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Demographics Outlook, Credit Conditions, and Property Prices

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Abstract

Many developed countries have experienced prolonged economic stagnation in the aftermath of property bubbles bursting. Such observations have led people to believe that economic stagnation accompanied by property bubbles has longer and more severe consequences than other forms of economic stagnation. This study conducts an empirical analysis to challenge this hypothesis and suggest that demographics are closely related to other aspects of long-term economic stagnation. Using panel data from 17 countries from 1974 to 2018, we investigate the residential property price dynamics by incorporating demographic factors and considering the interaction of those demographics with credit conditions. Our results shed new light on the importance of demographic factors in modeling the long-run equilibrium of residential property prices. We find that the effect of nominal interest rates determined by monetary policy on asset prices varies depending on the country and the degree of population aging at the time. We also find that persistently optimistic population projections lead to the over-supply of residential stocks in rapidly aging countries, resulting in stagnant residential property markets. We demonstrated that ignoring the demographic and credit factors in the dynamics may lead to misjudgment of the long-run equilibrium conditions and incorrect policy decisions.

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Demographics Outlook, Credit Conditions, and Property Prices*

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Abstract

Many developed countries have experienced prolonged economic stagnation in the aftermath of property bubbles bursting. Such observations have led people to believe that economic stagnation accompanied by property bubbles has longer and more severe consequences than other forms of economic stagnation. This study conducts an empirical analysis to challenge this hypothesis and suggest that demographics are closely related to other aspects of long-term economic stagnation. Using panel data from 17 countries from 1974 to 2018, we investigate the residential property price dynamics by incorporating demographic factors and considering the interaction of those demographics with credit conditions. Our results shed new light on the importance of demographic factors in modeling the long-run equilibrium of residential property prices. We find that the effect of nominal interest rates determined by monetary policy on asset prices varies depending on the country and the degree of population aging at the time. We also find that persistently optimistic population projections lead to the over-supply of residential stocks in rapidly aging countries, resulting in stagnant residential property markets. We demonstrated that ignoring the demographic and credit factors in the dynamics may lead to misjudgment of the long-run equilibrium conditions and incorrect policy decisions.

Keywords: asset pricing, population aging, residential property prices, nominal interest rates, credit cycle, panel cointegration analysis

Classification Codes: E31, R21, R31

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1 Property Bubbles and Population Dynamics

During the rapid economic growth in the post-World War II era, Japan experienced the largest property bubble of the 20th century (Shimizu and Watanabe, 2010[38]). At the peak of the Japanese bubble, land in central Tokyo was selling for as much as 50 million yen (approximately USD 400,000) per square meter, and the total value of land in Japan was estimated to be four times that of the United States. However, the Japanese economy suffered a prolonged period of economic stagnation following the property bubble bursting. Throughout our history, we have witnessed many episodes of bubble economies followed by recessions.*1

It has become evident that the economic growth model associated with a bubble economy does not return to its pre-bubble growth trajectory after the bubble's collapse.*2 Furthermore, the economic stagnation accompanying a bubble economy and its subsequent collapse tend to have more significant and enduring effects than other forms of economic stagnation. Hirano and Yanagawa (2017)[16] explain the long-term economic stagnation following the bubble's bursting by focusing on the distortions in resource allocation caused by the property bubble.*3

Did Japan's long-term economic stagnation really stem from the bubble economy?*4 Over the past decade, property prices have consistently risen in major Chinese cities such as Beijing, Shanghai, Hong Kong, and Shenzhen as well as in major Asian cities such as Singapore and Seoul. The same trend is observed in global cities such as London, Paris, San Francisco, Sydney, and Vancouver. If we misjudge these conditions and fail to anticipate a decline, it could lead to incorrect policy decisions, especially in the realm of monetary policy. However, existing research has not provided sufficient answers to such questions.

The primary purpose of this study is to argue that demographics exert a significant influence on these issues. Assuming that a property bubble impacts a recession, this study reveals that demographic factors are strongly related to long-term fluctuations in property price dynamics. A property bubble can be simplistically defined as a prolonged deviation from the fundamental level. Such a situation is unsustainable and eventually reverts to the fundamental level. Conversely, a situation may persist below the fundamental level, as seen in the "lost decade" following the bursting of Japan's property bubble in the 1980s. Therefore, in this study, we begin by incorporating demographic factors into the fundamental model of property price dynamics.

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*1 Kindleberger (2000)[21] introduces 38 episodes in the 1618–1998 period. Jorda et al. (2015)[19] analyze bubbles in housing markets in 17 countries. Hirano and Toda (2023a[14], 2023b[15]) extensively summarize the episodic and theoretical issues of the bubble economy.

*2 From a global perspective, Japan and Sweden in the late 20th century and various Western countries, including the United States, in the early 21st century (after the global financial crisis of 2008) faced the formation of property bubbles and then long-term economic stagnation following the collapse of those bubbles. According to Claessens et al. (2011)[4], not all property bubbles lead to financial crises and not all financial crises are caused by property bubbles. However, in many countries, economic slumps together with property market failures triggered by the formation and collapse of property bubbles are significant in terms of length and scope (Crowe et al., 2013[5]).

*3 During the bubble generation period, property is transferred to less productive investment entities, as it is traded at prices that are not commensurate with productivity. Furthermore, after the bursting of the property bubble, highly productive firms are more likely to borrow heavily in the bubble period and thus give up their property, which is then purchased by less productive firms. This mechanism, by which resource allocation is distorted due to a divergence from productivity, leading to long-term land price declines and economic stagnation, is consistent with the findings of Kiyotaki and Moor (1997)[22].

*4 Diewert et al. (2023)[6] find that excessively high land prices crowd out capital and labor, thereby lowering productivity.

The world economy is facing the unprecedented challenge of changing population dynamics, with some countries experiencing a rapid rise in the young population while others face rapid aging. If we examine Japan's demographics in 1990, the year the bubble burst, it was the same year as when Japan's working population reversed its declining trend. In the 21st century, Japan's population was in a period of both fast aging (the fastest rate in the world) and absolute decline.*5 In the coming decades, China, South Korea, and other major Asian countries will experience rapid population aging, accompanied by a phase of population decline. Global aging will progress swiftly, a challenge shared by Europe and various other countries.

When considered within the framework of a fundamental model of property prices, it is reasonable to assume that a declining and aging population will affect housing income or rent, which is the numerator. This is because it will reduce demand for housing, as demonstrated by Mankiw and Weil (1989)[25]. However, in the fundamental model, changes in the denominator—the discount rate—have a significantly larger impact than changes in the numerator, which represents income or rent. Further, the nominal interest rate plays an important role in the discount rate (determined by monetary policy). In such uncharted territory, it is crucial to ask: Is the effect of the change in nominal interest rates similar across countries or vastly different? Why is the recovery from the collapse of so-called property bubbles inconsistent across countries?

We attempt to answer these questions using an econometric approach based on the experiences of 17 divergent economies in demographic composition and stages of economic development over 45 years, focusing on population dynamics statistics. In the context of these phenomena, we explore what kinds of economic mechanisms—including demographics—are at work, focusing on their property markets.*6

Our empirical analysis presents a model incorporating demographics to explain the long-term decline in property prices and economic stagnation. Even before the global financial crisis triggered by the collapse of Lehman Brothers, numerous studies had attempted to properly understand the correlation between long-term economic stagnation and the mechanisms underlying large fluctuations in asset prices, such as property bubbles. Reinhart and Rogoff (2009)[36] examine the relationship between the two on the basis of long-term economic data from multiple countries covering over 100 years.*7 There are several theoretical frameworks to

*5 In light of this, the number of vacant houses has continued to increase to depress regional property markets. Within 10 years, it has been predicted that one-quarter of all residential houses will be vacant. Furthermore, ownership of more than 10% of the nation's land has been relinquished by the owners. The population decline in Japan's regional cities and aging trend had already begun in the second half of the 20th century. Some municipalities became financially insolvent in the early 21st century, leading to the coining of the phrase "extinct municipalities." In 2010, the town of Yubari in the northern Japanese region of Hokkaido suffered financial insolvency. Looking at population composition reveals that this insolvency occurred when the old age dependency ratio, which indicates the proportion of the old population (aged 65 and over) to the working-age population (aged 15 to 65), was over 90% (see Nakagawa and Shimizu, 2023[27]).

*6 In recent years, residential prices in many "superstar" cities such as New York, Boston, Washington, D.C., San Francisco, and Seattle in the United States; London, Frankfurt, and Amsterdam in Europe; and Tokyo, Seoul, Beijing, Singapore, Sydney, and Melbourne in the Asia-Pacific region have increased dramatically. However, substantial regional heterogeneity exists, making price behavior different between these superstar cities and the national average. This study examines average residential property prices nationally as opposed to those of superstar cities (see Gyourko et al., 2013[11]; Badarinza et al., 2021[1]).

*7 Reinhart and Rogoff (2009)[36] elucidate four common phenomena observed in countries that have suffered financial crises: 1) among asset prices, property prices in particular diverged significantly from earnings; 2) debts increased far beyond income/net assets and leveraging increased; 3) substantial capital inflows continued; and 4) productivity increases lagged behind increases in asset values and debt. Further, the study clarifies that when society as a whole is excessively optimistic, it leads to high growth explain Reinhart and Rogoff's (2009)[36] findings. For example, Kiyotaki and Moore (1997)[22] provide a micro-foundation theory of leveraging and de-leveraging during credit cycles. When leverage is high for the economy, even a small adverse shock can deteriorate economic conditions sharply to induce accelerated de-leveraging.*8 During the de-leveraging process, many assets, including property, are on "fire-sale," causing lasting damage to property markets.*9

By comparison, let us consider the Japanese experience. The post-war era of rapid economic growth was driven by the generation born during the baby boom reaching working age. This period is known as the "population bonus phase" (Ito and Hoshi, 2020[18]). Then, in the early 1980s, this baby-boomer generation became homebuyers and entered the housing market, generating the highest level of housing demand since the war and triggering the formation of a property bubble. At that time in Japan, optimism was extremely high. After the bubble collapse in 1990, Japan's working-age population continued to decline. In recent years, in conjunction with deflation and a low economic growth rate, the property market has struggled with a high vacant house rate and an increase of land with relinquished ownership: It has entered the "population onus phase." With the appearance of these problems, a pessimistic mood has spread across society (Tamai et al., 2017[40]). It follows logically that close relationships exist among population factors, large fluctuations in the property market, and economic downturns.*10

The literature review suggests that demographics and the property market strongly influence macroeconomic fluctuations such as economic growth and length of recessions. We attempt to decipher this mechanism by focusing on the relationship between the residential property market and demographics. We investigate the following two hypotheses using panel data from 17 countries spread over 45 years.

(1) Did changes in population composition influence the dynamics of residential property prices?
(2) Did changes in population composition amplify/dampen the effects of nominal interest rates on residential property prices?

This study's key contributions are highlighted below. First, there is no consensus among theories that simultaneously explain demographic changes, property price dynamics, and credit cycles; indeed, this theoretical strand is still being developed and is not ready for testing using data. Therefore, we base our study on the most basic theoretical relationships in the fundamental or present value relationship (PVR) model (Campbell and Shiller, 1988[2]; Hirano and Toda, 2023a[14], 2023b[15]; Walras, 1954[42]). Using international panel data from 17

via financial leverage, which fosters growth in a self-feeding manner. Conversely, the authors also note that once optimism turns into pessimism, regardless of the reason, the economy enters a cycle of contraction.

*8 In particular, the rapid expansion of credit corresponds to the financial instability theory of Minsky (1992)[26]. The financial crises from the late 20th century to the early 21st century were likely to happen given the presence of the strong cumulative interaction of factors such as marked shifts in population composition, property bubbles, and credit cycles.

*9 Nishimura (2016)[29] analyzes the systems, policies, and histories of the United States, various European countries, and Japan, noting the presence of two common factors in countries facing economic crises: (1) excessive optimism caused by favorable changes in population composition (rapid increase in the young population) and (2) rapid expansion of credit due to the spread of so-called new finance technologies and vehicles introduced during the excessively optimistic time. For example, large time deposits with no interest rate ceiling were introduced and Commercial Papers were allowed for large corporations around 1986, at the beginning of the bubble economy in Japan.

*10 Furthermore, models of secular stagnation support this issue. Secular stagnation is the theory that economic stagnation continues in the long term due to stricter borrowing constraints, aging and declining populations, and widening income inequality (Eggertsson and Mehrotra, Robbins, 2019[8]).

countries with diverse population compositions, population trends, economic growth rates, and housing market environments, over almost half a century, this study empirically examines the relationships among demographics, property price dynamics, and credit cycles. In previous research, limited residential property price data were obtained for only a limited period. It is, therefore, only possible to analyze at most one property boom and bust cycle. However, this study's dataset includes various cases, including countries with an increasing young population, countries that have already reached a high aging rate, and countries that have experienced two or more property boom and bust cycles in the period under study. Thus, it enables us to consider various cases necessary to analyze slow-moving long-term factors.*11

Second, to the best of our knowledge, this study is the first to analyze the effects of expectation errors in demographics. Mankiw and Weil's (1989)[25] central criticism of demographics and housing market-related analysis focuses on the fact that if economic agents' expectations are rational (i.e., with no persistent expectation errors) with respect to demographic projections, there should be little impact of demographic changes on residential property prices, since the supply will be adjusted accordingly when it is sufficiently elastic. However, we find evidence that demographic expectations are not rational and, for example, expectation errors about populations persist (see Figure 1). Thus, when young populations are growing and the underestimation of demand persists, housing supply shortages accumulate over time to increase residential property prices. By contrast, when the population is aging rapidly, and the overestimation of demand persists, a housing supply surplus becomes persistent, and residential property prices are depressed. Therefore, to assess the effects of possible persistent demographic expectation errors, we collect data by tracing the population projection data published by each country throughout the analysis period, as far back as possible, and estimate the difference between the actual figure and previous projections.

