Renting Insures Wage Risk*

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Abstract

Homeowners and renters have mirror-image exposures to the considerable volatility in housing costs. The welfare effects of these exposures depend on their correlations with the rest of household portfolios. Using 70 years of data on local markets in the U.S., we find that rent and home price changes are strongly positively correlated with wage changes at all horizons. As a result, uninsured wage income risk is hedged by renting and exacerbated by owning. These interactions with wage income risk are strong enough that for many households, renting is not only safer than owning, it is safer than full housing insurance. This highlights an important cost of owner-occupied housing and the many major policies that encourage it.

JEL Classification: D12, D14, E21, G11, G51
Keywords: housing tenure, rent risk, home price risk, wage income risk, local labor markets

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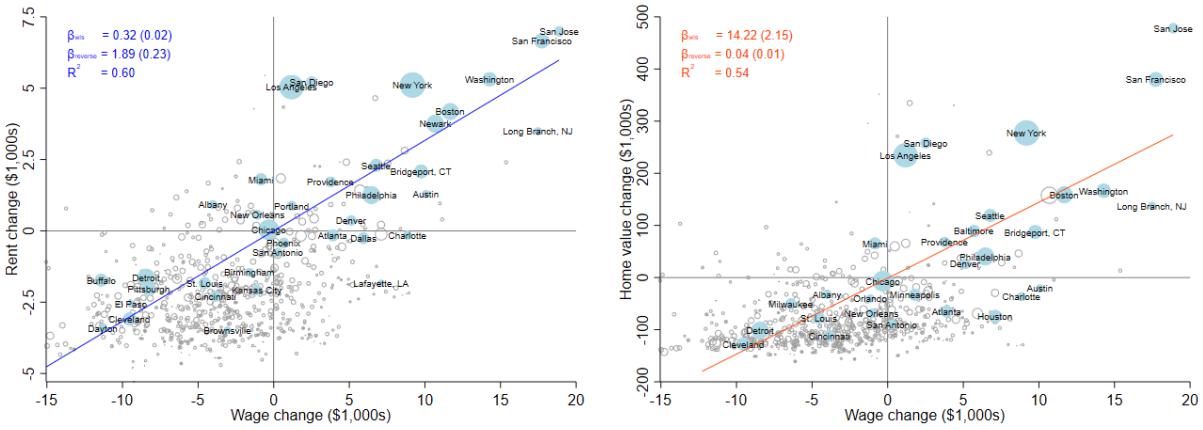
1 Introduction

Housing is a major consumption good and an important asset, comprising about one-fourth of household expenditures and about three-fifths of the median homeowner's assets (Consumer Expenditure Survey, 2020; Survey of Consumer Finances, 1989–2019). A household's exposure to the substantial volatility in rents and home prices ("housing prices") depends crucially on whether it owns or rents its home and for how long. Owning exposes a household to risk in the price at which it will sell its home but eliminates its exposure to risk in the rents it would face as a renter (Sinai and Souleles, 2005). Whereas a longer horizon in a home increases the volatility of a renter's housing costs by accumulating more draws of risky rent, it eventually decreases the volatility of an owner's housing costs by discounting the sale price farther into the future. Whether due to this or other considerations, a widespread view is that for many households owning is safer than renting.[1]

In this paper, we find instead that for many households, renting is not only safer than owning, it even is safer than full housing insurance. The key proximate mechanism is wage income risk. Home prices and rents are strongly positively correlated with wages both across locations at a point in time and, importantly, within locations over time. Within location, wage growth is 60% correlated with rent growth and over 40% with home price growth. Across locations at a point in time, log wages are 75% correlated with log rents and over 80% with log home prices. This creates strong, opposite-signed correlations between household earnings—a major source of hard-to-diversify risk—and the housing costs of owners versus renters. An owner's housing costs tend to be negatively correlated with wages, since wages are positively correlated with home prices and thus with the proceeds from selling a home. So an owner's housing costs tend to exacerbate wage income risk, increasing net income when wages are high and decreasing it when wages are low. A renter's housing costs, by contrast, are strongly positively correlated with wages, since wages are positively correlated with rents. So a renter's housing costs provide a valuable hedge against wage income risk, partially offsetting the effect of wage shocks on net income. These interactions with wage income risk are sufficiently strong that although owning leads to less volatility in *housing costs* for households with long horizons in a home, renting tends to lead to less volatility in net income—and so to less risk in consumption and welfare.

