusted R^2 are both 0.533, indicating that the policy, building characteristic, regional, and annual variables included in this study explain about 53.3% of the variation in the log of the housing unit price; the F value is 10346.121, significant at the 1% level, indicating that the overall model is statistically significant.

For the main explanatory variable, the coefficient of the New Youth Policy is 0.044 with a t value of 12.829, significant at the 1% level, indicating that after controlling for building age, floor area, floor level, parking, management organization, building type, region, and year, the housing unit price in the post-implementation period is still significantly higher than in the pre-implementation period; since the dependent variable is the log of the unit price, the policy coefficient can be interpreted as an increase of about 4.4% in the housing unit price after implementation. This result suggests that the New Youth Policy may, by lowering the burden of home-purchase loans, raising loan ceilings, and improving financing conditions, bring part of the purchase demand into the market earlier, thereby exerting upward pressure on housing prices in New Taipei City—consistent with the expectation of this study and with the inference of price capitalization theory.

The economic magnitude of this coefficient should be assessed against the contemporaneous appreciation of the market rather than in isolation. The year dummies in Table 6 indicate that, relative to 2021, the conditional unit price of housing in New Taipei City was 9.5% higher in 2022, 14.4% higher in 2023, and 21.3% higher in 2024, implying an average conditional appreciation of roughly 7.1% per year over the observation window. Measured against this benchmark, the 4.4% policy differential is equivalent to approximately one-fifth of the cumulative conditional appreciation recorded between 2021 and 2024, or about seven to eight months of typical market appreciation. Measured against the raw index, the Ministry of the Interior housing price index for New Taipei City rose by about 31% between the first quarter of 2021 and the third quarter of 2024, so the policy differential corresponds to roughly one-seventh of the unconditional price growth of the same period. The interpretation is therefore two-sided. The differential is not trivial: for a dwelling transacting at the sample median it is comparable in order of magnitude to several years of the interest saving that the subsidy itself confers, which is the pattern that price capitalization theory predicts when short-run supply is inelastic. At the same time, the differential is clearly a minority share of the observed price increase, so the New Youth Policy cannot be regarded as the dominant driver of housing price growth in New Taipei City over this period; the great majority of the increase is attributable to the broader market trend captured by the year effects and to locational and structural attributes. Readers should also bear in mind that, for the identification reasons set out in Section 3.4, this 4.4% figure represents an upper bound on the policy contribution to the extent that any unobserved time-varying demand shock coincides with the August 2023 cut-off. The interpretation is therefore two-sided. The differential is not trivial: for a dwelling transacting at the sample median it is comparable in order of magnitude to several years of the interest saving that the subsidy itself confers, which is the pattern that price capitalization theory predicts when short-run supply is inelastic. At the same time, it is clearly a minority share of the observed price increase, so the New Youth Policy cannot be regarded as the dominant driver of housing price growth in New Taipei City over this period. Whether even this minority share can be attributed to the policy is a separate question, and it is the one addressed in Section 4.5, where a difference-in-differences design based on differential exposure to the loan ceiling is used to test whether the differential is concentrated in the segment the policy actually reaches.

For the control variables: the coefficient of building age is −0.009 and significant, indicating that each additional year of age reduces the unit price by about 0.9%, reflecting depreciation; the coefficient of floor area is −0.001 and significant, confirming the "small unit, high unit price" phenomenon in the New Taipei City market; the coefficient of floor level is 0.006 and significant, indicating that each additional floor raises the unit price by about 0.6%. The coefficients of parking (−0.080) and management organization (−0.116) differ from the expected direction: the parking variable may be diluted because parking area or ancillary space is included in the per-ping unit price calculation, while management organization is highly correlated with elevator buildings, so its positive effect may have been absorbed by the building type variable; these two results should not be read directly as parking or management reducing housing value, but rather as overlapping explanatory effects and product-structure differences under the per-ping unit price model. The coefficient of elevator buildings is 0.150 and significant, indicating that the unit price of elevator buildings is about 15% to 16% higher than that of apartments. For the regional variables, the coefficients of high-price areas (0.964) and mid-price areas (0.511) are both significant at the 1% level, indicating that location remains one of the most important factors affecting housing prices. The coefficient of the year variable rises from 0.095 in 2022 to 0.213 in 2024, reflecting the overall continued rise of the New Taipei City housing market during the study period.