Third, after assessing the long-term relationship between residential property prices and demographic factors, this study examines the interaction effect of demographic factors and nominal interest rates. In particular, we examine whether the impact of declining nominal interest rates on residential prices is substantially smaller in an aged economy such as present-day Japan than in an economy with a growing young population such as Japan 30 years ago.

2 Empirical Model

In their seminal work, Mankiw and Weil (1989)[25] examine the relationship between demographics and the property market. They argue that in the 1980s, housing demand peaked because of the baby-boomer generation in the United States; subsequently, over the next 20 years, until 2007, real residential property prices decreased by 47% because of population decline. This so-called "asset meltdown hypothesis" subsequently caused considerable debate (Mankiw and Weil, 1989[25]).

Two major issues were central to this debate: supply elasticity and the accuracy of projected demographic changes. Researchers argued that demographic changes take place at an extremely slow pace, and thus, their predictions are often accurate. Therefore, if the housing supply is elastic, even in the event of a pessimistic future population projection, no residential

*11 After the financial crisis, the International Monetary Fund (IMF) and Bank for International Settlements (BIS) took the lead in developing the property price index internationally. The BIS and Organisation for Economic Co-operation and Development (OECD) began publishing the property price index in 2016 (see Diewert et al., 2020[7]).

property price slump should occur since supply will be adjusted via stock adjustment (Hamilton, 1991[12]; Hendershott, 1991[13]). Engelhardt and Poterba's (1991)[9] empirical research reports no statistically significant relationship between demographic changes and residential property price changes in Canada.

Studies focusing on Japan, such as Ohtake and Shintani (1996)[31], obtain similar results. They conclude that while demographic changes impact residential property price changes in the short term—when supply constraints exist—they do not impact residential property prices in the long run, since the housing supply is adjusted accordingly. Further, Shimizu and Watanabe (2010)[38] estimate housing demand using Mankiw and Weil's (1989)[25] framework and expand the model on the basis of panel data (by prefecture in Japan and by state in the United States); they show that housing demand fluctuation shocks do not impact residential property prices in the long run.

However, Japan's experiences in the most recent decade have cast doubts about the elasticity of the supply of houses in an increasingly rapidly aging society. Vacant houses and land with unknown ownership are increasing at an unprecedented rate in Japan, raising doubts about how elastic the housing supply is in a low-growth economy. In particular, land, an essential factor determining housing stocks, is a real, non-depreciable asset. The supply volume is rather limited (physically or by zoning) so that it is inelastic rather than elastic.[*12] If this is the case, even when we assume that demographic changes are perfectly predictable, demographics may influence residential asset prices. In this context, Nishimura and Takáts (2012)[30] and Takáts (2012)[39], using a two-generation overlapping generation model, demonstrate that increases in the working-age population lead to rises in real residential property prices in the ultra-long run (between generations). However, it cannot explain residential property price changes in the medium term (around 10 years) or over the business cycle period (around two years), which is the subject of our study.[*13]

Next, we consider the accuracy of projections of demographic changes. Figure 1 shows population statistics experts' predictions of the total fertility rate (TFR) starting in 1975, when the fertility rate dropped noticeably, to 2012, when the drop was reversed. This figure has several notable features of persistent projection errors at the unfavorable time of declining TFRs: 1) the recent unforeseen changes are considered temporary; 2) the level will eventually revert to a presupposed long-term level that is close to the "old normal" of the previous period; and 3), most importantly, even if the current period figure is constantly lower than the long-term level considered in the previous projection, the downward revision of the projection is very slow.

The last point is crucial since government experts' demographic projections, which are indispensable for private-sector actors to anticipate long-term total demand and total supply in an economy, may not be "rational" in the sense that the expectation (projection) errors are not zero on average over a given period, possibly due to the slow adjustment. The figure shows that overly optimistic projections persist over the period of analysis, and optimistic errors accumulate when the situation is unfavorable in terms of population dynamics. For instance, a projected TFR for 2010, made in 1992, was approximately 1.8. However, the actual TFR in

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[*12] Gyourko et al. (2013)[11] focus on the point that the most pronounced residential price movement is driven by the limited supply of land. Similarly, Knoll et al. (2017)[23] show that residential price movement over a century is mostly brought about by the movement in land prices and that the sharp increase in the second half of the 20th century was caused by the substantial appreciation of land.

[*13] Their model examines an intergenerational portfolio selection problem. Thus, while it can explain that population composition changes impact intergenerational or ultra-long-run price changes, it is unable to explain changes in the medium run and over business cycles.

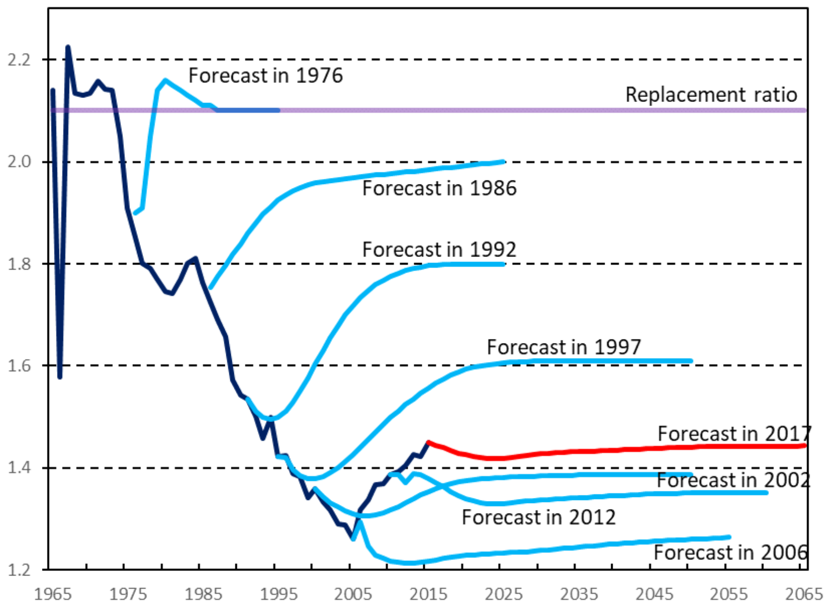

Figure 1 Persistently Optimistic Predictions of the Total Fertility Rate by Government Experts at the Unfavorable Time of Declining Rates (Japan)
Source: National Institute of Population and Social Security Research. Various issues.

2010 was 1.4. This gap (0.4) is likely to lead to over-capacity and over-supply. This suggests another important route through which demographic factors influence the economy: when demographic dynamics are unfavorable, it is likely to be translated into over-capacity and over-supply through persistent and accumulated expectation errors. Thus, if government experts' and regular people's overly optimistic forecast errors persist and accumulate, residential property prices decline more than in the contrasting scenario.

This study empirically investigates the effects of demographic dynamics on property prices considering the above-mentioned two points. Specifically, we elucidate the PVR's perfect foresight to determine long-term nominal residential property prices.[*14] We start with the following long-run equilibrium relations among the nominal residential property price index (P^{rppi}), consumer price index (CPI) (P^{cpi}), real rent (R) , nominal interest rate (i), and nominal expected rate of change in housing rents (g^e):[*15*16]

$$P^{rppi} = \frac{P^{cpi} \times R}{i - g^e}. \tag{1}$$

As Equation (1) is for the no-arbitrage condition in the competitive equilibrium, it may be

[*14] We perform a robustness check based on the PVR using real residential property prices, real rent, and real interest, but there are no major changes in the baseline results. The robustness check is performed using both perfect foresight and static expectation for the real interest rate. For the details, see Appendix 1.

[*15] Walras (1954)[42] is the first to model the PVR of asset prices. For further model details, see Campbell and Shiller (1988)[2].

[*16] In addition to the nominal interest rate, there is a nominal risk premium in the denominator since residential property is a risky asset. Thus, even if the nominal interest rate becomes negative, the denominator is still positive and the PVR is well defined. For simplicity, we assume that the nominal risk premium does not change over time and thus it becomes a constant in Equation (2).

considered common to all capital markets. Accordingly, we assume that this basic relationship is common to all the countries analyzed below; however, it will be tested statistically in the following analysis.

The lack of reliable housing rent data is a common problem for empirical analyses based on Equation (1). In this study, for a consistent definition across the 17 countries, we use working-age per-capita real gross domestic product (GDP) as a proxy variable for real rent R. This treatment can be justified as follows. If we consider GDP from the consumption side, the share of the household consumption basket paid toward housing services is roughly constant.[*17] Thus, if we consider that households mainly comprise working-age people, using working-age per-capita real GDP as a proxy variable for real rent R may be considered a reasonable assumption.[*18]

Meanwhile, as mentioned previously, Takáts (2012)[39] and Nishimura and Takáts (2012)[30] derive the relationship between working-age population increases and rises in real residential property prices using a two-generation intergenerational model. Further, empirical studies such as Saita et al. (2016)[37] and Takáts (2012)[39] indicate that population composition has a significant impact on residential property prices. From previous research findings, we specify the expected rate of change of housing rents as a function of population factors, namely, population composition and total population. As we have clarified, population variables may not be accurately estimated, and their errors may persist and accumulate over time (i.e., they may not be "rational"). To incorporate these possibilities into the perfect foresight framework of Equation (1), we assume the perfect foresight of the expectation errors of the demographic variables and do not impose zero-sum constraints; this assumption should be tested empirically.

Equation (2) is a benchmark specification derived by performing a logarithmic transformation on both sides of the long-term PVR established as above:[*19*20]

$$\begin{aligned}\log P_{jt}^{rppi} &= \beta_{0j} + \beta_1 \log P_{jt}^{cpi} + \beta_2 \log\left(\frac{Y_{jt}}{pop_{jt}^{wrk}}\right) + \beta_3 i_{jt} \\ &\quad + \beta_4 \log pop_{jt}^{total} + \beta_5 (n_{jt}^{yng} - n_{jt}^{wrk}) + \beta_6 (n_{jt}^{old} - n_{jt}^{wrk}) + \varepsilon_{jt}, \qquad (2)\end{aligned}$$

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*17 Housing rent data do not exist for all 17 countries. Therefore, per-capita GDP or income is used as a proxy variable in the empirical analysis. This assumption is extremely strong. However, housing services account for the largest proportion of household consumption, between 25% and 30%. In the Japanese case, where housing rent data exist, the correlation coefficient between housing rent and the rate of change in per-capita GDP is 0.97. Hirano and Toda (2023b)[15] compare land prices and per-capita income growth as identifiers of a property bubble. In other words, per-capita income and GDP are often used as proxy variables for housing rent. Therefore, we adopt per-capita GDP as a suitable proxy variable for housing rent.

*18 Chapter 6 ("Measuring the Services of Durables and Owner Occupied Housing") in Diewert et al. (2020)[7] presents the situation in various countries, along with a housing services estimation method.

*19 As the relationship between per-capita real GDP and per-capita real rent among the working-age population is not necessarily linear, β_1 is not necessarily 1. In actual empirical research, it is not 1.

*20 To understand the impact of population dynamics on fundamentals, we employ the following specification, based on Saita et al. (2016)[37], Takáts (2012)[39], and Tamai et al. (2017)[40]:

$$\text{Population factors}_{jt} = \delta_0 \log(pop_{jt}^{total}) + (\delta_1 n_{jt}^{yng} + \delta_2 n_{jt}^{wrk} + \delta_3 n_{jt}^{old}).$$

Population factors are categorized by generation (young generation, working generation, old generation). Each category's definition is explained below. Since the total of the population ratios by generation at time point t is always 1, δ_1, δ_2, and δ_3 cannot be estimated simultaneously. This is handled by imposing the restriction $\delta_1 + \delta_2 + \delta_3 = 0$ (Fair and Dominguez, 1991[10]). That is, the formula is modified as follows:

$$\text{Population factors}_{jt} = \delta_0 \log(pop_{jt}^{total}) + \delta_1 (n_{jt}^{yng} - n_{jt}^{wrk}) + \delta_3 (n_{jt}^{old} - n_{jt}^{wrk}).$$

where the subscript j represents the country and t represents the time point. For the coefficients of the variables other than the demographic variables, interpretation and sign conditions can be verified based on several economic theories. For instance, the absence of monetary illusion suggests $\beta_1 = 1$, the suggestion that an increase in housing rents raises residential property prices indicates $\beta_2 > 0$, and the suggestion that an increase in the nominal interest rate lowers residential property prices implies $\beta_3 < 0$. In Nishimura and Takáts' (2012)[30] and Takáts' (2012)[39] two-generation model, the main buyers of property are the younger generation; therefore, it is predicted that increases in the younger generation's population result in higher real residential property prices ($\beta_5 > 0$). Conversely, it is expected that increases in the older generation's population have a deflating effect on residential property prices ($\beta_6 < 0$).

