Demographics, Property Prices, and Credit Conditions: Analysis Based on Panel Data from 17 Countries Over a Half-Century*

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Abstract

Using panel data from 17 countries with varying economic circumstances from 1974 to 2019, we estimate regression models that explain residential property price dynamics by incorporating demographic factors and considering the interaction of those demographics with credit conditions. Our results show the importance of the 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 the persistently optimistic population projections lead to the oversupply of the residential stock in rapidly aging countries, resulting in stagnant residential property markets.

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1 Population Dynamics, Property Prices, and the Effects of Nominal Interest Rates

The world economy is facing the unprecedented challenge of changing population dynamics, with some countries experiencing a rapid rise of the young population while others are facing rapid aging. In such uncharted territory, it is crucial to ask: Is the effect of the change in nominal interest rates (determined by monetary policy) similar across countries or vastly different? Why is the recovery from collapse of the so-called property bubbles inconsistent across countries? Casual observation shows that property prices are not recovering to the pre-bubble era in rapidly aging economies as fast as in other economies, and their adjustment has taken very long (Crowe et al., 2013[8]). However, existing research has not provided sufficient answers to such questions. We attempt to answer these questions from an econometric approach through the experience of 17 divergent economies over 46 years, focusing on population dynamics statistics.

During the post-World War II rapid economic growth era, Japan experienced the largest property bubble of the 20th century (Shimizu and Watanabe, 2010[?]). However, following the collapse of the bubble in 1990, the Japanese economy faced an extended period of economic stagnation termed the "lost decade." At the peak of the Japanese bubble, land in central Tokyo was sold for as much as 50 million yen (about US $500,000) per square meter; however, property prices plummeted after the bubble's collapse, especially in regional markets. In the 21st century, Japan's population was in a period of both fast aging (the fastest rate in the world) and absolute decline. In light of this, the number of vacant houses has continued to increase to depress regional property markets, and it has been predicted that in 10 years, 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 the 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 the population composition reveals that this insolvency occurred when the old age dependency ratio, which indicates the proportion of the old population (ages 65 and over) to the working-age population (ages 15 to 65), was at a level exceeding 90%. In the context of these phenomena, we explore what kind of economic mechanisms—including demographics—are at work, focusing on their property markets.[*1]

From the 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 2007) faced the formation of property bubbles and then long-term economic stagnation following the collapse of those bubbles. According to Claessens et al. (2011)[5], 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

[*1] In recent years, there are 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 Asia Pacific, where residential prices have dramatically increased. However, there is substantial regional heterogeneity, making price behavior very different between these superstar cities and the national average. This study is about the national average of residential property prices, not that of these superstar cities.

and scope (Crowe et al., 2011[7]).

Even before the global financial crisis triggered by the collapse of Lehman Brothers, numerous studies have attempted to properly understand the correlation between long-term overall economic stagnation and the mechanisms underlying large fluctuations in asset prices, such as property bubbles. Reinhart and Rogoff (2009)[29] attempt to understand the relationship between the two based on long-term economic data from multiple countries, covering over 100 years. Reinhart and Rogoff (2009)[29] elucidate four common phenomena observed in countries that 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 debts. Furthermore, the study clarifies that when society as a whole is excessively optimistic, it leads to high growth 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.

There are several theoretical frameworks to explain Reinhart and Rogoff's (2009)[29] findings. For example, Kiyotaki and Moore (1997)[19] 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 makes economic conditions deteriorate sharply to induce accelerated de-leveraging. During the de-leveraging process, many assets, including property, are sometimes on "fire-sale," causing lasting damage to property markets.

In comparison, let us consider the Japanese experience. The post-war era of rapid economic growth was driven by the generation born during the postwar baby boom reaching the working age. This period is known as a "population bonus phase" (Ito and Hoshi, 2020[17]). Then, in the early 1980s, this baby-boomer generation became home buyers 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. Following the collapse of the bubble in 1990, Japan's working-age population has 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 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. Considering these facts, it follows logically that there is a close relationship between population factors, large property market fluctuations, and economic downturns.

The literature review suggests that demographics and the property market have a strong underlying influence on 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 46 years.

(1) Did changes in population composition influence the dynamics of residential property prices?
(2) Did changes in population composition amplify/dampen nominal interest rates' effects 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; and 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 involved in the present value relationship (PVR) model (Campbell and Shiller, 1988[3]). Using international panel data from 17 countries with diverse population compositions, population trends, economic growth rates, and housing market environments, over almost half a century, this study empirically examines the relationship between demographics, property price dynamics, and credit cycles. In the previous research, only limited residential property price data are obtained for 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 for this kind of analysis of slow-moving long-term factors.[*2]

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)[22] central criticism of demographics and housing market-related analysis focuses on the fact that if economic agents' expectations are rational (that is, 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 supply is sufficiently elastic. However, we find some 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 underestimation of demad persists, housing supply shortage accumulates over time to cause an increase in residential property prices. In contrast, when population is aging rapidly and overestimation of demand persists, housing supply surplus becomes persistent to depress residential property prices. Therefore, to assess the effects of possible persistent demographic expectation errors, we collect data by tracing 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 the ones projected before.