Table 6. Regression analysis

Variable	Coefficient	t value	p value
New Youth Policy	0.044	12.829	<0.001***
Building age	−0.009	−87.388	<0.001***
Floor area	−0.001	−19.028	<0.001***
Floor level	0.006	32.799	<0.001***
Has parking	−0.080	−30.929	<0.001***
Has management	−0.116	−27.854	<0.001***
Elevator building	0.150	33.041	<0.001***
High-price area	0.964	243.965	<0.001***
Mid-price area	0.511	126.682	<0.001***
2022	0.095	40.181	<0.001***
2023	0.144	52.217	<0.001***
2024	0.213	50.526	<0.001***
Intercept	2.959	477.348	<0.001***
F value	10346.121		<0.001***
R^2	0.533		
Adjusted R^2	0.533		

Note 1: N = 108,671; the dependent variable is the log of the housing unit price. Note 2: *** significant at 1%, ** at 5%, * at 10%.

4.4 Robustness Check

To confirm the reliability of the baseline estimates, this study, based on the standardized residuals of the baseline model, removes observations whose absolute standardized residual exceeds 3, so as to reduce the interference of abnormal transaction prices; after this treatment, the valid sample decreases from 108,671 to 107,518, removing 1,153 extreme observations, about 1.06% of the total. As shown in Table 7, the model after removing extreme values (Model 2) is highly consistent with the baseline model (Model 1) in the direction and statistical significance of the coefficients: the coefficient of the core explanatory variable, the New Youth Policy, is slightly adjusted from 0.044 to 0.043, remaining positive and significant at the 1% level; the coefficients and significance of the building characteristic variables are almost identical; and although the coefficients of high-price and mid-price areas decline slightly, their direction and significance are unchanged. The R^2 of Model 2 rises from 0.533 to 0.557 and its F value to 11273.002, indicating an improvement in the overall explanatory power after removing extreme values. In sum, the significant positive impact of the New Youth Policy on housing prices in New Taipei City does not change with the minor adjustment of the sample composition, and the empirical findings of this study exhibit good robustness.

Table 7. Estimation results of the baseline model and the robustness check model

Variable	Baseline (Model 1)	After removing outliers (Model 2)
New Youth Policy	0.044***	0.043***
Building age	−0.009***	−0.009***
Floor area	−0.001***	−0.001***
Floor level	0.006***	0.006***
Has parking	−0.080***	−0.078***
Has management	−0.116***	−0.115***
Elevator building	0.150***	0.149***
High-price area	0.964***	0.897***
Mid-price area	0.511***	0.451***
2022	0.095***	0.097***
2023	0.144***	0.149***
2024	0.213***	0.217***
Intercept	2.959***	3.018***
F value	10346.121	11273.002
R^2	0.533	0.557
Adjusted R^2	0.533	0.557
N	108,671	107,518

Note 1: The dependent variable is the log of the housing unit price. Note 2: *** significant at 1%, ** at 5%, * at 10%.

4.5 Sub-market Heterogeneity and a Difference-in-Differences Assessment

The New Youth Policy is not directed at the housing market as a whole. Eligibility is restricted to first-time owner-occupier buyers with no other self-use residence, and the loan is capped at NT$10 million, so the scheme reaches dwellings with a low total-price threshold far more than it reaches large residences. If the 4.4% differential estimated in Section 4.3 reflects capitalization of the subsidy rather than a market-wide trend, it should be concentrated in the segment the policy actually touches. This section tests that proposition in two ways: first by re-estimating the baseline model on size-split sub-samples, and then by using the exposure contrast described in Section 3.4 to construct a difference-in-differences estimator with an event study and placebo tests.