In the long-run relation Equation (2), the constant term differs by country, but we presume that the other variables' coefficients are homogeneous, with no differences between countries, based on the assumption that the no-arbitrage condition in the competitive equilibrium is common to all capital markets. This assumption is also tested as a part of the empirical research.

As an extension of the basic model, we perform two types of analyses. First, for the property bubble and collapse periods, we empirically analyze how population composition affects the impact of interest rates on property prices. Specifically, we add the population factor and interest rate interaction terms into the estimation model.

Nishimura (2016)[29] suggests the possibility that optimism induced by population bonuses and expanded credit conditions typical of low interest rates have a fairly synergistic effect on property demand and property bubbles. A decline in nominal interest rates has large positive impacts on property prices when the population is young and growing. By contrast, the experiences of Japan, the United States, and Ireland, following property booms, show that the effect of monetary easing measures such as lowering nominal interest rates is severely restricted in countries facing population onus (aging) periods.

To test whether the phenomenon observed in these three countries is simply a coincidence, we estimate a model that adds the interaction term of the nominal interest rate and population ratio into the model above. The following equation is the expanded long-run equilibrium relation:

$$\begin{aligned}\log P_{jt}^{rppi} = {} & \beta_{0j} + \beta_1 \log P_{jt}^{cpi} + \beta_2 \log\left(\frac{Y_{jt}}{pop_{jt}^{wrk}}\right) + \beta_3 i_{jt} + \beta_4 \log pop_{jt}^{total} \\ & + \beta_5 (n_{jt}^{yng} - n_{jt}^{wrk}) + \beta_6 (n_{jt}^{old} - n_{jt}^{wrk}) \\ & + \beta_7 i_{jt} \times (n_{jt}^{yng} - n_{jt}^{wrk}) + \beta_8 i_{jt} \times (n_{jt}^{old} - n_{jt}^{wrk}) + \varepsilon_{jt}. \end{aligned} \tag{3}$$

We examine Equation (3)'s estimated coefficient of i incorporating the interaction term to check whether the population bonus period's coefficient is significantly different from that of the onus period.

Our second analysis investigates the effect of persistent demographic expectation errors on residential property prices in the long-run equilibrium. Property is a durable good, and thus supply cannot be easily adjusted instantaneously to sudden fluctuations in demand. The adjustment may be possible but with a substantial cost. Therefore, we may assume that the supply side supplies housing by predicting demand for a certain period in advance.

In this section, we explore the type of impact on residential property prices if the population prediction at a given time turns out to be wrong in a future period. We examine how expectation errors change the results of the benchmark case in Equation (2) and the extension case in Equation (3). To do this, we decompose the young population ratio n_{jt}^{yng} used in the regression analysis in terms of the ratio $\breve{n}_{jt}^{yng}$ at time t predicted z years before (in the empirical analysis, $z = 5$) and the expectation error $n_{jt}^{yng} - \breve{n}_{jt}^{yng}$, which is the difference between the actual and predicted numbers. A similar procedure is applied to the old population ratio.

$$n_{jt}^{yng} = \underbrace{\breve{n}_{jt}^{yng}}_{\text{predicted}} + \underbrace{(n_{jt}^{yng} - \breve{n}_{jt}^{yng})}_{\text{error}}. \tag{4}$$

The long-run relation is modified by plugging in the following analysis:

$$\begin{aligned}
\log P_{jt}^{rppi} &= \beta_{0j} + \beta_1 \log P_{jt}^{cpi} + \beta_2 \log\left(\frac{Y_{jt}}{pop_{jt}^{wrk}}\right) + \beta_3 i_{jt} + \beta_4 \log pop_{jt}^{total} \\
&+ \beta_5(\breve{n}_{jt}^{yng} - \breve{n}_{jt}^{wrk}) + \beta_6(\breve{n}_{jt}^{old} - \breve{n}_{jt}^{wrk}) \\
&+ \beta_7\{(n_{jt}^{yng} - \breve{n}_{jt}^{yng}) - (n_{jt}^{wrk} - \breve{n}_{jt}^{wrk})\} \\
&+ \beta_8\{(n_{jt}^{old} - \breve{n}_{jt}^{old}) - (n_{jt}^{wrk} - \breve{n}_{jt}^{wrk})\} \\
&+ \beta_9 i_{jt} \times (\breve{n}_{jt}^{yng} - \breve{n}_{jt}^{wrk}) + \beta_{10} i_{jt} \times (\breve{n}_{jt}^{old} - \breve{n}_{jt}^{wrk}) \\
&+ \beta_{11} i_{jt} \times \{(n_{jt}^{yng} - \breve{n}_{jt}^{yng}) - (n_{jt}^{wrk} - \breve{n}_{jt}^{wrk})\} \\
&+ \beta_{12} i_{jt} \times \{(n_{jt}^{old} - \breve{n}_{jt}^{old}) - (n_{jt}^{wrk} - \breve{n}_{jt}^{wrk})\} + \varepsilon_{jt}.
\end{aligned} \tag{5}$$

This allows the effects of the predictable and unpredictable parts of the population ratios on residential property prices to be decomposed and analyzed.

3 Data and Methodology

3.1 Data

Following the global financial crisis in 2008, internationally comparable property price indexes have been developed, led by the IMF and BIS, with participation from the United Nations and OECD.*21 This study covers the 17 countries in the four regions indicated below, for which it is possible to obtain BIS data. The international panel data cover a wide range of countries, rather than just Western countries: three from the Asia-Pacific region (Australia, Japan, New Zealand), two from North America (Canada, the United States), 11 from Europe (Belgium, Switzerland, Germany, Denmark, France, the United Kingdom, Ireland, Italy, the Netherlands, Norway, Sweden), and one from Africa (South Africa) (see Table 1). We conduct the analysis using balanced panel data for these 17 countries over 46 years from 1974 to 2019.

We use the Residential Property Price Index (local currency denominated in nominal terms) published by the BIS. As this index is published quarterly, we use the simple average for each year. For nominal interest rates, the main source is the long-term interest rates from OECD Statistics. However, as data for Denmark, Italy, Japan, Norway, and Sweden are missing for a part of the study period, data on interest rates, government securities, and government bonds obtained from the International Financial Statistics by the IMF are used instead. We

*21 Led by the IMF and BIS and administered by Eurostat, a handbook was created to show the procedure for generating internationally comparable property price indexes. Two of this study's authors participated in this project. On the basis of the handbook, various national statistics offices have developed property price indexes as public statistics, recorded in a BIS database. For more on the development process and creation method of these internationally comparable property price indexes, see Diewert et al. (2020)[7].

Table 1 Complete List of Countries/Regions in Our Sample

Region	Country	Region	Country
Asia-Pacific (3)	Australia (AU)	Europe (11)	Belgium (BE)
	Japan (JP)		Switzerland (CH)
	New Zealand (NZ)		Germany (DE)
			Denmark (DK)
America (2)	Canada (CA)		France (FR)
	United States (US)		United Kingdom (GB)
			Ireland (IE)
Africa (1)	South Africa (ZA)		Italy (IT)
			Netherlands (NL)
			Norway (NO)
			Sweden (SE)

use values obtained by converting these nominal interest rates (annual rates) into continuous compounded interest rates in the regression analysis. For real GDP, we use the real GDP (local currency unit) published in the World Bank's World Development Indicators (WDI). The CPI is likewise obtained from the WDI.

For the population-related variables, we aggregate population data by country and age cohort (obtained from the United Nations' World Population Prospects database) into three generations, young (aged 0–14), working age (15–64), and old (65+), for each country and calculate the population ratios with respect to the total population. Total population data are also used as an explanatory variable.

In Equation (4), a realized population ratio is decomposed into a predicted ratio of some interval ago and a prediction error. The selection of this prediction interval is an empirical decision.[*22] To capture housing investment's characteristics (time lag between planning and construction start/completion) as well as consider the sample period required for the estimation, we select a five-year prediction interval. In the analysis of Equations (2) and (3), we use panel data for the 17 countries from 1974 to 2019. However, because the United Nations' country-level population projection data are available only from 1982, this inevitably results in the usable sample period of Equation (5) starting after 1982.[*23] Accordingly, the analysis of Equation (5) is based on a 31-year panel for the 17 countries from 1989 to 2019.

To understand the nature of these aggregated data, we perform two panel unit root tests: the Im–Pesaran–Shin (IPS) test and the Fisher-type augmented Dickey–Fuller (ADF–Fisher) test (Im et al., 2003[17]). The tests are performed based on the following regression model:

$$\Delta y_{it} = \rho_i y_{i,t-1} + \sum_{L=1}^{p_i} \theta_{iL} \Delta y_{i,t-L} + \alpha_i d_t + \varepsilon_{it}. \tag{6}$$

[*22] The procedure for creating the five-year prediction data for the population ratio is explained in Appendix A-3.

[*23] The United Nations publishes projections for individual countries along with the global population every two years (with some exceptions). Projections are made in five-year intervals. To date, population updates have been made in 1982 , 1984, 1988, 1990, 1992, 1994, 1996, 1998, 2000, 2002, 2004, 2006 , 2008, 2010, and 2012. For each update, the United Nations publishes population projections in five-year intervals (e.g., 1985, 1990, 1995, 2000, 2005, 2010, and 2015). By using these data, approximate predictions for each year based on linear interpolation can be calculated. We therefore analyze the actual population values in relation to past predictions and prediction errors over a certain period. For the method of calculating the five-year population ratio projections used in the analysis, see Appendix 2.

where d_t is a term representing the deterministic factors such as a constant and a trend. The null and alternative hypotheses are as follows:

$$H_0 : \rho_i = 0 \text{ for all } i \quad \text{vs} \quad H_1 : \rho_i < 0 \text{ for at least one } i$$

Thus, the null hypothesis assumes that all series are a non-stationary process, whereas the alternative hypothesis assumes that a proportion of the series are stationary.

Table 2 Panel Unit Root Tests

	IPS W test		ADF-Fisher test	
	Level	1st Difference	Level	1st Difference
$\log P^{rppi}$	−0.495	−10.364***	37.722	178.237***
$\log P^{cpi}$	−10.617***	−6.472***	189.104***	103.722***
$\log(Y/pop^{wrk})$	4.116	−18.205***	18.178	343.452***
i	5.136	−19.201***	4.621	363.352***
$\log pop^{total}$	7.496	−3.512***	12.576	68.326***
$n^{yng} - n^{wrk}$	−11.978***	−7.309***	218.792***	129.88***
$n^{old} - n^{wrk}$	0.0824	−1.376*	48.427*	46.837*

Note: ***, **, and * indicate rejection of the null hypothesis at 1%, 5%, and 10% significance levels, respectively. Lag length is selected by Schwarz information criteria. Andrews automatic bandwidth selection and Quadratic Spectral kernel are used. All test regression includes individual effects as the exogenous variable. Test regression of $\log P^{cpi}, n^{yng} - n^{wrk}, n^{old} - n^{wrk}$ add individual linear trends for the level unit root tests. IPS indicates Im-Pesaran-Shin. ADF indicates augmented Dickey-Fuller.

Table 2 summarizes the test results. The IPS and ADF–Fisher tests reach identical conclusions. Specifically, the test results show that residential property prices, working-age per-capita real GDP, nominal interest rates, and total population are integrated of order I(1). By contrast, the test results of the CPI are stationary of order I(0) with a constant and linear time trend. By definition, population ratios should be stationary processes, but the test results are mixed, possibly due to the small sample.[*24] Based on the IPS and ADF–Fisher tests, the null hypothesis is rejected if the series is differenced, suggesting the possibility that I(0) and I(1) processes are mixed. In either case, it is acceptable to consider that the maximum order of integration is 1 for all the variables used in this study. Because of this mixed order of integration, the panel autoregressive distributed lag (ARDL) approach is an appropriate framework for the following investigation.

3.2 Methodology

The procedure for testing and estimating the panel cointegration relations is as follows. First, we perform a panel cointegration test based on residuals by checking the stationarity of

[*24] As the population ratio variables have values that are restricted to the interval [0,1], by definition, they are stationary processes. However, with panel unit root tests such as in this case, local non-stationarity sometimes cannot be rejected. This is convenient for estimating a long-run equilibrium relation model that includes population ratios in the explanatory variables, which should be stationary processes under normal circumstances, as in this study. While the applications differ from the present study, Pedroni (2007)[34] and Cavalcanti et al. (2011)[3] are also able to estimate panel cointegration relations, including ratio variables (specifically, investment–income ratios) since they demonstrate local non-stationary processes.

the residuals and testing the presence of a cointegrating relation, where the null hypothesis is the "absence of a cointegrating relation." Here, we use the Kao test and Pedroni test as representative tests (Kao, 1999[20]; Pedroni, 1999[32], 2004[33]). The Kao test assumes that all cointegration vectors are common to each country. We assume the commonality of the discounted PVR (housing market no-arbitrage condition) across all the countries; therefore, the Kao test's assumption is possible. However, since the commonality assumption is typically an extremely strong hypothesis, the cointegration vectors may differ. In that case, it may be considered inappropriate to apply the Kao test.[*25] Using these in combination, it is possible to empirically demonstrate whether the variables showing the discounted PVR in Equation (1) are in a long-run equilibrium relation, or, at least, whether it is impossible to observe a long-run relationship between the housing markets in the 17 countries covered in this study (i.e., whether there is no cointegrating relation).