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

2 The Empirical Model

Mankiw and Weil (1989)[22] examine the relationship between demographics and the property market in their seminal work. They argue that in the 1980s, housing demand peaked due to the baby-boomer generation in the United States; subsequently, over the next 20 years, till 2007, real residential property prices decreased by 47% due to population decline. This so-called "asset meltdown hypothesis" subsequently caused a considerable debate (Mankiw and Weil, 1989[22]).

Two major issues were central to this debate: supply elasticity and accuracy of projected demographic changes. Researchers argued that demographic changes take place at an extremely slow pace, and, thus, they are accurately predictable. Therefore, if the housing supply is elastic, even in the event of a pessimistic future population projection, no residential property price slump should occur, since supply will be adjusted via stock adjustment (Hendershott,

[*2] After the financial crisis, the International Monetary Fund (IMF) and Bank of International Settlements (BIS) took the lead in developing the property price index internationally. BIS and the Organisation for Economic Co-operation and Development (OECD) began publishing the property price index in 2016 (see Diewert et al., 2020)[9].

1991[14]; Hamilton, 1991[13]). Engelhardt and Poterba (1991)[10]'s 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)[**?**], 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. Furthermore, Shimizu and Watanabe (2010)[**?**] estimate housing demand using Mankiw and Weil (1989)[22]'s framework and expand the model based on panel data (by prefecture in Japan and by state in the United States) and 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. Particularly, land, which is one of the essential factors 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.[*3] If this is the case, even when we assume that demographic changes are perfectly predictable, demographics may influence residential asset prices. In this context, Takáts (2012)[32] is using a two-generation overlapping generation model, demonstrate that increases in the working-age population lead to rises in real residential property price 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 our study's subject.[*4]

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 expectation (projection) errors are not zero on average over a given period. 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 the year 2010, made in 1992, was approximately 1.8. However, the actual TFR in the year 2010 was 1.4. This gap (0.4) is likely to lead to over-capacity and over-supply. This suggests another important

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[*3] Gyourko, Mayer and Sinai (2013)[12] focus their attention that most pronounced residential price movement is driven by the limited supply of land. Similarly, Knoll, Schularick and Steger (2017)[20] shows that residential price movement over a century is mostly brought on by the movement of the land price, and the sharp increase in the last half of the 20th century in particular is caused by the substantial appreciation of the land.

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

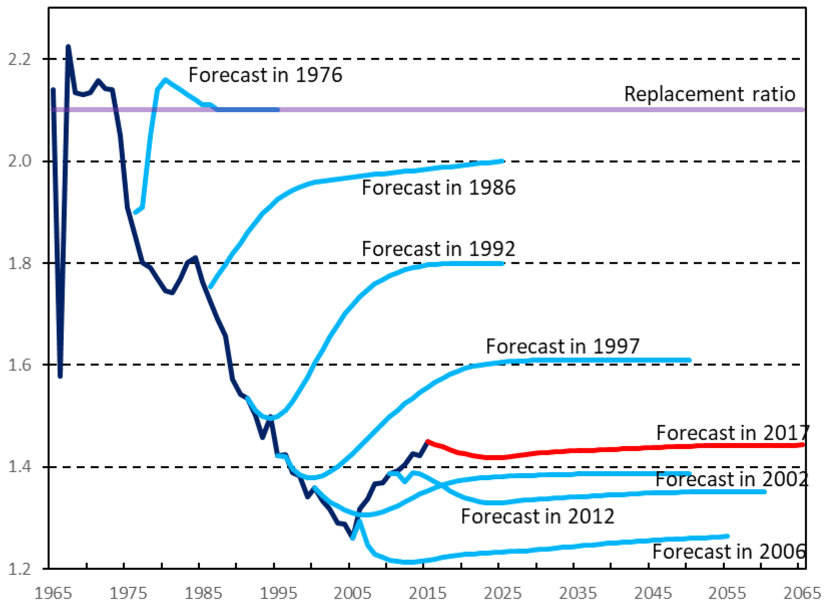

Figure 1 Persistently Optimistic Predictions of the Total Fertility Rate by Government Experts at the Unfavorable Time of Declining Rates (Japan)
Source: National Institute of Population and Social Security Research. Various issues.

route via 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 expectations errors. Thus, if government experts' and regular people's overly optimistic forecast errors persist and are accumulated, residential property prices decline more than in the contrasting scenario.

This study empirically investigates the demographic dynamics' effects on property prices considering the above-mentioned two points. Specifically, we elucidate that long-term nominal residential property prices are determined by the perfect foresight PVR.[*5] We start with the following long-run equilibrium relation between 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):[*6]

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

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

[*5] We perform a robustness check based on the PVR using real residential property price, real rent, and real interest, but there are no major changes in the basic results. As summarized in Cochrane (2011), recent studies on asset pricing have acknowledged that the time-varying discount rate is a major determinant.In this study, the robustness was checked using both perfect foresight and static expectation as the real interest rate, and the present value relationship was confirmed to be valid before proceeding with the analysis. For details, see Appendix 1.

[*6] In addition to the nominal interest rate there is a nominal risk premium in the denominator since the residential property is a risky asset. Thus even if the nominal interest rate becomes negative, the denominator is still positive and PVR is well-defined. For simplicity, we assume the nominal risk premium does not change over time and thus it becomes a constant in Equation (2). For further model details, see Campbell and Shiller (1988)[3].