4.5.1 Split-sample Estimation by Housing Unit Size

Table 8 reports the baseline specification estimated separately for dwellings below and at or above 25 ping, with heteroskedasticity-robust standard errors. The policy coefficient is 0.040 in the small-unit sub-sample and 0.045 in the large-unit sub-sample. Both are significant at the 1% level, but they are almost identical in magnitude, and the difference between them is −0.005 with a standard error of 0.008 ($z = -0.67$, $p = 0.503$). The difference is therefore statistically indistinguishable from zero, and its sign runs opposite to the direction that price capitalization would predict. The remaining coefficients differ substantially across the two sub-samples — the high-price-area coefficient is 1.241 among small units against 0.845 among large units, and the floor-area coefficient is an order of magnitude larger in absolute value among small units — which indicates that the two segments are governed by materially different price structures. This observation matters for the specification choices reported below.

Table 8. Split-sample estimation by housing unit size

Variable	Small units (< 25 ping)	Large units (≥ 25 ping)
New Youth Policy	0.040*** (0.007)	0.045*** (0.004)
Building age	−0.008*** (0.000)	−0.010*** (0.000)
Floor area	−0.003*** (0.000)	−0.000*** (0.000)
Floor level	0.004*** (0.000)	0.005*** (0.000)
Has parking	−0.015* (0.009)	−0.076*** (0.003)
Has management	−0.094*** (0.011)	−0.114*** (0.005)
Elevator building	0.130*** (0.013)	0.149*** (0.006)
High-price area	1.241*** (0.013)	0.845*** (0.007)
Mid-price area	0.788*** (0.013)	0.397*** (0.007)
2022	0.092*** (0.005)	0.095*** (0.003)
2023	0.145*** (0.006)	0.144*** (0.003)
2024	0.227*** (0.009)	0.209*** (0.005)
Intercept	2.743*** (0.019)	3.033*** (0.009)
R^2	0.591	0.526
Adjusted R^2	0.590	0.526
N	27,708	80,963
Difference in policy coefficient	−0.005 (0.008), z = −0.67, p = 0.503	

Note 1: The dependent variable is the log of the housing unit price. Note 2: Heteroskedasticity-robust (HC1) standard errors in parentheses. Note 3: *** significant at 1%, ** at 5%, * at 10%. Note 4: The difference in the policy coefficient is tested as an unpaired comparison of the two sub-sample estimates.

4.5.2 Difference-in-differences Based on Differential Exposure

Following the design set out in Section 3.4, dwellings of 35 ping or less are treated as the high-exposure group (N = 59,428) and dwellings larger than 50 ping as the low-exposure comparison group (N = 20,341), with the intermediate band excluded. The estimating equation adds a treatment indicator, period fixed effects for the five sub-periods, and the interaction of the treatment indicator with the post-policy indicator, which is the coefficient of interest. Table 9 reports the results.

Specification (1) constrains the coefficients on the control variables to be equal across the two exposure groups and yields an estimate of 0.015, significant at the 1% level. Specification (2) relaxes that constraint by interacting the treatment indicator with the full set of controls. The estimate falls to −0.001 with a p value of 0.825. Given the evidence in Table 8 that the price structures of the two segments differ sharply, specification (2) is the appropriate one, and specification (1) is misspecified: it imposes a common linear floor-area effect across a range running from below 35 ping to above 50 ping, and a common locational premium, neither of which the data support. The null result in specification (2) is not an artefact of the chosen threshold. Rows (5) to (7) re-estimate the same specification with three alternative definitions of the exposure groups, and the estimate remains between −0.001 and 0.000 with p values above 0.83 throughout. Rows (3) and (4) show, in addition, that even the restricted estimate of 0.015 ceases to be statistically significant once standard errors are clustered to allow for serial correlation within group-period cells, which is the concern raised by Bertrand et al. (2004). We note that our setting is the canonical two-group, multiple-period case rather than a staggered-adoption design, so the aggregation problems documented for two-way fixed-effects estimators under heterogeneous timing do not arise here; the binding constraint is instead the credibility of the parallel-trends assumption itself (Baker et al., 2026).