[1]This view is widespread in the economics literature, in popular financial advice, and among U.S. households, even in the wake of the housing bust of the Great Recession. For example, in a nationally representative survey of more than 50,000 households between 2010 and 2016, Adelino et al. (2018) find that 71% of U.S. households believe that housing is a "safe" investment, versus 55% for government and corporate bonds and 18% for stocks.

Figure 1: Relationships Between Wages, Rents, and Home Prices, 1960–2010



(a) rent vs. wage changes

(b) home price vs. wage changes

Notes. Changes in average location-level full-time wages, rents, and home values from 1960–2010 by Commuting Zone (CZ) relative to national-level changes. All prices are quality-adjusted, using the location-year fixed effects of the hedonic regression of prices on household and home characteristics as described in the text and shown in equation A.4 in Appendix A. Circle areas are proportional to the location's population share. β_{wls} and R^2 are obtained from regressing changes in housing costs on changes in wages, weighted by population. $\beta_{reverse}$ is obtained from the corresponding population-weighted reverse regression of wage changes on housing cost changes. Robust standard errors in brackets are clustered by state. Prices are adjusted to dollars of 2020 using the Consumer Price Index (CPI) for All Urban Consumers. Source: Decennial Census and ACS.

The key underlying mechanism is the substantial location-specific risk in wages and housing prices. This risk is characterized by strong positive relationships between location-specific changes in wages and housing prices. To fix ideas, consider the impact of owning versus renting in Detroit and San Francisco late in the 20th century. From 1960 to 2010, average annual wages declined by $8,400 in Detroit and increased by $17,700 in San Francisco (in 2020 dollars) relative to the national average (Figure 1). Detroit's relative decline imposed much greater financial losses on owners than renters. Owners experienced below-average capital gains on top of their depressed earnings ($101,000 reduction in average home prices relative to the national average), whereas renters received partial compensation for their depressed earnings in the form of lower rents ($1,600 reduction in average annual rents relative to the national average). Analogously, San Francisco's boom conferred much greater benefits on owners than renters. Owners experienced above-average capital gains on top of their increased earnings ($380,000 relative increase in average home prices), whereas renters experienced increases in their housing costs that partially offset their increased earnings ($6,700 relative increase in average annual rents). Ex ante, for households that do not know if their location will go the way of Detroit or San Francisco, renting provides a valuable hedge against the substantial location-specific risk in wages, whereas owning "doubles down" on it.

We begin by using 70 years of data on local markets in the U.S. (commuting zones) to characterize the implications of the within-location volatility of wages, rents, and home prices for the volatility of owners' versus renters' housing costs and net income (income net of housing costs) during a single housing spell. This analysis produces three main findings. First, the substantial volatility of housing costs is minimized by owning for long spells (longer than 25–30 years or so) and by renting otherwise. So were it not for other risk exposures, this finding would be in keeping with the view that owning is safer for households with long horizons in a home, though even in this case the requisite horizon is long, longer than typical lengths of time between moves (owners' average spell in a home is 15 years; American Housing Survey 1973–2021).

Second, growth in home prices and rents is highly correlated with growth in wages. Locations that experience faster growth in home prices and rents during a particular time period are much more likely to have experienced faster growth in wages as well, over all horizons from 10-year to 70-year periods. Hence, renting hedges wage income risk—lower rents "compensate" for lower wages and higher rents partially offset higher wages—and owning tends to exacerbate it.