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 rent data is a common problem for empirical analyses based on Equation (1). In this study, for a consistent definition across 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 25%–30% in each country. 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.*7

Meanwhile, as mentioned previously, Takáts (2012)[32] derive the relationship between working-age population increases and rises in real residential property price using a two-generation intergenerational model. Furthermore, empirical studies such as Takáts (2012)[32] and Saita et al. (2016)[30] indicate that population composition has a significant impact on residential property prices. Based on the previous research's findings, we specify the expected change rate of housing rents as a function of population factors, namely the population compositions and the total population. As we have clarified, population variables may not be accurately estimated, and their errors may persist and accumulate over time (that is, they may not be "rational"). To incorporate these possibilities in the perfect foresight framework of Equation (1), we assume perfect foresight of demographic variables' expectation errors and do not impose zero-sum constraints.

Equation (2) is a benchmark specification derived by performing a logarithmic transformation on both sides of the long-term PVR established based on the above:*8*9

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

where P^{rppi} denotes the residential property price index, P^{cpi} is the consumer price index, Y is the real GDP, i is the nominal interest rate, pop^{total} is the total number of population, and n^{yng}, n^{wrk}, and n^{old} are the shares of the total population that belongs to three generations, young (ages 0-14), working-age (15-64), and old (65+), respectively. The subscript j represents the country, and t represents the time period. ε_{jt} denotes the disturbance term. Economic theory expects the following restrictions in the coefficients: $\beta_1 = 1$ (absence of money illusion), $\beta_2 > 0$ (an increase in housing rents raises residential property prices), and $\beta_3 < 0$ (an increase in the nominal rate of interest lowers residential property prices), as will be verified below.

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

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

*9 To understand population dynamics' impact on fundamentals, we employ the following specification, based on Takáts (2012), Saita et al. (2016)[30], and Tamai et al. (2017)[34].

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

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 a time point t is always $1, \delta_1, \delta_2$, 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[11]). That is, the formula is modified as follows:

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

In Takáts (2012)[32] 's two-generation model, the main buyers of property are the younger generation; therefore, it is predicted that increases in the younger generation's population will result in higher real residential property prices ($\beta_5 > 0$). Conversely, it is expected that increases in the older generation's population will 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 other variables' coefficients are homogenous, 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 will also be tested as part of the empirical research.

As an extension of the basic model, we perform two types of analysis. First, for the property bubble and collapse periods, we empirically analyze what kind of effect population composition has on the impact of interest rates on property prices. Specifically, we add the population factor and interest rate interaction terms to the estimation model.*10

A decline in nominal interest rates has large positive impacts on property prices when population is young and growing. In contrast, the experiences of Japan, the United States, and Ireland, following property booms, show that the effect of monetary easing measures, such as lowering of 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 nominal interest rate and population ratio interaction terms to 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 will examine Equation (3)'s estimated coefficient of i incorporating the interaction term to see 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 the residential property price in the long-run equilibrium. Property is a durable good, and so it is difficult for supply to adjust 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 there will be 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 results of the benchmark case Equation (2) and the extention case 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 number. A similar procedure is applied to the old population

*10 As Bielecki et al (2018) point out, demographics itself can also be a pathway through which monetary policy is affected. In this paper, the relationship between demographics and monetary policy is considered as a cross effect of the discount rate.

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 makes it possible to decompose and analyze the effects of predictable and unpredictable parts of population ratios on residential property price.

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 the BIS, with participation from the United Nations and the OECD.*11 The 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 Asia-Pacific (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 I). 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 BIS for P^{rppi}. As this index is published quarterly, we use the simple average for each year. For nominal interest rates, i, the main source is long-term interest rates from OECD.Stat. However, as data for Denmark, Italy, Japan, Norway, and Sweden are missing for part of the study period, percent per annum data on interest rates, government securities, and government bonds obtained from the IFS are used as a substitute. We use values obtained by converting these nominal interest rates (annual rates) into continuous compounded interest rates in the regression analysis. For real GDP, Y, we use the real GDP (local currency unit) published in the World Bank's World Development Indicators (WDI). The CPI, P^{cpi}, is likewise obtained from the WDI.

*11 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. Based on 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 internationally comparable property price indexes, see Diewert et al. (2020)[9].

Table I 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)

For 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 (ages 0–14), working-age (15–64), and old (65+) for each country, and calculate the population ratios with respect to the total population, n^{yng}, n^{wrk}, and n^{old}. Total population data are also used as an explanatory variable, pop^{total}.

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. To capture housing investment's characteristics (time lag between planning and construction start/completion), we select a five-year prediction interval for this study. In the analysis of Equations (2) and (3), we use panel data for 17 countries covering 46 years 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.[*12] Accordingly, the analysis of Equation (5) is based on a 31-year panel for 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[16]). The test is performed based on the following regression model:

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

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

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

[*12] 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, 2015, etc.). By using this data, it is possible to calculate approximate predictions for each year based on linear interpolation. 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 A-3.

Table II 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.