Table 9. Difference-in-differences estimates of the policy effect

Specification	DID estimate	p value
(1) Control coefficients restricted to be equal across groups	0.015*** (0.005)	0.002
(2) Control coefficients allowed to differ across groups	−0.001 (0.005)	0.825
(3) Specification (1), SE clustered by group × period (10)	0.015 (0.015)	0.296
(4) Specification (1), SE clustered by group × period × tier (30)	0.015 (0.017)	0.359
(5) Threshold ≤ 30 / > 55 ping, specification (2)	−0.001 (0.006)	0.889
(6) Threshold ≤ 40 / > 50 ping, specification (2)	−0.001 (0.005)	0.832
(7) Threshold ≤ 35 / > 45 ping, specification (2)	−0.000 (0.005)	0.991

Note 1: The dependent variable is the log of the housing unit price. The reported coefficient is the interaction of the high-exposure indicator with the post-policy indicator. Note 2: All specifications include period fixed effects for the five sub-periods and the full set of control variables. Note 3: HC1 robust standard errors in parentheses unless clustering is stated. Note 4: *** significant at 1%, ** at 5%, * at 10%.

4.5.3 Event Study and Placebo Tests

Table 10 reports an event-study specification in which the treatment indicator is interacted with each period, taking 2021 as the base, together with two placebo tests estimated on pre-policy data alone. Under the unrestricted specification the interaction for the first half of 2023 — a period that precedes the policy — is −0.029 and significant at the 1% level, which is a direct violation of the parallel-trends assumption: the two exposure groups were already diverging before the policy took effect. The interaction for the first post-policy period, the last five months of 2023, is −0.038 and also significant, meaning that the relative price of the high-exposure segment fell rather than rose in the months immediately following implementation. The coefficient for 2024 is 0.007 and not significant. Under the restricted specification the pattern is similarly non-monotonic, with a significantly negative coefficient in the first post-policy period followed by a positive coefficient a full year later. Neither timing pattern is consistent with the immediate capitalization of a subsidy into the prices of the dwellings that the subsidy finances.

The placebo tests point in the same direction. Assigning a false cut-off at the start of 2022 produces a coefficient of 0.002 that is not significant, which is reassuring. Assigning a false cut-off at the start of 2023, however, produces a coefficient of −0.016 that is significant at the 5% level, even though no policy change occurred at that date. Taken with the event-study evidence, this indicates that differential movements between the two exposure groups arise in the pre-policy period as well, and that the identifying assumption required for a causal reading of the difference-in-differences estimate does not hold in these data.

Table 10. Event study and placebo tests

Period (base = 2021)	Restricted specification	Unrestricted specification
2022 (pre-policy)	0.006 (0.006)	−0.009 (0.006)
2023 Jan-Jul (pre-policy)	−0.006 (0.007)	−0.029*** (0.007)
2023 Aug-Dec (post-policy)	−0.017** (0.008)	−0.038*** (0.008)
2024 (post-policy)	0.035*** (0.006)	0.007 (0.006)
Placebo: cut-off set at 2022	0.002 (0.005)	—
Placebo: cut-off set at 2023 Jan	−0.016** (0.007)	—

Note 1: The dependent variable is the log of the housing unit price. Entries are the interaction of the high-exposure indicator with the stated period. Note 2: The restricted specification constrains control coefficients to be equal across groups; the unrestricted specification allows them to differ. Note 3: Placebo tests are estimated on pre-policy data only (2021 to July 2023). Note 4: HC1 robust standard errors in parentheses. *** significant at 1%, ** at 5%, * at 10%.

4.5.4 Interpretation

Three findings emerge, and they point the same way. The split-sample estimation finds no significant difference in the policy coefficient between the segment the policy targets and the segment it does not. The difference-in-differences estimate, once the specification is allowed to accommodate the different price structures of the two segments, is indistinguishable from zero across every threshold we examined. And the identifying assumption that would license a causal reading of that estimate is rejected by both the event study and one of the two placebo tests. We therefore do not find robust evidence that the 4.4% differential reported in Section 4.3 is attributable to price capitalization of the New Youth Policy.

This conclusion requires one important qualification, which follows from the third assumption discussed in Section 3.4. A difference-in-differences estimator recovers a differential effect, not a level effect. If the chain move-up mechanism documented by Acolin et al. (2016) operates in the New Taipei City market, price pressure originating among first-time buyers of small dwellings would be transmitted to the larger dwellings that constitute our comparison group, and a uniform market-wide policy effect would be differenced away. Our results therefore establish that the policy did not raise prices in its target segment relative to the non-target segment; they do not establish that the policy had no effect on the market as a whole. What they do remove is the principal empirical support for the capitalization interpretation as it is usually stated, namely that the subsidy is bid into the prices of exactly those dwellings that eligible buyers compete for. The more parsimonious reading of the evidence is that the post-August-2023 price differential mainly reflects the broader upswing of the New Taipei City market during the observation window, which the year dummies capture only imperfectly.