Next, we estimate the long-run relationship based on the panel ARDL approach (Pesaran et al., 1999[35]):

$$y_{it} = \mu_i + \sum_{j=1}^{p} \lambda_{ij} y_{i,t-j} + \sum_{j=0}^{q} \delta'_{ij} X_{i,t-j} + \varepsilon_{it}. \tag{7}$$

Equation (7) is a typical ARDL(p, q) model, where p is the maximum lag length of the dependent variable y_{it} and q is the maximum lag length of the explanatory variables $X_{i,t}$.[*26]

Since Equation (7) includes I(1) variables, one can derive its error correction form as

$$\Delta y_{it} = \theta_i (y_{i,t-1} - \beta'_i X_{i,t-1}) + \sum_{j=1}^{p-1} \lambda^*_{ij} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \delta^{*\prime}_{ij} \Delta X_{i,t-j} + \mu_i + \varepsilon_{it}, \tag{8}$$

which is called a mean group (MG) model. The first term in the right-hand side of Equation (8) corresponds to the long-run equilibrium relationship, whereas the second and third terms capture the short-run adjustment processes. The parameters in Equations (7) and (8) are associated as follows: $\theta_i = -(1 - \sum_{j=1}^{p} \lambda_{ij})$, $\beta_i = \sum_{j=0}^{q} \delta_{ij} / (1 - \sum_{k=0}^{q} \lambda_{ik})$, $\lambda^*_{ij} = -\sum_{m=j+1}^{p} \lambda_{im}$, and $\delta^*_{ij} = -\sum_{m=j+1}^{q} \delta_{im}$. All the coefficients vary by country. For example, the coefficient θ_i in Equation (8) represents the speed of the adjustment of equilibrium errors, which varies by country. If a long-run equilibrium relationship exists between the variables, the sign of the coefficient θ_i may be expected to be negative and statistically significant. Given that the long-run equilibrium coefficient or cointegration coefficient β_i also varies by country, this specification permits variation in the coefficients of the variables in the levels included in the discounted PVR, such as the elasticity value (degree of money illusion) of nominal residential property prices with respect to the CPI. Moreover, the coefficients of the variables in differences expressing short-run adjustment vary by country. If housing market structure/policy and consumer preferences vary, the short-run adjustment process and correction of equilibrium errors also vary by country.

It is worthwhile to examine whether the theoretically predicted homogeneity restriction on the coefficient of the long-run relationship is valid. With regard to this, the pooled mean group (PMG) model is estimated:

$$\Delta y_{it} = \theta_i (y_{i,t-1} - \beta' X_{i,t-1}) + \textstyle\sum_{j=1}^{p-1} \lambda^*_{ij} \Delta y_{i,t-j} + \sum_{j=0}^{q-1} \delta^{*\prime}_{ij} \Delta X_{i,t-j} + \mu_i + \varepsilon_{it}. \tag{9}$$

[*25] By comparison, the Pedroni test may be considered more generally representative than the Kao test, since it permits cointegration vectors that vary by country.

[*26] It is technically possible to set different lag orders for each country; however, for simplicity, we use a common order.

Since Equation (9) is a non-linear model in terms of the parameter imposing the homogeneity restriction on the cointegration coefficients ($\beta_i' = \beta'$), maximum likelihood estimation is used (Pesaran et al., 1999[35]). This restriction may, at first glance, seem strong. However, for the 17 countries covered in the analysis, we can assume that the financial conditions are homogeneous based on the effects of no-arbitrage because of international capital flow. Econometrically, suppose the true long-run equilibrium coefficients are common among the sample countries. In that case, the PMG estimates will be more efficient than the MG estimates, making them preferable for the estimation. Conversely, if the true long-run equilibrium relation coefficients vary by country, the PMG estimates will not be consistent, but the MG estimates will. Therefore, we verify this by performing a Hausman test on the null hypothesis that the PMG model is more appropriate than the MG model, thereby testing the merits of a formulation that imposes homogeneity on the long-run equilibrium.

4 Estimation Results

This section summarizes the analytical results of the benchmark model (Equation (2)), two panel cointegration tests, and various panel ARDL specifications. We also report the empirical results based on the interaction between demographic composition and interest rates and its effect on residential property prices (Equation (3)) and the impact of persistent demographic expectation errors on residential property prices (Equation (5)).

4.1 Long-Run Equilibrium Relation Estimation: Statistically Significant Demographic Factors

First, we consider the panel cointegration test of Equation (2). The Kao test rejects the null hypothesis at the 1% level, indicating the presence of cointegration (see Table 3). Further, in

Table 3 Panel Cointegration Tests

	Statistic	Prob.
Kao test		
ADF t	-6.914^{***}	0.000
Pedroni tests: Within-dimension		
Panel v	1.897^{**}	0.029
Panel ρ	2.093	0.982
Panel PP	1.079	0.859
Panel ADF	-2.864^{***}	0.002
Pedroni tests: Between-dimension		
Group ρ	3.572	0.999
Group PP	2.142	0.984
Group ADF	-2.914^{***}	0.002

Note: ***, **, and * indicate the significance at 1%, 5%, and 10% levels, respectively. For Kao test: No deterministic trend. Lag length is 1, and the bandwidth is 2 with Bartlett Kernel. For Pedroni test: No deterministic trend. Automatic lag length selection on SIC with a maximum lag of 2. Bandwidth is 2 with Bartlett Kernel. PP indicates Phillips-Perron. ADF indicates augmented Dickey-Fuller.

the results using seven types of test statistics based on Pedroni, the panel v test and panel ADF test reject the null hypothesis that a cointegrating relation does not exist at the standard level of significance (see Table 3). The group ADF test also rejects the null hypothesis at the 1% level.*27 Based on these, we conclude that while some of the variables comprising the PVR include I(1) processes, we cannot rule out a long-run relationship between these variables.

Table 4 shows the PMG and MG model estimates of the long-run parameters using panel data from 1974 to 2019 for the 17 countries.*28 At first glance, the coefficients of the MG model are largely nonsignificant, whereas the coefficients of the PMG model are significant with the expected signs. As shown at the bottom of Table 4, the Hausman test statistic is 10.53 and its p value is 0.1042.*29 Therefore, for the housing markets of the countries covered in our analysis at least, the null hypothesis cannot be rejected, and the results support the commonality of the long-run equilibrium coefficients between the housing markets. On the basis of these results, we present the estimation results of the PMG model below.

Table 4 Estimation Results of Baseline Model, Equation (9)

	Parameters	PMG		MG	
$\log P^{cpi}$	β_1	1.080***	(0.102)	0.023	(0.434)
$\log(Y/pop^{wrk})$	β_2	0.410**	(0.201)	0.845	(1.213)
i	β_3	−8.705***	(1.038)	−8.056**	(3.653)
$\log pop^{total}$	β_4	1.153***	(0.426)	1.745	(1.697)
$n^{yng} - n^{wrk}$	β_5	5.579***	(0.853)	−8.326	(5.818)
$n^{old} - n^{wrk}$	β_6	−5.705***	(0.834)	0.566	(4.522)
error correction term		−0.111***	(0.023)	−0.309***	(0.038)
N		782		782	
$\log L$		1606.4		1808.1	
Hausman test					
Statistic		10.53			
p-value		0.1042			

Note: This table reports the estimate of long-run coefficients and error-correction term. The column labeled "Parameters" corresponds to the parameters of Equation (9). Standard errors are in parentheses. ***, **, and * indicate the significance at 1%, 5%, and 10% levels, respectively. The lag of the autoregressive distributed lag model, $p = 2$ and $q = 1$, is selected by Schwarz information criteria. PMG indicates pooled mean group. MP indicates mean group. The Hausman statistic refers to the test statistic on the long-run homogeneity restriction, examining if the PMG estimator should be preferred to the MG estimator.

As shown in Table 4, first, the CPI coefficient is approximately 1, suggesting that money illusion does not exist. Second, the coefficient of working-age per-capita real GDP is 0.410. This is a proxy variable of housing rent, and the fact that this coefficient is positive is consistent with discounted present value theory. Third, the estimated value of the impact of the nominal interest rate is −8.705, which is also significant at the 1% level. This result is likewise

*27 According to Pedroni (2004)[33], if the sample size in the time series dimension is less than 100, as in this study, the group ADF test and panel ADF test have the greatest test power.

*28 The ARDL model's lag order is taken as $p = 2, q = 1$, based on the Schwarz information criteria.

*29 The test statistics here follow a chi-square distribution with 6 degrees of freedom based on the null hypothesis.

consistent with the theoretical prediction based on the present value model wherein interest rate increases lower asset prices. Finally, and most importantly, with regard to the population ratio coefficients, the coefficient of the young population ratio is 5.579, whereas the coefficient of the old population ratio is −5.705. These coefficients are statistically significant at the 1% level. Thus, if the other parameters are constant, a 1% increase in the young population ratio increases residential property prices by 5.579%; conversely, a 1% increase in the old population ratio has an equivalent downward effect on residential property prices.

4.2 Long-Run Equilibrium Interaction Between Demographic Composition and Nominal Interest Rates

As before, we perform a PMG estimation of Equation (3), based on the ARDL specification. The estimation results are shown in Table 5. The interest rate coefficient is nonsignificant, whereas the coefficient of the interest rate and old population ratio interaction term is significant.

Table 5 Estimation Results for Model Including Interaction Terms, Equation (3)

	Parameters	PMG		MG	
$\log P^{cpi}$	β_1	1.053***	(0.089)	0.312	(0.278)
$\log(Y/pop^{wrk})$	β_2	0.602***	(0.190)	0.963	(1.194)
i	β_3	3.949	(5.920)	−29.438	(86.637)
$\log pop^{total}$	β_4	1.402***	(0.414)	1.519	(1.570)
$n^{yng} - n^{wrk}$	β_5	6.262***	(1.191)	−22.389**	(10.728)
$n^{old} - n^{wrk}$	β_6	−6.207***	(0.924)	13.693	(10.255)
$i \cdot (n^{yng} - n^{wrk})$	β_7	−8.526	(6.358)	158.237*	(91.106)
$i \cdot (n^{old} - n^{wrk})$	β_8	28.291**	(11.699)	−134.716	(181.930)
error correction term		−0.130***	(0.023)	−0.372***	(0.046)
N		782		782	
$\log L$		1638.3		1900.9	
Hausman test					
Statistic			11.42		
p-value			0.1789		

Note: This table reports the estimate of long-run coefficients and error-correction term. Standard errors are in parentheses. ***, **, and * indicate the significance at 1%, 5%, and 10% levels, respectively. The lag of the autoregressive distributed lag model, $p = 2$ and $q = 1$, is selected by Schwarz information criteria. PMG indicates pooled mean group. MP indicates mean group. Hausman statistic is for testing the null hypothesis of PMG as a correct specification against MG.

Because of the interaction terms in Equation (3), the impact of interest rate cuts on residential property prices depends on population composition conditions. To evaluate the magnitude numerically, we derive the marginal effect using the coefficient estimates of Equation (3) as follows.

First, define the average marginal effect of interest rate changes on the residential property price of country j over time:

$$\frac{\partial \log P_j^{rppi}}{\partial i_j} = \beta_3 + \beta_7 \overline{n_j^{yng}} - (\beta_7 + \beta_8)\overline{n_j^{wrk}} + \beta_8 \overline{n_j^{old}}, \tag{10}$$

where $\overline{n_j^{yng}}, \overline{n_j^{wrk}}$, and $\overline{n_j^{old}}$ are the historical averages for each population ratio in country j. Then, the total marginal effect of interest rate changes on residential property prices is

$$\frac{\partial \log P_{jt}^{rppi}}{\partial i_{jt}} = \frac{\partial \log P_{j}^{rppi}}{\partial i_{j}} + \beta_7 \widetilde{n_{jt}^{yng}} - (\beta_7 + \beta_8)\widetilde{n_{jt}^{wrk}} + \beta_8 \widetilde{n_{jt}^{old}}, \tag{10a}$$

where

$$\widetilde{n_{jt}^{k}} = n_{jt}^{k} - \overline{n_j^{k}} \text{ for } k = yng, wrk, \text{ and } old$$

Using the estimation results in Table 5, Equation (10) is rewritten as follows:

$$\frac{\widehat{\partial \log P_{j}^{rppi}}}{\partial i_{j}} = 3.949 - 8.526\,\overline{n_j^{yng}} - 19.765\,\overline{n_j^{wrk}} + 28.291\,\overline{n_j^{old}}. \tag{10b}$$

Table 6 shows the average marginal effect of interest rate increases on residential property prices by country, calculated according to the average values of the population ratios for the sample period (1974 to 2019) using Equation (10b). The average population ratios for the 17 countries during this period are 20.4% for the young population, 65.4% for the working-age population, and 14.1% for the old population. The average marginal effect of interest rate

Table 6 Effect of 1% Nominal Interest Rate Increase

	historical average shares (%)			coefficients of
	n^{yng}	n^{wrk}	n^{old}	interest rate
AU	21.8	66.2	11.9	−7.635
BE	18.2	65.7	16.1	−6.045
CA	19.8	68.0	12.2	−7.714
CH	17.4	67.3	15.3	−6.512
DE	15.9	66.9	17.2	−5.752
DK	18.5	65.7	15.8	−6.165
FR	20.0	64.4	15.6	−6.082
GB	19.2	64.8	16.0	−5.980
IE	25.0	63.9	11.2	−7.646
IT	16.7	66.2	17.1	−5.724
JP	17.3	66.6	16.1	−6.152
NL	19.1	67.1	13.8	−7.045
NO	19.9	64.7	15.4	−6.162
NZ	23.3	65.0	11.7	−7.565
SE	18.2	64.2	17.6	−5.321
US	21.5	65.9	12.6	−7.325
ZA	35.6	60.0	4.4	−9.713
Average	20.4	65.4	14.1	−6.738

Note: The historical average share of generations are calculated by using the data from the UN's World Population Prospects for the period from 1974 to 2019. The average marginal effect of interest rate change (the numbers in the rightmost column) is calculated using Equation (10b). AU: Australia, BE: Belgium, CA: Canada, CH: Switzerland, DE: Germany, DK: Denmark, FR: France, GB: the United Kingdom, IE: Ireland, IT: Italy, JP: Japan, NL: Netherlands, NO: Norway, NZ: New Zealand, SE: Sweden, US: the United States, ZA: South Africa.

increases obtained by plugging in these values is the value at the bottom of Table 6, -6.738. This result signifies that a 1% interest rate cut raises residential property prices by 6.738% on average, which is somewhat smaller than the result obtained with the benchmark model in Section 4.1 (8.705%).