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

Here, we summarize the test results (see Table II). The IPS and ADF–Fisher tests reach identical conclusions. Specifically, the test results show that the residential property price, working-age per-capita real GDP, nominal interest rate, and total population are integrated of order I(1). In contrast, 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.[*13] 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. Due to 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 panel cointegration relation testing and estimation is as follows. First, we perform a panel cointegration test based on residuals by checking the stationarity of 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[18]; Pedroni, 1999[25], 2004[26]). The Kao test assumes that all cointegration vectors are common to each country. We assume the commonality of the dis-

*13 As the population ratio variables have values that are restricted to interval $[0, 1]$, by definition, they are stationary processes. However, with panel unit root tests such as in this case, it is sometimes not possible to reject local non-stationarity. 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)[27] and Cavalcanti et al. (2011)[4] were also able to estimate panel cointegration relations including ratio variables (specifically, investment-income ratios) since they demonstrated local non-stationary processes.

counted PVR (housing market no-arbitrage condition) across all countries; therefore, the Kao test's assumption is not impossible. However, since the commonality assumption is typically an extremely strong hypothesis, there is also a possibility that the cointegration vectors will be non-homogenous. In that case, it may be considered inappropriate to apply the Kao test. In comparison, the Pedroni test may be considered more generally representative than the Kao test, since it permits cointegration vectors that vary by country. 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).*14

Next, we estimate the long-run relationship based on the panel ARDL approach (Pesaran et al., 1999[28]):

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

Equation (7) is a typical ARDL(p, q) model, where p is the maximum lag length of dependent variable y_{it}, and q is the maximum lag length of a vector of explanatory variables $X_{i,t}$. It is technically possible to set different lag orders for each country, but, for simplicity, we have chosen to use a common order.

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 the third terms capture the short-run adjustment processes. The parameters in Equations (7) and (8) are associated as:$\theta_i = -(1 - \sum_{j=1}^{p} \lambda_{ij}), \beta_i = \sum_{j=0}^{q} \delta_{ij}/(1 - \sum_{k=0}^{q} \lambda_{ik}), \lambda^*_{ij} = -\sum_{m=j+1}^{p} \lambda_{im}$, and $\delta^*_{ij} = -\sum_{m=j+1}^{q} \delta_{im}$. Note that all coefficients vary by country. For example, the coefficient θ_i in Equation (8) represents the speed of adjustment of equilibrium errors, and this speed of adjustment varies by country. If a long-run equilibrium relationship exists between the variables, the sign for 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 also 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, coefficients for variables in difference expressing short-run adjustment vary by country. It may easily be expected that if housing market structure/policy and consumer preferences vary, the short-run adjustment process and correction of equilibrium errors will also vary by country.

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

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

*14 As pointed out by Cochrane (2011), there is a problem that the present value relationship as shown in equation (1) does not hold in the determination of asset prices. However, in the case of the determination of residential prices, the results of the panel cointegration quantile test indicate that a long-run relationship cannot be said not to exist between variables that indicate a present value relationship. The following discussion is based on the present value relationship.

Since Equation (9) is a non-linear model in terms of the parameter imposing the restriction of homogeneity on the cointegration coefficients ($\beta_i' = \beta'$), maximum likelihood estimation is used (Pesaran et al., 1999[28]). This restriction may, at first glance, seem strong. However, for the 17 countries covered in the analysis, we believe that we can assume that the financial conditions are homogenous 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 be. 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), the two panel cointegration tests, and various panel ARDL specifications. We also report the empirical results based on the interaction between demographic composition and interest rate and its effect on the residential property price (Equation 3) and the impact of persistent demographic expectation errors on the residential property price (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 III). Fur-

Table III 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.

thermore, in the results using 7 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 III). The group ADF test also rejects the null hypothesis at the 1% level.*15 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 IV shows the PMG model and MG model estimates of the long-run parameters using panel data from 1974 to 2019 for the 17 countries.*16 At a glance, the coefficients of the MG model are largely insignificant, whereas the coefficients of the PMG model are significant with expected signs. As shown in the bottom of Table V, the Hausman test statistic is 10.53, and its p value is 0.1042.*17 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. Based on these results, we present the PMG model results below.

Table IV Estimation Results of Baseline Model, Equation (2)

	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 (2). 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.

The analyses of the PMG model estimation results in Table IV are as follows. First, the CPI coefficient is approximately 1, suggesting the possibility that money illusion does not exist. Second, the coefficient for the working-age per-capita real GDP is 0.410. This is a housing rent proxy variable, and the fact that this coefficient is positive is consistent with the discounted present value theory. Third, the estimated value of the impact of the nominal interest rate is

*15 According to Pedroni (2004)[26], 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.

*16 The ARDL model's lag order was taken as $p = 2, q = 1$, based on Schwarz information criteria (SIC).

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

−8.705, which is also significant at the 1% level. This result is likewise consistent with the theoretical prediction based on the present value model wherein interest rate increases will hurt asset prices. Finally, and most importantly, with regard to population ratio coefficients, the coefficient for the young population ratio is 5.579, whereas the coefficient for the old population coefficient is −5.705. These coefficients are statistically significant at the 1% level. Thus, if 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 as shown in Table V. The interest rate coefficient is insignificant, but the interest rate and old population ratio interaction term coefficient is significant.

Due to the interaction terms in Equation (3), the impact of interest rate cuts on residential property prices depends on the 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 change on residential property price of country j over time such as

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

Table V 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.