5. Conclusions and Recommendations

5.1 Conclusions

This study investigates the impact of the New Youth Policy on housing prices in New Taipei City, using transaction-level actual price registration data as the sample, building a multiple regression model through OLS with the log of the housing unit price as the dependent variable, and incorporating building age, floor area, floor level, building type, parking, management organization, district, and year as control variables to reduce the interference of housing heterogeneity.

According to the pooled empirical results, the New Youth Policy variable is significantly and positively associated with housing prices, indicating that after controlling for building characteristics, locational conditions, and year effects, housing prices after implementation remain significantly higher than before. This suggests that the policy may stimulate housing demand reflected in rising prices by lowering the purchase threshold, raising loan ceilings, and easing the financial pressure of home purchase. The further tests reported in Section 4.5, however, do not sustain that reading. Splitting the sample by housing unit size yields policy coefficients of 0.040 and 0.045 for small and large dwellings respectively, a difference that is statistically indistinguishable from zero. A difference-in-differences estimator that exploits the NT$10 million loan ceiling to separate high-exposure from low-exposure dwellings returns an estimate of −0.001 once the control coefficients are allowed to differ between the two segments, and this null is stable across four alternative thresholds. An event study shows a significant divergence between the two groups in the first half of 2023, before the policy took effect, and a placebo cut-off placed at the start of 2023 is significant at the 5% level, so the parallel-trends assumption underpinning the design is not supported. Taken together, these results indicate that housing prices in New Taipei City rose markedly over the observation window, but that the data do not provide robust evidence that the increase is attributable to price capitalization of the New Youth Policy in the segment the policy targets. The 4.4% pooled differential should therefore be read as a conditional association that most probably overstates the policy contribution, and the more parsimonious explanation is that it largely reflects the broader market upswing that the year dummies capture only imperfectly. We note one qualification: because a difference-in-differences estimator recovers a differential rather than a level effect, and because the chain move-up mechanism may transmit price pressure from the target segment to the comparison segment, our results do not rule out a uniform market-wide policy effect. What they remove is the principal empirical support for the capitalization interpretation as it is conventionally stated. This finding is consistent with the view in the literature that the price response to a demand-side subsidy is conditional on local supply elasticity and on the design of the scheme rather than automatic, and it shows that when housing supply is difficult to increase quickly in the short run, a purchase subsidy policy may be partly converted into a housing price increase.

For building characteristics, building age has a significant negative effect on housing prices, reflecting depreciation and market preference; floor level has a significant positive effect, indicating that higher floors are more favored owing to better lighting, ventilation, and views; and floor area has a significant negative relationship, confirming the "small unit, high unit price" phenomenon, as smaller products with lower total-price thresholds are favored by first-time buyers and investment demand. In terms of product characteristics, the prices of elevator buildings are significantly higher than those of apartments, indicating that product type has a clear influence on price; although the parking and management variables are significant, their direction differs from expectation, reflecting the effects of product structure, pricing method, and overlapping explanatory power under the per-ping unit price model, and showing the complexity of the price formation mechanism. For regional factors, the prices of high-price and mid-price areas are both significantly higher than those of low-price areas, with the high-price area effect most pronounced, indicating that location remains a core factor. The year variable results show that the New Taipei City housing market was in an upward price phase during the study period. The robustness check shows that after removing extreme observations, the direction and significance of the policy effect remain unchanged, confirming that the findings are robust.

In sum, the results show that while the New Youth Policy enhances young people's purchasing capacity, it may also drive up market demand and be reflected in rising housing prices; housing prices are jointly affected by building characteristics, product type, and locational conditions, indicating that price formation in the housing market is highly heterogeneous and structural. Beyond supplementing the limited empirical research on the New Youth Policy, this study can serve as a reference for the government in evaluating housing subsidy policies and housing market regulation. This overall assessment should be read subject to the qualification set out above: the evidence establishes a strong contemporaneous association between the policy period and higher prices, but not a robust causal link between the policy and the price increase in its target segment.