Third, the interaction with wage income risk is sufficiently strong that for working households, renting *reduces* overall risk in net income—net income is less variable by renting than it would be even with full housing insurance—whereas owning increases it. This holds regardless of the horizon in a home. This suggests that for working households, renting is not only safer than owning, it is safer than full (hypothetical) housing insurance.

We then investigate, using a variety of approaches, the implications of these strong empirical regularities for the welfare consequences of the financial risk aspect of owning versus renting. Each approach points to the same conclusion: For most households, renting has a considerable risk benefit over owning. For example, we estimate that for a household that never moves, the risk benefit from always renting over always owning is worth on the order of $10,000 in lifetime present value.

Our investigation of mechanisms reveals the crucial role of the interaction with wage income risk. Wage income risk transforms a renter's exposure to rent risk from quite costly (welfare cost on the order of $8,000 for a household that never moves when failing to account for wage income risk) to quite valuable (welfare benefit on the order of $9,000). This is because renting provides a highly valuable hedge against wage income risk (worth on the order of $17,000) that dominates the cost of its contribution to cost-of-living risk (on the order of $8,000). The hedge against wage income risk is more beneficial than the increase in cost-of-living risk is costly mainly because wage income

risk is greater than cost-of-living risk. In other words, our estimates suggest that many households' exposure to wage income risk is sufficiently costly that the benefit of hedging it by renting exceeds the direct cost of being exposed to rent risk. For them, renting has a sizable risk benefit over not only owning, but even full housing insurance.

While multiple factors likely contribute to the strong positive relationships between location-specific changes in housing prices and wages, one set of mechanisms that the previous literature has identified as being especially important is the combination of significant location-specific shocks to labor demand (e.g., Blanchard and Katz, 1992; Autor et al., 2013; Hornbeck and Moretti, 2019), relatively inelastic supply of housing in many locations (Glaeser and Gyourko, 2005, 2018), and spatial equilibrium forces toward different locations being similarly attractive to potential residents (Rosen, 1974; Roback, 1982). Together, these factors can generate the large, persistent, strongly-positively-correlated differences across locations in the growth rates of wages and housing prices that drive our conclusions.

The crucial role of wage income risk highlights systematic heterogeneity in the risk consequences of renting versus owning. Households with less exposure to wage income risk enjoy a smaller risk benefit from renting than households with more exposure do (e.g., one- versus two-earner households), and retirees enjoy a benefit that is smaller still; in fact, for retirees with long horizons in a home, owning would be safer. Hence, the risk-minimizing life cycle profile involves renting during prime working years and, if the horizon in a home is long enough, potentially owning during retirement.

Households in locations with more volatile labor demand and less elastic housing supply enjoy a larger risk benefit from renting, since in such locations wages are both more volatile and more strongly correlated with rents and home prices.

Our main contribution is to quantify the effects of renting versus owning on the volatility of housing costs and net income, and the welfare consequences. This complements and extends previous research in three main ways. First, we use 70 years of data to characterize risk over life cycle-relevant horizons. Second, we use a wide range of approaches, including flexible non-parametric ones, to account for the full richness of the strong relationships between housing prices and wages over different horizons. Third, we use a variety of approaches to quantify the welfare consequences of different housing risk exposures.[2] Our central conclusion that for many households renting is safer than

[2] A related literature considers how wage income risk affects optimal or actual housing tenure choices. Campbell and Cocco (2003) and Yao and Zhang (2005) investigate optimal tenure choice in life cycle models with correlated wages and house prices. Davidoff (2006) investigates how heterogeneity in the correlation between earnings and home prices relates to heterogeneity in the amount of housing people own. Ortalo-Magné and Rady (2002) show that in theory a higher covariance between earnings and

owning contrasts with widely-held views, and our findings highlight important dimensions of heterogeneity that also differ from prevailing views. Although risk is just one determinant of optimal housing choices, our results suggest that it is an important determinant that is widely misunderstood. Our back-of-the-envelope calculations suggest that such misperceptions could increase the lifetime efficiency cost of the substantial net public subsidies to owner-occupied over rental housing by about $4,000 per household and could decrease welfare by even more.[3]