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

$$\frac{\partial \log P_{jt}^{rppi}}{\partial i_{jt}} = \frac{\partial \log P_j^{rppi}}{\partial i_j} + \beta_7 \widetilde{n_{jt}^{yng}} - (\beta_7 + \beta_8)\widetilde{n_{jt}^{wrk}} + \beta_8 \widetilde{n_{jt}^{old}}, \tag{11}$$

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 V, Equations (10) is re-written as follows:

$$\frac{\partial \widehat{\log P_j^{rppi}}}{\partial i_j} = 3.949 - 8.526\, \overline{n_j^{yng}} - 19.765\, \overline{n_j^{wrk}} + 28.291\, \overline{n_j^{old}}, \tag{12}$$

Table VI shows the average marginal effect of interest rate increases on property prices by country, calculated based on the population ratio average values obtained for the sample period (1974 to 2019) using Equation (12). 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 increases obtained by plugging in these values is the value at the bottom of Table VI, −6.738. This result signifies that a 1% interest rate cut raises property prices by around 6.738% on average, which is somewhat smaller in scope 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 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 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 based on historical averages.

The key takeaway is that divergence from historical averages further enhances the impact of the above monetary measures on the property market. Using the estimates reported in Table V, Equation (11) is re-written as:

$$\frac{\partial \widehat{\log P_{jt}^{rppi}}}{\partial i_{jt}} = \frac{\partial \widehat{\log P_j^{rppi}}}{\partial i_j} - 8.526 \widetilde{n_{jt}^{yng}} - 19.765 \widetilde{n_{jt}^{wrk}} + 28.291 \widetilde{n_{jt}^{old}}. \tag{13}$$

Equation (13) 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 will considerably strengthen the positive effect of interest rate cuts (monetary expansion), while conversely, population onuses will 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.

Table VI 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 (10). 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.

4.3 PERSISTENT DEMOGRAPHIC EXPECTATION ERRORS AND LONG-RUN EQUILIBRIUM

The estimation results of Equation (5) are summarized in Table VII.*18 There are changes in the estimation values that are statistically significant, such as the CPI coefficient (1.386), which is higher than the estimates in the previous sections, and the coefficient for 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 for the total population is negative. The shortened estimation interval may also have had an influence.

Equation (5) implies that marginal effect of interest rate change on residential property

*18 $p = 2, q = 1$ was selected as the optimum lag length for ARDL, based on SIC.

Table VII 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.

price of country j is:[*19]

$$\begin{aligned}\frac{\partial \log P^{rppi}_{jt}}{\partial i_{jt}} &= -1.886 - 33.013\,\breve{n}^{yng}_{jt} - 7.001\,\breve{n}^{wrk}_{jt} + 40.014\,\breve{n}^{old}_{jt} \\ &\quad - 413.758\,(n^{yng}_{jt} - \breve{n}^{yng}_{jt}) - 105.990\,(n^{wrk}_{jt} - \breve{n}^{wrk}_{jt}) \\ &\quad + 519.748\,(n^{old}_{jt} - \breve{n}^{old}_{jt}). \end{aligned} \tag{14}$$

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

The second term of Equation (14) represents the change in residential property prices with respect to the portion of the young population ratio increase that was 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 10% level. In addition, the fourth term suggests

[*19] Using Equation (5), the marginal effect of interest rate change on residential property price of country j is derived as:

$$\begin{aligned}\frac{\partial \log P^{rppi}_{jt}}{\partial i_{jt}} &= \beta_3 + \beta_9 \breve{n}^{yng}_{jt} - (\beta_9 + \beta_{10})\breve{n}^{wrk}_{jt} + \beta_{10}\breve{n}^{old}_{jt} + \beta_{11}(n^{yng}_{jt} - \breve{n}^{yng}_{jt}) - (\beta_{11} + \beta_{12})(n^{wrk}_{jt} - \breve{n}^{wrk}_{jt}) \\ &\quad + \beta_{12}(n^{old}_{jt} - \breve{n}^{old}_{jt}).\end{aligned}$$

By substituting the corresponsing estimates from Table VII, Equation (14) is obtained.

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

Furthermore, 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 VIII 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 (14). 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 the country, we empirically clarify the relationship between changes in demographics (including the aging of the population), nominal interest rates determined by monetary policy and residential property prices.*20 We also find the importance of demographic expectation formation: considerable difference exists between expected and unexpected change in demography.

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 relationship between residential property prices, demographics, and nominal interest rates. Also we find significant differences among countries in the short-run dynamics.*21 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. In terms of specific results, the study shows that if the ratio of the young population to the total population rises by 1%, residential property prices will increase by 5.579%, while, conversely, if the old population ratio increases by 1%, residential property prices will fall by 5.705%.

Furthermore, in Section 4.2, we estimate the relationship to credit conditions during property bubble and collapse periods by adding a cross-term interest rate and population factors to 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, due to 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. Furthermore, we determine that the effect of monetary easing measures, such as lowering of nominal interest rates, is severely restricted during population onus (aging) periods and in countries facing them.

As a criticism of various studies analyzing the relationship between demographics and the housing market, starting with Mankiw and Weil (1989)[22], it has been pointed out that since

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*20 We examine the case of real interest rates and real residential property prices in Appendix A-1, and find our resuts in Section 4 are robust.