With regard to individual countries, in South Africa, where the average ratio for the young population is markedly high at 35.6%, a 1% interest rate decrease increases residential property prices by 9.713%. Similarly, in Ireland, which has the next highest young population ratio (25.0%), the figure is 7.646%, and in New Zealand, which has the third-highest ratio (23.3%), it is 7.565%; therefore, the impact on the property market is considerable.

The opposite phenomenon occurs in countries with a high old population ratio. Among the sample countries, Sweden has the highest ratio at 17.6%, followed by Denmark (17.2%) and Italy (17.1%). In these countries, a 1% interest rate cut raises residential property prices only by 5.321% (in Sweden) to 5.752% (in Denmark), which is about half the extent of the increase in South Africa. These figures show the effect of monetary easing calculated on the basis of historical averages.

The key takeaway is that divergence from historical averages further enhances the impact of the above monetary measures on the residential property market. Using the estimates in Table 5, Equation (10a) is rewritten as follows:

$$\frac{\partial \widehat{\log P_{jt}^{rppi}}}{\partial i_{jt}} = \frac{\partial \widehat{\log P_{j}^{rppi}}}{\partial i_{j}} - 8.526\widetilde{n_{jt}^{yng}} - 19.765\widetilde{n_{jt}^{wrk}} + 28.291\widetilde{n_{jt}^{old}}. \tag{10c}$$

Equation (10c) implies that an increase in the young age population ratio ($\widetilde{n_{jt}^{yng}} > 0$) enhances the interest rate effects (since $\beta_7 = -8.526 < 0$), while an increase in the old age population ratio ($\widetilde{n_{jt}^{old}} > 0$) reduces the interest rate effects (since $\beta_8 = 28.291 > 0$). In other words, population bonuses considerably strengthen the positive effect of interest rate cuts (monetary expansion), while population onuses considerably reduce the positive effect of interest rate cuts. These findings strongly support the hypothesis of a strong interaction between population statistics and monetary policy proposed by Nishimura (2016)[29] and Nishimura and Takáts (2012)[30].

4.3 Persistent Demographic Expectation Errors and Long-Run Equilibrium

The estimation results of Equation (5) are summarized in Table 7.[*30] There are statistically significant changes in the estimation values such as the CPI coefficient (1.386), which is higher than the estimates in the previous sections, and the coefficient of working-age per-capita GDP almost doubling (from 0.410 to 0.973). Nonetheless, in each case, there is no qualitative change in the interpretation of these variables, except for the coefficient of the total population being negative. The shortened estimation interval may also have had an influence.

Equation (5) implies that the marginal effect of interest rate changes on the residential

[*30] $p = 2, q = 1$ is selected as the optimum lag length for ARDL based on the Schwarz information criteria.

property prices of country j is as follows:[*31]

$$\begin{aligned}\frac{\partial \log P_{jt}^{rppi}}{\partial i_{jt}} &= -1.886 - 33.013\, \breve{n}_{jt}^{yng} - 7.001\, \breve{n}_{jt}^{wrk} + 40.014\, \breve{n}_{jt}^{old} \\ &\quad - 413.758\, (n_{jt}^{yng} - \breve{n}_{jt}^{yng}) - 105.990\, (n_{jt}^{wrk} - \breve{n}_{jt}^{wrk}) \\ &\quad + 519.748\, (n_{jt}^{old} - \breve{n}_{jt}^{old}). \end{aligned} \tag{11}$$

Thus, the marginal effect of interest rate changes has seven parts: constant, three predicted population ratios $\breve{n}_{jt}^{yng}, \breve{n}_{jt}^{wrk}$, and $\breve{n}_{jt}^{old}$, and three prediction errors, $n_{jt}^{yng} - \breve{n}_{jt}^{yng}, n_{jt}^{wrk} - \breve{n}_{jt}^{wrk}$, and $n_{jt}^{old} - \breve{n}_{jt}^{old}$.

Table 7 Estimation Results for Model Including Population Prediction Errors, Equation (5)

	Parameters	PMG	
$\log P^{cpi}$	β_1	1.386***	(0.273)
$\log(Y/pop^{wrk})$	β_2	0.973***	(0.229)
i	β_3	−1.886	(9.764)
$\log pop^{total}$	β_4	−1.884**	(0.916)
$\breve{n}^{yng} - \breve{n}^{wrk}$	β_5	19.374***	(2.759)
$\breve{n}^{old} - \breve{n}^{wrk}$	β_6	−8.630***	(1.243)
$(n^{yng} - \breve{n}^{yng}) - (n^{wrk} - \breve{n}^{wrk})$	β_7	39.702***	(4.554)
$(n^{old} - \breve{n}^{old}) - (n^{wrk} - \breve{n}^{wrk})$	β_8	−33.230***	(4.237)
$i \cdot (\breve{n}^{yng} - \breve{n}^{wrk})$	β_9	−33.013*	(17.989)
$i \cdot (\breve{n}^{old} - \breve{n}^{wrk})$	β_{10}	40.014**	(15.551)
$i \cdot \{(n^{yng} - \breve{n}^{yng}) - (n^{wrk} - \breve{n}^{wrk})\}$	β_{11}	−413.758***	(48.649)
$i \cdot \{(n^{old} - \breve{n}^{old}) - (n^{wrk} - \breve{n}^{wrk})\}$	β_{12}	519.748***	(90.639)
error correction term		−0.092**	(0.042)
N		527	
$\log L$		1401.217	

Note: This table reports the estimate of long-run coefficients and error-correction term. The column labeled "Parameters" corresponds to the parameters of Equation (5). Standard errors are in parentheses. ***, **, and * indicate the significance at 1%, 5%, and 10% levels, respectively. The lag of the autoregressive distributed lag model, $p = 2$ and $q = 1$, is selected by Schwarz information criteria. PMG indicates pooled mean group.

The second term of Equation (11) represents the change in residential property prices with respect to the proportion of the young population ratio increase predicted in advance. The sign of the estimate, −33.013, suggests that interest rate cuts with an increase in the predicted young population ratio cause additional upward pressure on residential property prices, and the coefficient is significant at the 10% level. In addition, the fourth term suggests that a predicted

[*31] Using Equation (5), the marginal effect of interest rate changes on the residential property prices of country j is derived as

$$\frac{\partial \log P_{jt}^{rppi}}{\partial i_{jt}} = \beta_3 + \beta_9 \breve{n}_{jt}^{yng} - (\beta_9 + \beta_{10})\breve{n}_{jt}^{wrk} + \beta_{10}\breve{n}_{jt}^{old} + \beta_{11}(n_{jt}^{yng} - \breve{n}_{jt}^{yng}) - (\beta_{11} + \beta_{12})(n_{jt}^{wrk} - \breve{n}_{jt}^{wrk}) + \beta_{12}(n_{jt}^{old} - \breve{n}_{jt}^{old}).$$

Substituting the corresponding estimates from Table 7, Equation (11) is obtained.

increase in the old population ratio decreases the residential property price-increasing effect of interest rate cuts even if they were predicted and their impact is statistically significant. This result is consistent with our findings in the previous sections that population onuses decrease the positive effect on residential property prices.

Finally, and most importantly, we consider the impact of the supply and demand mismatch due to prediction errors based on the fifth and seventh terms. As the sign of the coefficient of the young population ratio prediction error is negative, unforeseen increases in the young population amplify the effect of interest rate cuts. Conversely, the sign is positive for the old population ratio-related prediction errors; therefore, if the aging of the population proceeds more rapidly than expected, the effect of interest rate cuts on increasing residential property prices decreases further.

Further, on comparing the prediction error coefficients, we find that old population (519.748) prediction errors have a greater impact than young population errors (−413.758) on the marginal effect of interest rate cuts in absolute value. These implications of prediction errors are likewise consistent with the analysis results in the previous sections.

Table 8 Marginal Effect of Nominal Interest Rate i

	historical average (%)						coefficients of interest rate
	predicted ratios			prediction errors			
	n^{yng}	n^{wrk}	n^{old}	$n^{yng} - \breve{n}^{yng}$	$n^{wrk} - \breve{n}^{wrk}$	$n^{old} - \breve{n}^{old}$	
AU	20.2	66.8	13.0	0.06	−0.09	0.03	−8.033
BE	17.1	65.9	17.0	0.23	−0.20	−0.03	−6.232
CA	18.3	68.1	13.6	−0.12	0.18	−0.06	−7.250
CH	15.9	67.5	16.6	0.32	0.14	−0.47	−9.157
DE	14.6	67.2	18.2	0.11	−0.22	0.11	−3.790
DK	17.5	66.1	16.3	0.15	−0.16	0.02	−6.108
FR	18.8	64.9	16.3	0.09	−0.34	0.25	−4.781
GB	18.3	65.3	16.4	0.05	−0.19	0.14	−5.255
IE	22.8	65.7	11.5	−0.34	0.50	−0.16	−9.350
IT	14.7	66.6	18.7	−0.23	−0.11	0.34	−1.102
JP	15.0	65.7	19.2	−0.37	0.20	0.17	−1.581
NL	17.6	67.5	14.9	0.20	−0.17	−0.02	−7.204
NO	18.9	65.3	15.9	0.16	0.02	−0.17	−7.893
NZ	21.5	66.0	12.5	0.13	−0.28	0.15	−8.057
SE	17.3	64.4	18.3	0.38	−0.20	−0.18	−7.101
US	20.8	66.1	13.1	0.03	−0.06	0.02	−8.067
ZA	33.7	61.7	4.6	−0.70	0.70	0.00	−13.324
Average	19.0	65.9	15.1	0.01	−0.02	0.01	−6.723

Note: The historical average are calculated by using the data from the UN's World Population Prospects for the period from 1989 to 2019. The average marginal effect of interest rate change, the numbers in the rightmost column, are calculated by using Equation (11). AU: Australia, BE: Belgium, CA: Canada, CH: Switzerland, DE: Germany, DK: Denmark, FR: France, GB: the United Kingdom, IE: Ireland, IT: Italy, JP: Japan, NL: Netherlands, NO: Norway, NZ: New Zealand, SE: Sweden, US: the United States, ZA: South Africa.

5 Discussion: Demographics, Residential Property Prices, and Credit Conditions

In this study, focusing on residential property, which, as a means of wealth accumulation, represents the largest share of household assets regardless of country, we empirically clarify the relationships among changes in demographics (including the aging of the population), nominal interest rates determined by monetary policy, and residential property prices.*32 We also identify the importance of the formation of demographic expectations: a considerable difference exists between expected and unexpected changes in demographics.

The large set of empirical analyses in Section 3 demonstrate that residential property price changes form a PVR in the long term and are determined as fundamental prices. These empirical findings reveal the relationships among residential property prices, demographics, and nominal interest rates. Additionally, we find significant differences among the countries in the short-run dynamics.*33 In this section, we once again verify the consistency of our findings with the existing literature.

First, in Section 4.1, we show that population factors are key variables for a PVR for long-term changes in residential property prices, based on long-term panel data covering a diverse range of circumstances in 17 countries over 46 years. In other words, we find that residential property prices are determined by population composition ratios, in addition to working-age per-capita GDP (a proxy variable for rents) and nominal interest rates. As most of these variables demonstrate unit root process characteristics, we conduct an analysis that treats the long-term relationship between the variables as a cointegrating relation. We find that if the ratio of the young population to the total population rises by 1%, residential property prices increase by 5.579%; conversely, if the old population ratio increases by 1%, residential property prices drop by 5.705%.

Further, in Section 4.2, we estimate the relationship with credit conditions during property bubble and collapse periods by adding a cross-term interest rate and population factors into the model in Section 4.1. For example, in the case of Japan, whose property bubble that started in the mid-1980s has been dubbed the biggest of the 20th century, baby boomers entered the housing market in the early 1980s. They generated the most tremendous housing demand in the country's history, leading to a wave of excess optimism. Conversely, because of a decline in the working-age population following the bubble's collapse and then a decline in the total population in the 21st century, excess pessimism became prevalent. The country hit a period of long-term economic stagnation known as the "lost decade." When we consider the results not only for Japan but also for the 17 countries over 46 years, we find that the optimism caused by population bonuses and credit expansion conditions (typified by low-interest rates) have a synergistic effect on property demand and property bubbles, as emphasized by Nishimura (2016)[29]. Further, we determine that the effect of monetary easing measures such as lowering nominal interest rates is severely restricted during population onus (aging) periods and in the countries facing them.