5.2 Managerial Implications

For government agencies, the results show that although housing finance subsidies help enhance young people's purchasing capacity, without supply-side complementary measures they may produce price capitalization, so that the policy effect is partly offset by rising prices. In promoting youth home-purchase policies, the government should, in addition to loan preferences, simultaneously strengthen housing supply policies such as increasing social housing supply, promoting urban renewal, or enhancing the elasticity of housing market supply—to reduce the risk that demand stimulus is overly concentrated in price increases. Moreover, since location strongly affects housing prices, public construction, transportation, and improvements in living amenities will continue to influence regional prices, and local governments should be attentive to possible price-spillover effects when promoting regional development and major projects, so as to avoid further widening the housing affordability problem.

For financial institutions, the New Youth Policy may drive up purchase demand and expand the housing loan market; while capturing the market opportunities brought by first-time and young buyers, banks should also be alert to the credit risk that may arise from rising prices and market overheating. Especially when the interest rate environment changes or the business cycle reverses, an excessive price correction could raise credit and default risk, so banks should balance credit quality with risk control in promoting youth mortgage business, so as to maintain financial market stability. For housing market participants, characteristics such as floor area, floor level, management organization, and parking all affect prices, indicating that beyond price, consumers also value housing quality and living conditions; developers and real estate operators may adjust product design according to market preferences for example, improving community management quality, planning parking space, or enhancing housing functionality to raise product competitiveness and market acceptance.

5.3 Recommendations and Research Limitations

In terms of policy recommendations, when promoting housing subsidies and preferential loan policies, the government should consider not only young people's affordability but also the possible impact of the policy on market prices. In housing policy planning, demand-side subsidies can be paired with supply-side policies such as increasing social housing supply, promoting urban renewal, improving housing development efficiency, or increasing the supply of affordable housing—to reduce the price pressure caused by a rapid short-run increase in demand; in regional policy, the housing market characteristics of different areas can be further considered, appropriately strengthening market supervision and credit risk control in high-price, demand-concentrated areas, while improving living attractiveness in relatively low-price areas through transportation and amenity improvements, so as to promote balanced development; and in financial supervision, the authorities should continue to monitor capital flows and price movements to avoid excessive leverage or increased speculative demand, balancing market stability with the goal of housing justice.

In terms of research limitations and future research: this study takes New Taipei City as its scope, and future studies may extend to Taipei City, Taoyuan City, or other metropolitan areas for comparative analysis to understand differences in how housing markets respond to policy; this study takes the housing unit price as the object of analysis, and future work may further incorporate variables such as transaction volume, housing type, or the pre-sale market to more comprehensively examine the policy's impact; in addition, the observation period of this study leans toward the early effects of the policy, and if longer-term data can be obtained in the future, the long-term and persistent effects of the policy can be further analyzed to enhance the completeness and policy value of the research. Beyond these extensions, four limitations of the present research design deserve explicit emphasis. First, the pooled model of Section 3.3 captures a conditional correlation rather than a causal effect, and the difference-in-differences design of Section 4.5, which was implemented in order to move closer to causal identification, does not survive validation: the parallel-trends assumption is rejected by the event study and by a placebo cut-off placed in the pre-policy period. We report these failures rather than suppress them, because they are the substantive result of this exercise; but they mean that no causal estimate of the policy effect can be defended on the basis of these data. Second, treatment status is proxied by dwelling size rather than observed directly. The actual price registration data do not record whether a buyer drew on the subsidized loan, so the high-exposure group necessarily contains ineligible buyers and the low-exposure group may contain eligible ones, which attenuates the estimate towards zero. Third, and in the opposite direction, the comparison group may not be untreated: if the chain move-up effect transmits price pressure from small to large dwellings, a uniform market-wide effect is differenced away, so our null result speaks only to the differential effect. Fourth, the data contain no district identifier, only a three-way price-tier classification, which limits the number of clusters available for inference and the scope for controlling unobserved local heterogeneity; the sample is also restricted to completed transactions of existing collective dwellings, excluding the pre-sale market in which policy-driven expectations are typically expressed earliest. Addressing these limitations is the central priority for future research. The most promising route is to link loan-level administrative records from participating banks to transaction data so that programme take-up is measured directly, which would permit a genuinely untreated comparison group, an intention-to-treat and a treatment-on-the-treated estimate, and placebo cut-off dates that carry real falsification content; inference in such a design should follow Bertrand et al. (2004) in correcting for serial correlation. A second route is to exploit cross-city variation in the intensity of programme take-up, since the price response to a demand-side subsidy is expected to vary with local supply elasticity (Saiz, 2010). A third is to obtain transaction dates at monthly resolution, which would support a regression-discontinuity design around the August 2023 implementation date and would allow announcement effects to be separated from implementation effects.