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

populations change only slowly if perfect foresight exists with regard to population changes and aging in the long term, price falls should not occur due to stock being adjusted through supply adjustments (Engelhard and Poterba, 1991[10]; Hamilton, 1991[13]; Hendershott, 1991[14]). However, if changes in the 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 for prediction errors relating to the old population ratio in Equation (14) 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 based on the population ratio averages from 1989 to 2019, are presented in Table VIII.[*22] Several interesting suggestions may be obtained from this table.

First, in the case 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 cuts effects may be determined by predictable factors (the second and third terms in Equation (14)).

Second, in Belgium, Canada, Denmark, the Netherlands, and South Africa, the average old population ratio prediction errors are 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 predictions errors and decreases it in the other two countries with negative ones.

Third, as for the countries with positive old population ratio prediction errors on average, the marginal effects of interest rate are as follows: Germany (average prediction error: 0.11%, a 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 young population coefficient signs are negative, while both old population coefficient signs are positive, which is consistent with the series of estimation results and the prediction results. Notably, among the two young population coefficients, the prediction error coefficient is substantial. In comparison, interestingly, the prediction error coefficient is also very large among the two old population coefficients. 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 concerning monetary policy. According to the Taylor rule (Taylor, 1993[33]), when the economy is struggling (i.e., in times when the GDP gap is negative), monetary policy shores up the economy by lowering interest rates. When the economy is thriving (i.e., in times when the GDP gap is positive), excessive growth is

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

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 countires in Europe, Asia Pacific, Africa and Americas, 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 the proportion of the working-age population increased, resiendital property prices soared as housing demand increased. This tendency is strengthened further when credit conditions are loose with low nominal interest rates. Moreover, if demographic changes are unanticipated, then the credit conditions' effects become larger. In contrast, "population onus period", when the portion of old population is increased substantially, residential property markets are stagnated, and looseing of credit conditions do not have as strong positive effects as in the case of population bonus period.

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

However, some issues still prevail. First, in the current analysis, the definitions of working-age population and 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 down to some degree due to workers deferring retirement. We may incorporate retirement decision in defining working age population.

Second, our model is based on the assumption of relatively homogeneous population. However, population has become heterogeneous as immigration/emigration is increasingly important. As population becomes heterogeneous, its compositional effect may change over time. Such change will be incorporated in our future reseach.

Third, there is another kind of heterogeneity with respect to property markets. In fact, bipolarization in residential property markets is under way: some parts of urban areas (so-called superstar cities) experience rapidly rising residential property prices while the rest of a country suffers declining prices in regional property markets. Interestingly, some researchers argue that aging population has caused this bipolarization or bifurcation of national property markets. This is an important subject of future research.

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

Finally, this study is descriptive, and has not provided any suggestions concerning the issue of resource and welfare distribution. Previous research (Hirano and Yanagawa, 2017[15]) 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 have a linkage with demographics, which is an issue we would like to examine in the future.

Appendix:

A-1. Robustness Check: Analysis of Real Relationship

Theoretically, it is possible to interpret a discounted PVR as a long-run equilibrium relationship between real variables (Walras, 1954[**?**]). In this study, thus far, we have performed empirical analysis using nominal residential property price and interest rate values. 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 formation 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 consumer price index (CPI), while the explanatory variables are the working-age per-capita real gross domestic product (GDP), the real interest rate r, and population factors.[*23]

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

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

Real interest rate based on static expectations (SE): $r^{SE}_{jt} = i_{jt} - \Delta \log P^{cpi}_{j,t}$.

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

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

First, we consider the panel unit root tests' 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}. By Im–Pesaran–Shin (IPS) and Fisher–type augmented Dickey–Fuller (ADF–Fisher) tests, just as we did for the nominal variables, for real residential property prices and static expectation interest rates, the null hypothesis may be rejected at the 1% significance 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 homogenous, the Kao test rejects the null hypothesis that no cointegrating relation is present at the 1% level for both interest rate models (Table A2). Furthermore, with 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.

Finally, the PMG model's long-term coefficient estimation results are summarized in Table A3. The coefficient for 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 with either model, which is rather high compared with the value with the nominal

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

model (0.606). Next, although the real interest rate coefficient is negative and statistically significant for the static expectation case, it is positive and insignificant for the perfect foresight case. It reflects the difficulty of creating suitable real interest rate data based on annual data. Meanwhile, the results obtained with regard to population ratios are the same as for the nominal model.

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.

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.

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

The analysis thus far 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, it is difficult to instantly realize the fundamental value. Therefore, based on the PMG estimation results of Equation (9) in Section 4.1, we analyze the adjustment path until residential property prices reach long-run equilibrium when an exogenous shock occurs to residential property price fundamentals.