As a criticism of various studies analyzing the relationship between demographics and the housing market, starting with Mankiw and Weil (1989)[25], research has pointed out that since

*32 We examine the case of real interest rates and real residential property prices in Appendix A-1 and find that our results in Section 3 are robust.

*33 In Appendix A-2, we analyze the accumulated responses of a one-unit shock to rents and interest rates and show significant differences in the short-run dynamics of the countries.

populations change only slowly if perfect foresight exists with regard to population changes and aging in the long term, price drops should not occur because of stock being adjusted through supply changes (Engelhard and Poterba, 1991[9]; Hamilton, 1991[12]; Hendershott, 1991[13]). However, if changes in population composition are worse than forecast, deflation will occur if production capacity adjustments (based on predictions) are too small. Therefore, in Section 4.3, in response to these criticisms, we expand the model to include the effect of the difference between population predictions and the actual populations (i.e., surprise) in various countries on the residential property price inflation rate (deflation rate).

As expected, the obtained results show that unforeseen increases in the young population ratio amplify the effect of interest rate cuts and raise residential property prices. Meanwhile, the sign of the prediction errors relating to the old population ratio in Equation (11) implies that the aging of the population, which advances faster than expected, will curb the effect of interest rate cuts in boosting residential property prices. The marginal effects of increases in interest rates on property prices by country, derived on the basis of the population ratio averages from 1989 to 2019, are presented in Table 8.[*34] Several interesting suggestions may be obtained from this table.

First, in the cases of Australia and the United States, three average prediction errors are effectively zero (or less than 0.1%), which shows that while the possibility of errors occurring at a given point during the period cannot be dismissed, in general, there is no bias in the predictions. Therefore, in these two countries, the relationship between demographic changes and interest rate cut effects may be determined by predictable factors (the second and third terms in Equation (11)).

Second, in Belgium, Canada, Denmark, the Netherlands, and South Africa, the average old population ratio prediction error is zero (or less than 0.1%), and there is no bias for the period. Nonetheless, prediction errors occur for the young population, which enhances the marginal impact of interest rate increases in the three European countries with positive prediction errors and decreases it in the other two countries with negative ones.

Third, for the countries with positive old population ratio prediction errors on average, the marginal effects of interest rate changes are as follows: Germany (average prediction error: 0.11%, marginal effect: 3.790%), France (0.25%, 4.781%), the United Kingdom (0.14%, 5.255%), Italy (0.34%, 1.102%), Japan (0.17%, 1.581%), and New Zealand (0.15%, 8.057%). Although the marginal effect is impacted by a combination of four population ratio factors and not by old ratio prediction errors alone, the fact that the marginal effect in five of these six countries (the exception being New Zealand) is significantly below the overall average of 6.723% may be considered important in terms of the impact that population aging that exceeds predictions has on monetary policy.

To summarize, both the coefficient signs of the young population are negative, while both those of the old population are positive, which is consistent with the series of estimation results and prediction results. Notably, among the two coefficients of the young population, the prediction error is substantial. By comparison, interestingly, the prediction error is also very large among the two coefficients of the old population. These results suggest that unforeseen changes in population composition have a considerable effect on the impact of interest rate cuts on property prices.

Results such as these also have implications for monetary policy. According to the Taylor rule (Taylor, 1993[41]), when the economy is struggling (i.e., when the GDP gap is negative),

[*34] The population ratios in the table are the averages for the sample period.

monetary policy shores up the economy by lowering interest rates. When the economy is thriving (i.e., when the GDP gap is positive), excessive growth is curbed by raising interest rates. However, this is not necessarily true for property prices for all countries for all periods. Thus, the finding that the effects of monetary policy are related to population composition (e.g., the degree of population aging) in various countries has important implications for policymakers.

6 Conclusion

From the estimation results of a large set of econometric models based on data of diverse countries in Europe, the Asia-Pacific region, Africa, and America, we reveal that fluctuations in property prices are determined by the PVR in the long run and are also strongly influenced by population dynamics. Focusing on these dynamics, we attest that during the "population bonus period," when the population and proportion of the working-age population increase, residential property prices soar as housing demand increases. This tendency is strengthened further when credit conditions are loose with low nominal interest rates. Moreover, if demographic changes are unanticipated, the effects of the credit conditions rise. By contrast, in the "population onus period," when the proportion of the old population increases substantially, residential property markets stagnate, and the loosening of credit conditions does not have as strong positive effects as it does during the population bonus period.

We also show that the interaction between demographic factors and credit conditions (nominal interest rates) determined by monetary policy varies across countries. This result has an important policy implication: different countries should have different policies to counter the undesirable effects of demographic changes and credit conditions on property prices.

However, some issues prevail. First, in the current analysis, the definitions of the working-age population and the old population are fixed and exogenous. In the future, increases in the rate of capital accumulation through investment in residential property and decreases in the rate of return could be slowed to some degree because of workers deferring retirement. Hence, we could incorporate the retirement decision when defining the working-age population.

Second, our model assumes a relatively homogeneous population. However, the population has become heterogeneous as immigration/emigration has become increasingly more important. As the population becomes heterogeneous, its composition may change over time. Such a change could be incorporated into future research.

Third, another kind of heterogeneity exists with respect to property markets. In fact, bi-polarization in residential property markets is underway: some parts of urban areas (superstar cities) experience rapidly rising residential property prices, while the rest of the country suffers declining prices. Interestingly, some researchers argue that the aging population has caused this bi-polarization of national property markets. This is an important direction for future research.

Fourth, our model uses nominal interest rates describing credit conditions. However, concern about the effective lower band of nominal interest rates has recently been growing, and central banks have been increasingly relying on quantitative easing. Incorporating this unconventional policy is also important in future research.

Finally, this study is descriptive and does not provide suggestions on the issue of resource and welfare distribution. Previous research (Hirano and Yanagawa, 2017[16]) has shown that productivity and economic growth rates do not return to their pre-bubble levels due to the impact of property bubble formation and collapse on resource distribution. The structure underlying this phenomenon could also be linked to demographics, which is an issue to examine in the future.

Online Appendices

A-1. Robustness Check: Analysis of the Real Relationship

Theoretically, it is possible to interpret a discounted PVR as a long-run equilibrium relationship between real variables (Walras, 1954[42]). In this study, we have thus far performed the empirical analysis using nominal residential property prices and interest rates. This is because, for real values, there are multiple definitions based on expectation formation hypotheses. Accordingly, in this appendix, to verify the robustness of the estimation results reported in this study, we create real variables based on two types of expectation formations that appear frequently in the empirical analysis and present the cointegration test and cointegration vector estimation results based on the unit root tests for the benchmark model used in Section 4.1.

The explained variable is the real residential property price index $realP^{rppi}$, which is deflated by the CPI, while the explanatory variables are working-age per-capita real GDP, the real interest rate r , and population factors:[*35]

$$\log\left(realP^{rppi}_{jt}\right) = \mu_j + \alpha_1 \log\left(\frac{Y_{jt}}{pop^{wrk}_{jt}}\right) + \alpha_2 r_{jt} + \text{population factors}_{jt} + \varepsilon_{jt}. \tag{A1}$$

Here, the real interest rate r_{jt} is defined using the following two formulas:

$$\text{Real interest rate based on static expectations:} \quad r^{SE}_{jt} = i_{jt} - \Delta \log P^{cpi}_{j,t}.$$
$$\text{Real interest rate based on perfect foresight:} \quad r^{PF}_{jt} = i_{jt} - \Delta \log P^{cpi}_{j,t+1}.$$

Further, population factors show an effect corresponding to the real expected rate of change in housing rents.

First, we consider the panel unit root test results. Table A1 summarizes the test results for the three new real variables. The real interest rate based on static expectations is r^{SE} and the real interest rate based on perfect foresight is r^{PF}. Using IPS and ADF–Fisher tests, just

Table A1 Panel Unit Root Test — Real Variables

	IPS W test		ADF-Fisher test	
	Level	1st Difference	Level	1st Difference
$\log realP^{rppi}$	2.491	−11.272***	16.876	194.775***
r^{SE}	−2.605***	−28.315***	55.795**	563.27***
r^{PF}	−7.439***	−28.405***	131.482***	553.03***

Note: ***, **, and * indicate the significance at 1%, 5%, and 10% levels, respectively. Lag length is selected by Schwarz information criteria. Andrews automatic bandwidth selection and Quadratic Spectral kernel are used. All test regression includes individual effects as the exogenous variable. IPS indicates Im-Pesaran-Shin. ADF indicates augmented Dickey-Fuller.

[*35] It is assumed that the expected inflation rate frequently used by market players is equal to the ex-post inflation rate.

as we did for the nominal variables, for real residential property prices and static expectation interest rates, the null hypothesis can be rejected at the 1% level for the first differenced series. Meanwhile, for the perfect foresight interest rate, the null hypothesis is rejected before differencing. Thus, it is acceptable to consider the maximum order of integration for these three series to be 1.

Next, we perform cointegration tests, again using two types of tests. Assuming that the cointegration vectors are homogeneous, the Kao test rejects the null hypothesis that no cointegrating relation is present at the 1% level for both interest rate models (Table A2). Further, using the Pedroni test, the null hypothesis is also rejected at the standard level of significance for the Panel ADF and Group ADF (Table A2). Judging by these results in combination, a cointegrating relation may be deemed to exist.

Table A2 Panel Cointegration Tests — Real Variables

	Static Expectation		Perfect Foresight	
	Statistic	Prob.	Statistic	Prob.
Kao test				
ADF t	-3.915^{***}	0.000	-4.302^{***}	0.000
Pedroni tests: Within-dimension				
Panel v	1.779^{**}	0.037	1.624^{*}	0.052
Panel ρ	0.377	0.647	1.362	0.913
Panel PP	-1.262	0.103	0.200	0.579
Panel ADF	-3.279^{***}	0.001	-2.988^{***}	0.001
Pedroni tests: Between-dimension				
Group ρ	2.235	0.987	3.253	0.999
Group PP	0.353	0.638	1.956	0.974
Group ADF	-2.516^{***}	0.006	-1.977^{**}	0.024

Note: ***, **, and * indicate the significance at 1%, 5%, and 10% levels, respectively. For Kao test: No deterministic trend. Lag length is 1, and the bandwidth is 2 with Bartlett Kernel. For Pedroni test: No deterministic trend. Automatic lag length selection on Schwarz information criteria with a maximum lag of 2. Bandwidth is 2 with Bartlett Kernel. PP indicates Phillips-Perron. ADF indicates augmented Dickey-Fuller.

Finally, the PMG model's long-term coefficient estimation results are summarized in Table A3. The coefficient of working-age per-capita real GDP (a proxy variable for real rent) is positive and consistent with the discounted PVR. However, the estimated value is approximately 2 in either model, which is rather high compared to the nominal model's value (0.606). Next, although the real interest rate coefficient is negative and statistically significant for the static expectation case, it is positive and nonsignificant for the perfect foresight case. This reflects the difficulty of creating suitable real interest rate data based on annual data. Meanwhile, the results obtained with regard to the population ratios are the same as for the nominal model.

Table A3 Estimation Results of Baseline Model with Real Residential Property Price

	Static Expectation		Perfect Foresight	
$\log(Y/pop^{wrk})$	1.848***	(0.118)	1.885***	(0.134)
r^{SE}	−1.044**	(0.522)		
r^{PF}			0.237	(0.495)
$\log pop^{total}$	1.045***	(0.219)	1.307***	(0.224)
$n^{yng} - n^{wrk}$	3.160***	(0.698)	3.978***	(0.732)
$n^{old} - n^{wrk}$	−3.315***	(0.692)	−3.486***	(0.686)
error correction term	−0.173***	(0.033)	−0.166***	(0.034)
N	782		765	
$\log L$	1544.0		1515.1	
Hausman Test				
Statistic	3.29		3.45	
p-value	0.6546		0.6305	

Note: This table reports the estimate of long-run coefficients and error-correction term. This table reports the estimation results by pooled mean group (PMG) estimator. Standard errors are in parentheses. ***, **, and * indicate the significance at 1%, 5%, and 10% levels, respectively. The lag of the autoregressive distributed lag model, $p = 2$ and $q = 1$, is selected by Schwarz information criteria. For brevity, the mean group (MG) estimation results are not reported. The Hausman statistic refers to the test statistic on the long-run homogeneity restriction, examining if the PMG estimator should be preferred to the MG estimator.

A-2. Residential Property Price Short-Run Adjustment Process

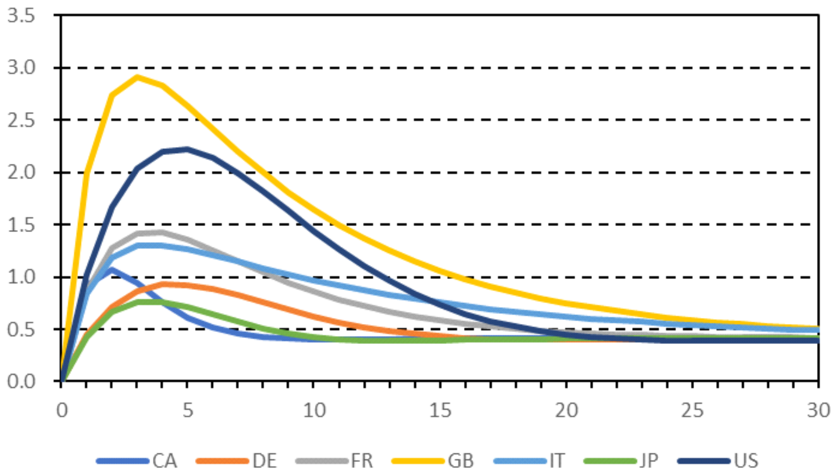
Thus far, the analysis has focused on the long-run equilibrium relation based on the discounted PVR. However, in the housing market, where transaction costs and information asymmetry exist, the fundamental value cannot be easily instantly realized. Therefore, based on the PMG estimation results of Equation (9) in Section 4.1, we analyze the adjustment path until residential property prices reach the long-run equilibrium when an exogenous shock occurs to residential property price fundamentals.