This paper also contributes new evidence on labor earnings risk, including its determinants and welfare consequences. This complements and extends the large literature on labor earnings risk (see Altonji et al., 2022, for a review), and the strand on local labor demand shocks in particular (see Moretti, 2011, for a review), by quantifying the implications of such shocks for lifetime risk in earnings and net income, accounting for interactions with housing costs. While idiosyncratic risk drives most of the risk in earnings over short horizons, we find that market-level risk drives much of the risk in lifetime earnings. Our findings also help clarify the welfare effects of wage income risk by quantifying the effect of housing tenure in either exacerbating or hedging it. This reveals an underappreciated yet powerful way to hedge this key risk: by renting one's home.

2 Theory: Housing Tenure and Wage Risk

This section uses a simple model to show how wage income risk can transform the riskiness of renting versus owning housing.[4] A household's tenure spell starts at time t and lasts T periods. In each future period $s \in \{1, 2, ...T\}$ and state of the world ω, it supplies one unit of labor perfectly inelastically, earning income $y_{\omega,t}$, and demands one unit of housing, also perfectly inelastically. Financial markets are perfect but incomplete: There are no borrowing constraints but earnings risk cannot be traded. There are two tradable assets, owner-occupied housing and a risk-free financial asset.[5]

rents reduces the riskiness of renting. Cocco (2000) finds that house price shocks are strongly positively correlated with income shocks in household-level panel data and investigates the implications for optimal house size in a life cycle model for a household that always owns. Research on housing risk that does not focus on wage income risk includes Rosen et al. (1984); Englund et al. (2002); Sinai and Souleles (2005, 2013).

[3]Such misperceptions are one reason our analysis centers on budget constraints rather than revealed preferences and sufficient welfare statistics.

[4]To show this as simply as possible, we use a minimal model that abstracts from many factors that could affect the riskiness of renting versus owning. Our quantitative analyses use richer models.

[5]Letting households invest in a richer set of financial assets such as stocks does not affect our analysis. Our baseline only uses location-specific risk which by construction is uncorrelated with aggregate returns. Moreover, historically, stock returns have been only weakly correlated with both aggregate housing returns (Jordà et al. 2019) and labor income risk (Davis and Willen 2014), although there is an active debate about

If the household rents, it pays the going rent $R_{\omega,t+s}$ in each period. If it owns, it pays known price P_t to buy a home immediately before period one and receives $P_{\omega,t+T}$ when it sells its home after living in it in period $t+T$ (received after $R_{\omega,t+T}$ is realized, i.e., the ex dividend price). The present value of non-housing consumption, C_ω, equals the present value of net income, i.e., the present value of labor income, Y_ω, less the present value of housing costs, H_ω, in state of the world ω:

$$C_\omega = Y_\omega - H_\omega = \sum_{s=1}^{T} \frac{y_{\omega,t+s}}{(1+r_t)^s} \begin{cases} -\sum_{s=1}^{T} \frac{R_{\omega,t+s}}{(1+r_t)^s} & \equiv -R_\omega \quad \text{if rent,} \\ +\frac{P_{\omega,t+T}}{(1+r_t)^T} - P_t & \equiv -O_\omega \quad \text{if own,} \end{cases} \tag{1}$$

where r_t is the risk-free rate over the tenure spell known at t. This shows that renters are short housing risk during their stay (the present value of rent risk enters non-housing consumption with a negative sign), while owners are long housing risk (the present value of sale price risk enters with a positive sign).