Using the estimates of common long-run parameters as well as the country-specific short-run paramters of the PMG model, and by successive substitutions of Equation (9), it is possible to express a 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 the residential property price level can be obtained by comparing two paths: one with a fundamental shock, and the other without a fundamental shock.*24 Here, we investigate the effects of two fundamental shocks: housing rent increases and nominal interest rate drops. Note that we are using a single equation model, variables other than the series giving the shock and the residential property price 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 the working-age per-

*24 For a similar analysis with a vector autoregressive (VAR) model including exogenous variables, see Lutkepohl (1993, p. 327)[21].

capita GDP increases by 1 unit. The horizontal axis is the number of years since the shock occurred, while the vertical axis is the residential property price increase rate. Based on the assumption of a long-run adjustment process, the values are illustrated over three decades.[*25]

The first characteristic is that the residential property price increase rate converges at around 0.4% over the long run. This is also the result expected based on 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 reached a peak was 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 the G7 countries, the increase rate was highest in the United Kingdom, followed in order by the United States, France, Italy, Canada, Germany, and Japan.[*26] Increases in working-age per-capita GDP directly produce housing demand and, therefore, the overshoot observed in the residential property price reaction 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 displayed divergent tendencies, is omitted from this graph. In terms of characteristics that may be observed from the graph, residential property prices rose 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 continued for at least four years. In the other six countries, residential property prices decreased at first immediately after the interest rate cut, but in four countries (the exceptions being Italy and Norway), they stopped decreasing after one year, while in Italy and Norway, they stopped decreasing after two years.

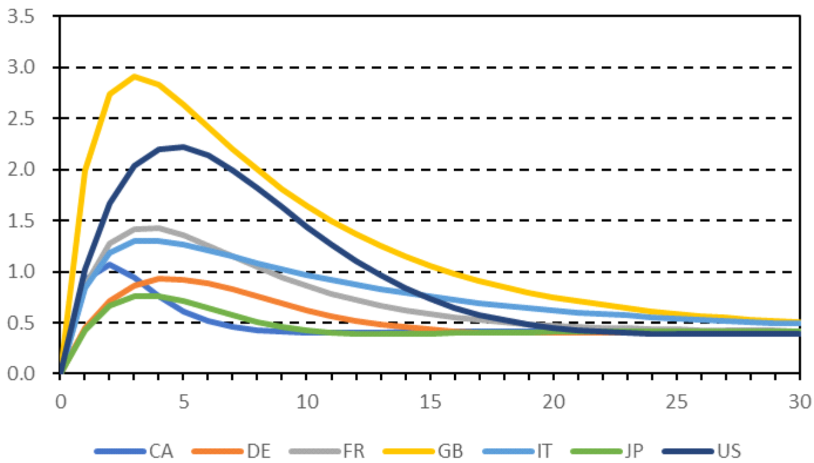
On examining G7 countries in detail, we observe that the reaction differs from housing rent shocks. One notable characteristic compared to housing rent is that adjustment of residential property prices in response to credit expansion requires a longer period. As the reduction of interest rates affects 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.

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

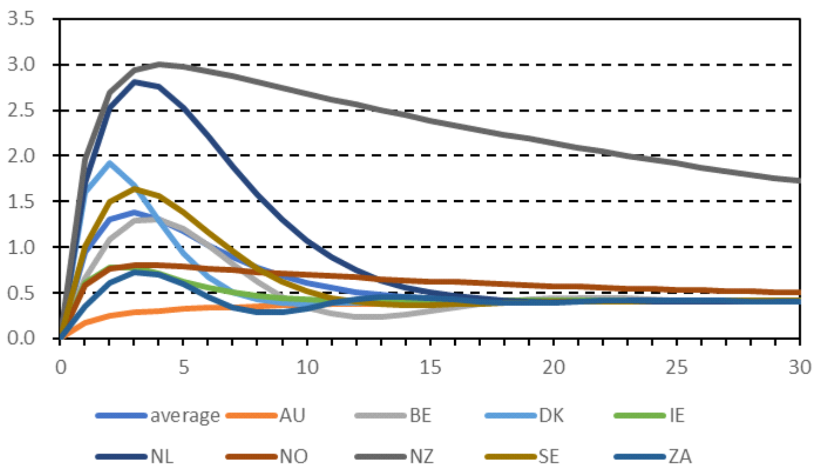
The United Nations publishes its revisions of the World Population Prospects every two years (with some exceptions). To date, updates have been made in 1982, 1984, 1988, 1990, 1992, 1994, 1996, 1998, 2000, 2002, 2004, 2006, 2008, 2010, and 2012. Each revision publishes population projections for years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

[*25] Among the 17 countries, Switzerland alone showed divergent behavior, and it was therefore excluded from the figure.