Using the estimates of the common long-run parameters, the country-specific short-run parameters of the PMG model, and successive substitutions of Equation (9), it is possible to express the residential property price of country j at time t as the sum of a deterministic component, past fundamental factors, and residential property price shocks of its own. Hence, the effect of a fundamental shock on residential property prices can be obtained by comparing two paths: one with a fundamental shock and another without a fundamental shock.*36 Here, we investigate the effects of two fundamental shocks: housing rent increases and nominal interest rate cuts. We use a single equation model, and the variables other than the series giving the shock and residential property prices are assumed to be constant.

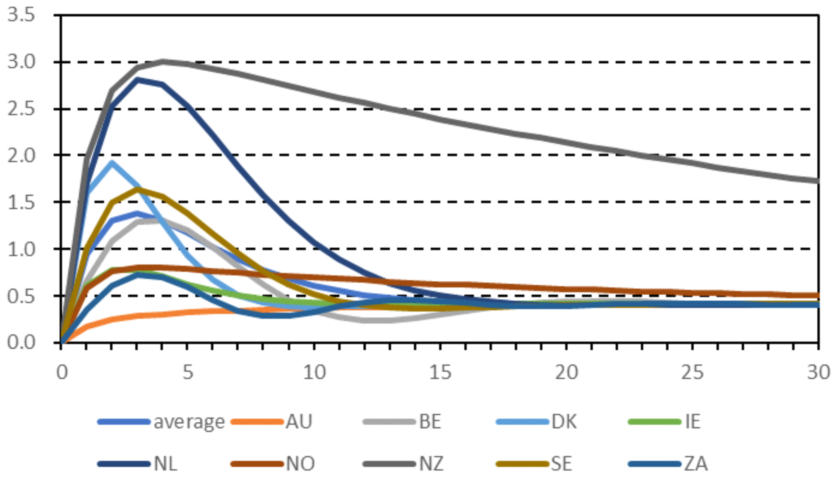
First, we consider the reaction of residential property prices to housing rent shocks. Figure A1 shows the accumulated effect on residential property prices when working-age per-capita GDP increases by one unit. The horizontal axis is the number of years since the shock

*36 For a similar analysis with a vector autoregressive model including exogenous variables, see Lutkepohl (1993, p. 327)[24].

occurred, while the vertical axis is the rate of increase in residential property prices. Based on the assumption of a long-run adjustment process, the values are illustrated over three decades.*37



(a) G7 Countries



(b) 10 Non-G7 Countries (Excluding Switzerland)

Figure A1 Accumulated Responses of a Positive One Unit Shock to $\log(Y/pop^{wrk})$ on Own $\log P^{rppi}$

Notes: The figure shows the reaction of nominal residential property prices in each country to a 1-unit increase in working-age per-capita real gross domestic product. Panel (a) is the reaction in G7 countries. Panel (b) shows the average for the 17 countries and the reaction of nine other countries (Switzerland is excluded). For abbreviations used in these figures, see Table 1.

*37 Among the 17 countries, only Switzerland showed divergent behavior, and it was therefore excluded from the figure.

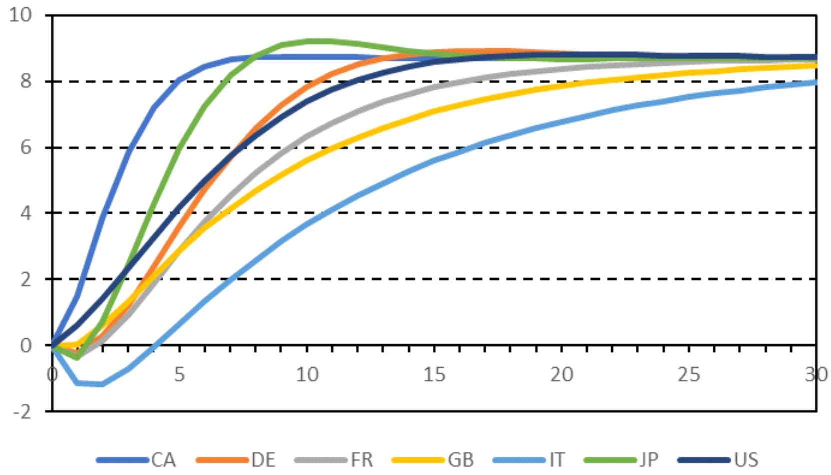
The first characteristic is that the rate of increase in residential property prices converges at around 0.4% in the long run. This is also the result expected according to the estimated long-run equilibrium relationship of the benchmark model. However, the second characteristic is that significant variation by country may be seen in the pattern of convergence to the long-run level. For example, countries other than Australia exhibit a residential property price overshoot. Moreover, the number of years required until the post-shock residential property price increase peaked is two years for Canada, Denmark, and Ireland; three years for the United Kingdom, the Netherlands, Norway, Sweden, and South Africa; four years for Belgium, France, Germany, Italy, Japan, and New Zealand; and five years for the United States.

Among G7 countries, the rate of increase is the highest in the United Kingdom, followed in order by the United States, France, Italy, Canada, Germany, and Japan.[*38] Increases in working-age per-capita GDP directly produce housing demand and, therefore, the overshoot observed in the reaction of residential property prices may be considered an understandable phenomenon.

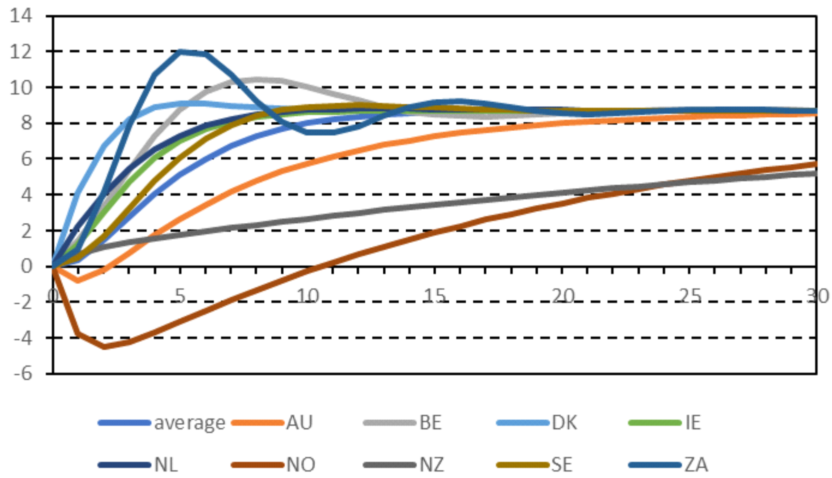
Next, we consider the decreasing effect of nominal interest rates. Figure A2 illustrates the effect of credit expansion, which is defined as a 1% decrease in the nominal interest rate. Switzerland, whose reaction path displays divergent tendencies, is omitted from this graph. In terms of characteristics observed from the graph, residential property prices rise in 10 countries (Belgium, Canada, Denmark, the United Kingdom, Ireland, the Netherlands, New Zealand, Sweden, the United States, and South Africa) immediately after the interest rate cut, and in these 10 countries, the increase continues for at least four years. In the other six countries, residential property prices decrease immediately after the interest rate cut, but in four countries (the exceptions being Italy and Norway), they stop decreasing after one year, while in Italy and Norway, they stop decreasing after two years.

Examining G7 countries in detail shows that the reaction differs from housing rent shocks. Compared with housing rent, one notable characteristic is that the adjustment of residential property prices in response to credit expansion requires a longer period. As the reduction in interest rates affects the information asymmetry between borrowers and lenders, the impact on housing demand is indirect. It is, therefore, understandable that the reaction to interest rate cuts is slow.

[*38] Since Figure A1 shows point estimates without confidence intervals, caution is required when interpreting the magnitude.



(a) G7 Countries



(b) 10 Non-G7 Countries (Excluding Switzerland)

Figure A2 Accumulated Responses of a Negative One Unit Shock to i on Own log P^{rppi}

Notes: The figure shows the reaction of nominal residential property prices in each country to a 1-unit decrease in the nominal interest rate. Panel a is the reaction in G7 countries. Panel b shows the average for the 17 countries and the reaction of nine other countries (Switzerland is excluded). For abbreviations used in these figures, see Table 1.

A-3. Procedure for Creating the Population Ratio Five-Year Prediction Data

The United Nations publishes its revisions of the World Population Prospects every two years (with some exceptions). To date, updates have been made in 1982, 1984, 1988, 1990, 1992, 1994, 1996, 1998, 2000, 2002, 2004, 2006, 2008, 2010, and 2012. Each revision publishes population projections for years ending in 0 or 5 (i.e., 1985, 1990, 1995, 2000, 2005, 2010, and 2015, among others). Table A4 shows the predicted and actual values for Australia's 0-to-14-year-old population ratio, obtained from the revised reports.

Column B shows the predictions for 1985, 1990, 1995, and 2000 in the 1982 revision. More long-term predictions exist; however, for this study, data are collected by taking 20 years in the future as the limit for long-term predictions. Past values relative to the 1982 update (e.g., values as of 1980) are actual values, not predictions.

Table A4 Predicted and Actual Values for the 0-to-14-year-old Population Ratio in Australia

	A	B	C	D	E	F	G	H	I	J	K	L
1		1982	1984	1988	1990	1992	1994	1996	1998	2000	2002	2004
2	1980	25.60	25.30									
3	1985	24.20	23.60	23.60								
4	1990	22.80	22.50	22.20	22.10	21.90	21.90					
5	1995	22.40	22.20	21.60	21.50	21.70	21.60	21.50	21.50			
6	2000	22.00	21.60	20.80	20.60	21.50	21.00	21.00	20.60	20.50	20.50	21.20
7	2005			20.10	19.80	21.40		20.40	19.60	19.60	19.40	19.60
8	2010		20.00	19.50	19.20	21.20	19.70	19.90	18.70	18.50	18.10	18.30
9	2015				18.80	20.60		19.50	18.30	18.00	17.30	17.60
10	2020						19.30	19.40	18.20	17.80	16.90	17.60
11	2025									17.70	16.80	17.60
12	2030											
13	2035											

Note: For brevity, only a subset of revision years (columns of this table) and a part of published population projections with five-year intervals (rows) are illustrated. For instance, the value of a cell B3 of this table (24.20) shows the predicted 0-to-14-year-old population ratio reported in the 1982 revision.

From these figures, approximate five-year predictions are calculated for all the years in the sample period. The calculation methods for the three cases are explained below.

[1] For revision years ending in 0 or 5

In this case, the five-year prediction published by the United Nations is used. For example, in the 1990 revision, the prediction for 1995 is 21.5%. Therefore, we use this value as a five-year prediction for 1995 as of 1990.

[2] For revision years that do not end in 0 and 5

We calculate the five-year prediction through linear interpolation based on the published five-year-interval predictions. To explain this using the 1982 update as an example, the prediction for 1987, five years after 1982, is calculated according to the predictions for 1985 (24.2) and 1990 (22.8) as follows: $24.2 + (22.8 - 24.2)/5 \times 2 = 23.64$.

Table A5 Prediction for Revision Years that Do Not End in 0 and 5, through Linear Interpolation

	A	B	C	D	E	F	G	H	I	J	K	L
1		1982	1984	1988	1990	1992	1994	1996	1998	2000	2002	2004
2	1980	25.60	25.30									
3	1981	25.32	24.96									
4	1982	25.04	24.62									
5	1983	24.76	24.28									
6	1984	24.48	23.94									
7	1985	24.20	23.60	23.60								
8	1986	23.92	23.38	23.32								
9	1987	23.64	23.16	23.04								
10	1988	23.36	22.94	22.76								
11	1989	23.08	22.72	22.48								
12	1990	22.80	22.50	22.20	22.10	21.90	21.90					
13	1991	22.72	22.44	22.08	21.98	21.86	21.84					

Note: For brevity, only a subset of revision years (columns of this table) and a subset of predicted year (rows) are illustrated.

Figure A3 shows the predictions for Australia's 0- to 14-year-old population ratio at five-year intervals for each update year, calculated by employing linear interpolation using the above method. The vertical axis is the proportion (%), while the variously colored lines show the differences by revision year.