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Author contributions

Prof. Ching-Hsing Chang and Prof. Chun-Hsien Wang were responsible for the study design and the specification of the empirical model. Yi Lin was responsible for the collection, cleaning, and organization of the actual price registration data and for conducting the regression analysis. Dr. Wen-Tsung Wu drafted the manuscript and, as corresponding author, was responsible for the revision, including the difference-in-differences analysis added at the revision stage. All authors read and approved the final manuscript. The authors declare no special agreements concerning authorship.

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Competing interests

The authors declare that they have no known competing financial interests or personal relationships that could have appeared to influence the work reported in this paper.

Informed consent

This study analyzes anonymized, publicly released administrative transaction records and does not involve human participants.

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Data availability statement

The data that support the findings of this study are openly available from the real estate transaction actual price registration system maintained by the Ministry of the Interior of the Republic of China (Taiwan) at https://plvr.land.moi.gov.tw. The analysis sample was constructed from transactions of collective dwellings in New Taipei City between January 2021 and September 2024. The cleaned analysis file is available from the corresponding author on request.

Data sharing statement

The replication code used to produce all tables in this article, including the difference-in-differences, event-study and placebo specifications reported in Section 4.5, is available from the corresponding author on request.

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References

Acolin, A., Bricker, J., Calem, P., & Wachter, S. (2016). Borrowing constraints and homeownership. *American Economic Review, 106*(5), 625-629. https://doi.org/10.1257/aer.p20161084

Arkhangelsky, D., & Imbens, G. (2024). Causal models for longitudinal and panel data: A survey. *The Econometrics Journal, 27*(3), C1-C61. https://doi.org/10.1093/ectj/utae014

Baker, A., Callaway, B., Cunningham, S., Goodman-Bacon, A., & Sant'Anna, P. H. C. (2026). Difference-in-differences designs: A practitioner's guide. *Journal of Economic Literature, 64*(2), 498-557. https://doi.org/10.1257/jel.20251650

Berger, D., Turner, N., & Zwick, E. (2020). Stimulating housing markets. *The Journal of Finance, 75*(1), 277-321. https://doi.org/10.1111/jofi.12847

Bertrand, M., Duflo, E., & Mullainathan, S. (2004). How much should we trust differences-in-differences estimates? *The Quarterly Journal of Economics, 119*(1), 249-275. https://doi.org/10.1162/003355304772839588

Best, M. C., & Kleven, H. J. (2018). Housing market responses to transaction taxes: Evidence from notches and stimulus in the U.K. *The Review of Economic Studies, 85*(1), 157-193. https://doi.org/10.1093/restud/rdx032

Carozzi, F., Hilber, C. A. L., & Yu, X. (2024). On the economic impacts of mortgage credit expansion policies: Evidence from Help to Buy. *Journal of Urban Economics, 139*, 103611. https://doi.org/10.1016/j.jue.2023.103611

Case, K. E., & Shiller, R. J. (2003). Is there a bubble in the housing market? *Brookings Papers on Economic Activity, 2003*(2), 299-362. https://doi.org/10.1353/eca.2004.0004

Chen, Y., Ma, L., & Orazem, P. F. (2025). Capitalization of state tax rates in housing values at state borders, 2000-2017. *Real Estate Economics, 53*(6), 1178-1199. https://doi.org/10.1111/1540-6229.12533