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



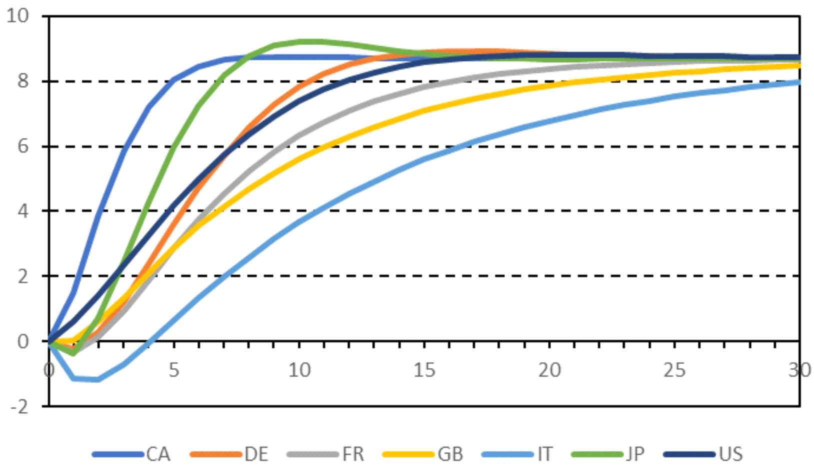
(a) G7 Countries



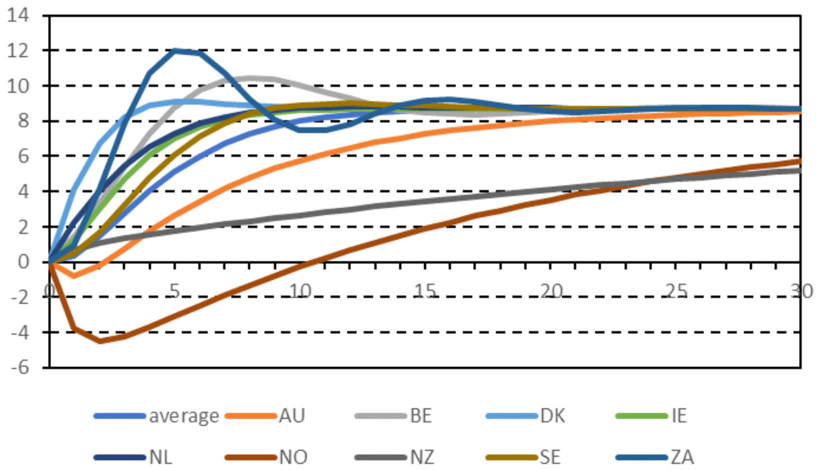
(b) 10 Non-G7 Countries (Excluding Switzerland)

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 I.



(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 I.

Column B is the predictions for 1985, 1990, 1995, and 2000 in the 1982 revision. More long-term predictions existed, but, for this study, data were 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.

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

[1] For revision years ending in 0 or 5

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.

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 year 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 based on 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 below 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 differences by revision year.

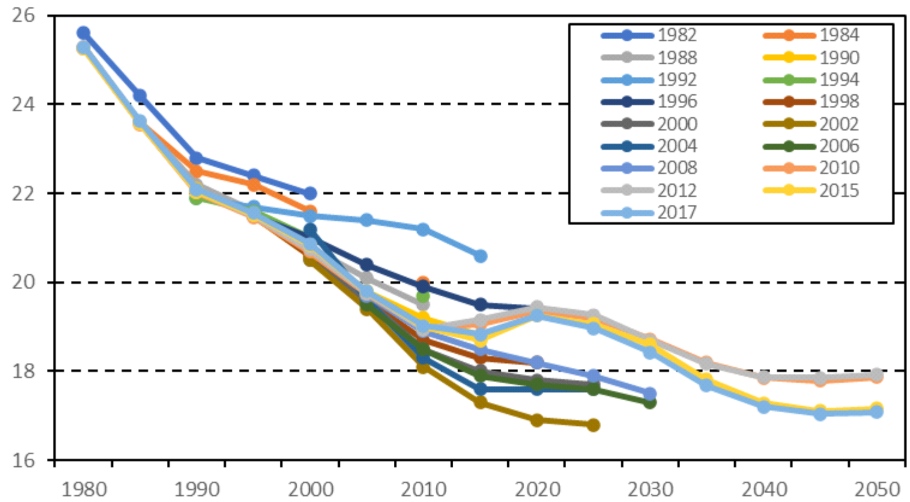


Figure A3 Predictions for Australia's 0- to-14-Year-Old Population Ratio in Five-Year Intervals by Update Year

[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 thusly: $23.6 + (22.5 - 23.6)/5 \times 3 = 22.94$. 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, fives-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.

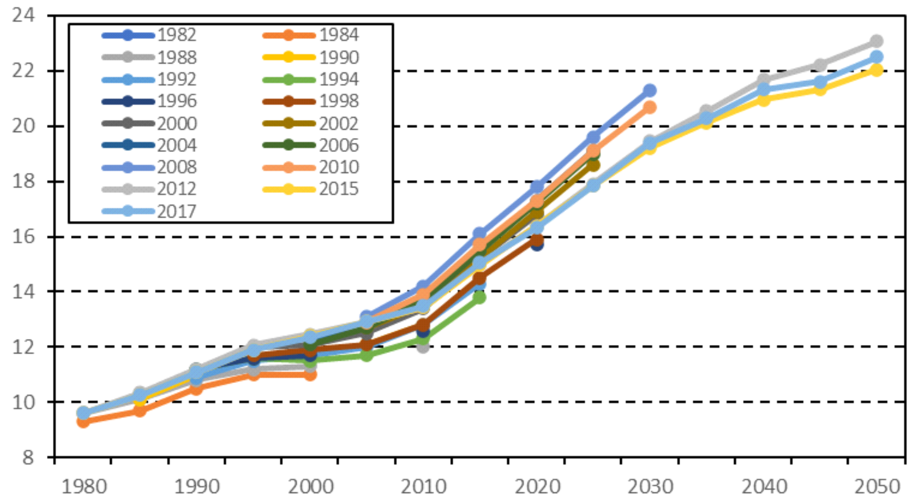


Figure A4 Predictions for Australia's 65-and-Over Population Ratio in Five-Year Intervals by Update Year

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