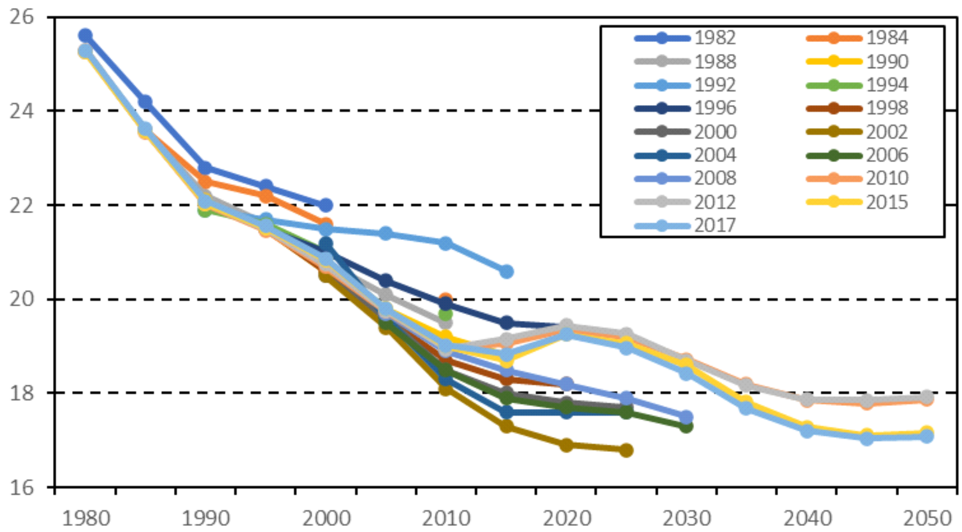


Figure A3 Predictions for Australia's 0- to-14-Year-Old Population Ratio in Five-Year Intervals by Update Year

[3] For non-revision years

Predictions for non-revision years are obtained by performing linear interpolation using the predictions from the previous and next revision years. The five-year prediction for 1988 as of 1983 is obtained using the following method. First, the prediction for 1988 based on the 1982 revision is calculated as follows: $24.2+(22.8-24.2)/5\times 3 = 23.36$. Similarly, the prediction for

1988 based on the 1984 revision is obtained thus: $23.6+(22.5-23.6)/5\times 3 = 22.94$. Next, the five-year prediction as of 1983 is obtained by performing linear interpolation of these values for 1988: $23.15 = (23.36 + 22.94)/2$.

Table A6 Predictions for Non-Revision Years Using Linear Interpolation

	A	B	C	D	E	F	G	H	I	J	K	L
1		1982	1983	1984	1985	1986	1987	1988	1989	1990	1991	1992
2	1980	25.60	25.45	25.30								
3	1981	25.32	25.14	24.96								
4	1982	25.04	24.83	24.62								
5	1983	24.76	24.52	24.28								
6	1984	24.48	24.21	23.94								
7	1985	24.20	23.90	23.60	23.60	23.60	23.60					
8	1986	23.92	23.65	23.38	23.37	23.35	23.34					
9	1987	23.64	23.40	23.16	23.13	23.10	23.07					
10	1988	23.36	23.15	22.94	22.90	22.85	22.81					
11	1989	23.08	22.90	22.72	22.66	22.60	22.54					
12	1990	22.80	22.65	22.50	22.43	22.35	22.28	22.20	22.15	22.10	22.00	21.90
13	1991	22.72	22.58	22.44	22.35	22.26	22.17	22.08	22.03	21.98	21.92	21.86

Note: For brevity, only a subset of years (columns of this table) and a subset of predicted years (rows) are illustrated.

Using these methods, five-year projections are calculated for each year. The projection value trends for the 65-and-over ratio for each update year obtained using a similar approach are shown below.

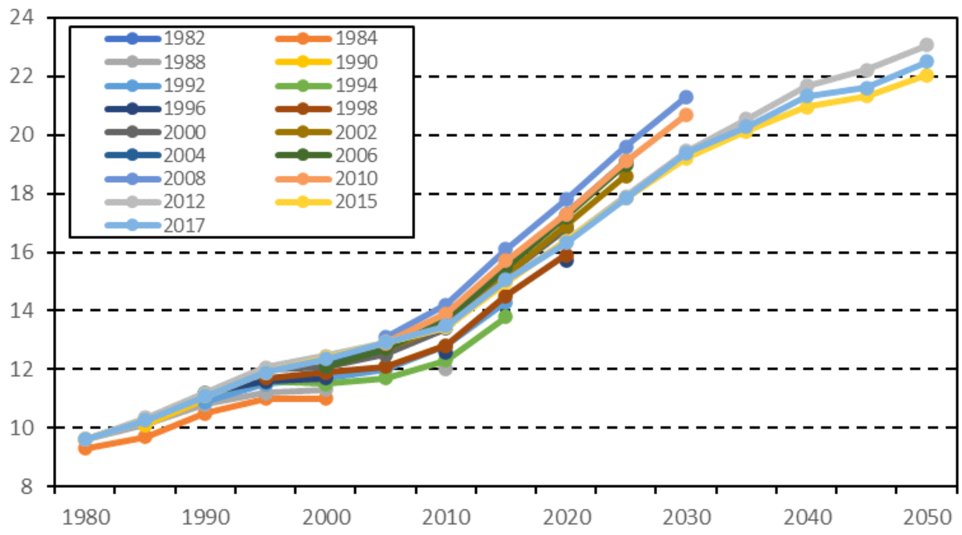


Figure A4 Predictions for Australia's 65-and-Over Population Ratio in Five-Year Intervals by Update Year

References

[1] Badarinza, C., Ramadorai, T., Shimizu, C., 2021. Gravity, Counterparties, and Foreign Investment, J. Financ. Econ, 145 (2), 532-555. Doi: http://dx.doi.org/10.1016/j.jfineco.2021.09.011.

[2] Campbell, J.Y., Shiller, R.L., 1988. The dividend-price ratio and expectations of future dividends and discount factor. Rev. Financ. Stud. 1 (3), 195–228. Doi:10.1093/rfs/1.3.195.

[3] Cavalcanti, T.V. de V, Mohaddes, K., Raissi, M., 2011. Growth, development and natural resources: new evidence using a heterogeneous panel analysis. Q. Rev. Econ. Finance 51 (4), 305–318. Doi:10.1016/j.qref.2011.07.007.

[4] Claessens, S., Kose, A., Terrones, M.E., 2011. Financial cycles: what? how? when?" NBER International Seminar on Macroeconomics 7 (1), paper WP-11-76. Doi:10.1086/658308.

[5] Crowe, C., Dell'Ariccia, G., Igan, D., Rabanal P., 2013. How to deal with real estate booms: lessons from country experiences. J. Financ. Stab. 9 (3), 300–319. Doi:10.1016/j.jfs.2013.05.003

[6] Diewert, E., Nomura, K., Shimizu., 2023. Improving the SNA: Alternative Measures of Output, Input, Income and Productivity for China. TCER Working Paper Series, E-178.

[7] Diewert, E., Nishimura, K., Shimizu, C., Watanabe, T., 2020. Property Price Indexes. Springer, Dordrecht.

[8] Eggertsson, G.B., Mehrotra, N.R., Robbins, J.A. 2019. A Model of Secular Stagnation, Am. Econ. Rev, Macroeconomics, 11(1), 1-48. Doi: 10.1257/mac.20170367.

[9] Engelhardt, G.V., Poterba, J.M., 1991. House prices and demographic change. Reg. Sci. Urban Econ. 21 (4), 539–546. Doi:10.1016/0166-0462(91)90017-H.

[10] Fair, R.C., Dominguez, K.M., 1991. Effects of the changing U.S. Age distribution on macroeconomic equations. Am. Econ. Rev. 81 (5), 1276-1294. Doi. https://www.jstor.org/stable/2006917.

[11] Gyourko, J., Mayer, C., Sinai, T., 2013. Superstar Cities. AEJ: Economic Policy. 5 (4), 67-199. Doi: 10.1257/pol.5.4.167.

[12] Hamilton, B. W., 1991. The baby boom, the baby bust, and the housing market: a second look. Reg. Sci. Urban Econ. 21 (4), 547–552. Doi:10.1016/0166-0462(91)90018-I.

[13] Hendershott, P.H., 1991. Are real house prices likely to decline by 47 percent? Reg. Sci. Urban Econ. 21 (4), 553–563. Doi:10.1016/0166-0462(91)90019-J.

[14] Hirano, T., Toda, A.A., 2023a. Unbalanced Growth, Elasticity of Substitution, and Land Overvaluation. Doi: https://doi.org/10.48550/arXiv.2307.00349

[15] Hirano, T., Toda, A.A., 2023b. Bubble Economics. Doi: https://doi.org/10.48550/arXiv.2311.03638

[16] Hirano, T., Yanagawa, N., 2017. Asset bubbles, endogenous growth, and financial frictions. Rev. Econ. Stud. 84 (1), 406–443. Doi:10.1093/restud/rdw059.

[17] Im, K.S., Pesaran, M.H., Shin, Y., 2003. Testing for unit roots in heterogeneous panels. J. Econom. 115 (1), 53–74. Doi:10.1016/S0304-4076(03)00092-7.

[18] Ito, T., Hoshi, T., 2020. The Japanese Economy, 2nd ed. The MIT Press, Cambridge, Massachusetts.

[19] Jord'a, 'O., Schularick,M., Taylor, A.M., 2015. Leveraged Bubbles. J. Monet. Econ. 76, S1–S20. Doi: 10.1016/j.jmoneco.2015.08.005.

[20] Kao, C., 1999. Spurious regression and residual-based tests for cointegration in panel

data. J. Econom. 90 (1), 1–44. Doi:10.1016/S0304-4076(98)00023-2.
[21] Kindleberger, C. P. (2000). *Manias, Panics, and Crashes*. 4th ed. New York: John Wiley & Sons.
[22] Kiyotaki, N., Moore, J., 1997. Credit cycles. J. Pol. Econ. 105 (2), 211–248. Doi:10.1086/262072.
[23] Knoll, K., Schularick, M., Steger, T. (2017). No price like home: Global house prices, 1870–2012. Am. Econ. Rev, 107(2), 331–353. Doi: 10.1257/aer.20150501
[24] Lutkepohl, H., 1993. Introduction to Multiple Time Series Analysis, 2nd ed. Springer-Verlag, Berlin.
[25] Mankiw, N.G., Weil, D.N., 1989. The baby boom, the baby bust, and the housing market. Reg. Sci. Urban Econ. 19 (2), 235–258. Doi:10.1016/0166-0462(89)90005-7.
[26] Minsky, H.P., 1992. The financial instability hypothesis. Working paper 74. Levy Economics Institute, Bard College.
[27] Nakagawa, M., Shimizu, C., 2023, Aging city and house prices: Impact of Aging Condominium Stock on the Housing Market in the Tokyo-, Int. Real Estate Rev, forthcoming.
[28] National Institute of Population and Social Security Research, (1976, 1986, 1992, 1997, 2002, 2006, 2012, 2017) Population Projections for Japan.
[29] Nishimura, K.G., 2016. Three "seismic shifts" in the global economy and the policy challenges they pose. Int. Finance 19 (2), 219–229. Doi:10.1111/infi.12089.
[30] Nishimura, K.G., Takáts, E., 2012. Ageing, property prices and money demand. BIS working papers, No. 385.
[31] Ohtake, F., Shintani, M., 1996. The effect of demographics on the Japanese housing market. Reg. Sci. Urban Econ. 26 (2), 189–201. Doi:10.1016/0166-0462(95)02113-2.
[32] Pedroni, P., 1999. Critical values for cointegration tests in heterogeneous panels with multiple regressors. Oxf. Bull. Econ. Stat. 61 (s1), 653–670. Doi:10.1111/1468-0084.61.s1.14.
[33] Pedroni, P., 2004. Panel cointegration: asymptotic and finite sample properties of pooled time series tests with an application to the PPP hypothesis. Econ. Theory 20 (3), 597–625. Doi:10.1017/S0266466604203073.
[34] Pedroni, P., 2007. Social capital, barriers to production and capital shares: implications for the importance of parameter heterogeneity from a nonstationary panel approach. J. Appl. Econom. 22 (2), 429–451. Doi:10.1002/jae.948.
[35] Pesaran, M.H., Shin, Y., Smith, R.P., 1999. Pooled mean group estimation of dynamic heterogeneous panels. J. Am. Stat. Assoc. 94 (446), 621–634. Doi:10.1080/01621459.1999.10474156.
[36] Reinhart, C.M., Rogoff, K.S., 2009. This Time Is Different: Eight Centuries of Financial Folly. Princeton University Press, Princeton, NJ.
[37] Saita, Y., Shimizu, C., Watanabe, T., 2016. Aging and real estate prices: evidence from Japanese and US Regional Data. Int. J. Hous. Mark. Anal. 9 (1), 66–87. Doi:10.1108/IJHMA-11-2014-0053.
[38] Shimizu, C., Watanabe T., 2010. Housing bubble in Japan and the United States. Pub. Pol. Rev. 6 (3), 431–472.
[39] Takáts, E., 2012. Aging and house prices. J. Hous. Econ. 21 (2), 131–141. Doi:10.1016/j.jhe.2012.04.001.
[40] Tamai, Y., Shimizu, C., Nishimura, K.G., 2017. Aging and property prices: a theory of very-long-run portfolio choice and its predictions on Japanese municipalities in the 2040s. Asian Econ. Pap. 16 (3), 48–74. Doi:10.1162/asep_a_00548.
[41] Taylor, J.B., 1993. Discretion versus policy rules in practice. Carnegie-Rochester Conf. Ser. Public Policy 39, 195–214.

[42] Walras, L., 1954. Elements of Pure Economics, translated by W. Jaffe (first published in 1874). George Allen & Unwin, London.