Clemente-Casinhas, L., & Vale, S. (2026). Exemption or illusion? The impact of a youth tax policy on house asking prices in Portugal. *Real Estate Economics, 54*, 469-503. https://doi.org/10.1111/1540-6229.70020

Crispino, M., & Loberto, M. (2025). Bridging policy and prices: Causal evidence of housing renovation subsidies on property values in Italy. *Regional Science and Urban Economics, 112*, 104095. https://doi.org/10.1016/j.regsciurbeco.2025.104095

Fack, G. (2006). Are housing benefit an effective way to redistribute income? Evidence from a natural experiment in France. *Labour Economics, 13*(6), 747-771. https://doi.org/10.1016/j.labeco.2006.01.001

Favilukis, J., Ludvigson, S. C., & Van Nieuwerburgh, S. (2017). The macroeconomic effects of housing wealth, housing finance, and limited risk sharing in general equilibrium. *Journal of Political Economy, 125*(1), 140-223. https://doi.org/10.1086/689606

Hendershott, P. H., & Slemrod, J. (1982). Taxes and the user cost of capital for owner-occupied housing. *Real Estate Economics, 10*(4), 375-393. https://doi.org/10.1111/1540-6229.00270

Hilber, C. A. L., & Turner, T. M. (2014). The mortgage interest deduction and its impact on homeownership decisions. *The Review of Economics and Statistics, 96*(4), 618-637. https://doi.org/10.1162/REST_a_00427

Himmelberg, C., Mayer, C., & Sinai, T. (2005). Assessing high house prices: Bubbles, fundamentals and misperceptions. *Journal of Economic Perspectives, 19*(4), 67-92. https://doi.org/10.1257/089533005775196769

Hsieh, Y.-S., & Yang, C.-W. (2024). The dominant risks in the interest rate channel: Evidence from the urban housing market in Taiwan. *Applied Economics, 56*(58), 8091-8111. https://doi.org/10.1080/00036846.2023.2289927

Lancaster, K. J. (1966). A new approach to consumer theory. *Journal of Political Economy, 74*(2), 132-157. https://doi.org/10.1086/259131

Malpezzi, S. (1999). A simple error correction model of house prices. *Journal of Housing*

Economics, 8(1), 27-62. https://doi.org/10.1006/jhec.1999.0240

Malpezzi, S., & Mayo, S. K. (1987). The demand for housing in developing countries: Empirical estimates from household data. *Economic Development and Cultural Change, 35*(4), 687-721. https://doi.org/10.1086/451618

Mieszkowski, P., & Zodrow, G. R. (1989). Taxation and the Tiebout model: The differential effects of head taxes, taxes on land rents, and property taxes. *Journal of Economic Literature, 27*(3), 1098-1146.

Oates, W. E. (1969). The effects of property taxes and local public spending on property values: An empirical study of tax capitalization and the Tiebout hypothesis. *Journal of Political Economy, 77*(6), 957-971. https://doi.org/10.1086/259584

Rosen, H. S. (1983). *Housing subsidies: Effects on housing decisions, efficiency, and equity* (NBER Working Paper No. 1161). National Bureau of Economic Research. https://doi.org/10.3386/w1161

Rosen, S. (1974). Hedonic prices and implicit markets: Product differentiation in pure competition. *Journal of Political Economy, 82*(1), 34-55. https://doi.org/10.1086/260169

Saiz, A. (2010). The geographic determinants of housing supply. *The Quarterly Journal of Economics, 125*(3), 1253-1296. https://doi.org/10.1162/qjec.2010.125.3.1253

Sinai, T., & Waldfogel, J. (2005). Do low-income housing subsidies increase the occupied housing stock? *Journal of Public Economics, 89*(11-12), 2137-2164. https://doi.org/10.1016/j.jpubeco.2004.06.015

Taltavull de La Paz, P., & White, M. (2016). The sources of house price change: Identifying liquidity shocks to the housing market. *Journal of European Real Estate Research, 9*(1), 98-120. https://doi.org/10.1108/JERER-11-2015-0041

Taylor, J. B. (2007). *Housing and monetary policy* (NBER Working Paper No. 13682). National Bureau of Economic Research. https://doi.org/10.3386/w13682