Consumption risk, measured as the variance of the present value of non-housing consumption, can be written as:

$$Var(C_\omega) = Var\left(Y_\omega\right) + \overbrace{\left[\underbrace{Var(H_\omega)}_{\text{“Partial effect”}} - \underbrace{2Cov(Y_\omega, H_\omega)}_{\text{“Portfolio effect”}}\right]}^{\text{“Net housing risk”}}, \tag{2}$$

where the risk comes from risk in wages and housing prices.[6] The contribution of housing to overall consumption risk, in brackets, depends on not only its riskiness as viewed in isolation, the “partial effect,” but also its covariance with the rest of the household’s portfolio, the “portfolio effect” (adopting the terminology of Lockwood, 2022). Other things equal, housing costs that covary positively with income are less costly because they reduce the welfare cost of income risk. Such costs are on average lower when income is lower and higher when income is higher, which partially compensates for and offsets income shocks.

In the special case in which housing prices are uncorrelated with wages, the riskiness of renting versus owning depends solely on the “partial effect”—the variance of housing costs. It amounts to a comparison of the variance of the present value of rents during

the latter for long-run risk (Cocco 2005; Benzoni et al. 2007; Huggett and Kaplan 2016).

[6]To be sure, the effect of a particular risk exposure on the variance of consumption or net income is not a sufficient statistic for its welfare cost (Rothschild and Stiglitz, 1970). We use it here only as a simple measure to fix ideas. Our quantitative analyses use exact welfare measures. A related point is that risk in prices is different from risk in income; see footnote 10.

the household's stay to the variance of the present value of the home price at the end of the household's stay:[7]

$$Var(C_\omega^{rent}) > Var(C_\omega^{own}) \Leftrightarrow Var(R_\omega) > Var(O_\omega) \Leftrightarrow Var\left(\sum_{s=1}^{T} \frac{R_{\omega,t+s}}{(1+r_t)^s}\right) > Var\left(\frac{P_{\omega,t+T}}{(1+r_t)^T}\right).$$

In this case, the riskiness of renting versus owning depends crucially on the household's horizon, the time it will spend in the home, since lengthening the horizon has opposite-signed effects on the variance of housing costs of owners versus renters. An increase in horizon increases a renter's exposure to rent risk, since it adds exposure to rents farther into the future. By contrast, under standard assumptions about discounting and home prices, an increase in horizon eventually decreases an owner's exposure to home price risk, since it pushes farther into the future—and so discounts to a greater extent the risk from—the eventual sale of the home. In the limit with an infinite horizon, there would be no sale price risk. Hence, without a portfolio effect, for households with sufficiently long horizons, owning is safer than renting: $Var(C_\omega^{rent}) - Var(C_\omega^{own}) \to Var(R_\omega) > 0$. This logic, due to Sinai and Souleles (2005), supports the prevailing view that for households with long enough horizons in a home, owning is safer than renting.

More generally, the riskiness of renting versus owning depends on the portfolio effect. The portfolio effect can have a major impact on the riskiness of owning versus renting because renters and owners have opposite-signed exposures to housing prices and hence opposite-signed portfolio effects. Renters have a negative exposure to rents during their stay in a home, whereas owners have a positive exposure to expected future rents after their stay through their sale price. If housing prices covary positively with wages, the portfolio effect reduces the riskiness of renting (since rent risk hedges wage income risk) and increases the riskiness of owning (since home price risk exacerbates wage income risk). If the covariance is positive enough, the portfolio effect can even make renting safer than full housing insurance. Exposure to rent risk reduces the variance of net income if $2Cov(Y_\omega, R_\omega) > Var(R_\omega)$, i.e., if the slope of the 'reverse regression' of income on rents exceeds one half:

$$\beta_{Y|R} \equiv \frac{Cov(Y_\omega, R_\omega)}{Var(R_\omega)} > \frac{1}{2}.$$

More generally, a household that could not otherwise hedge its wage income risk bene-

[7]This adopts a standard perspective of risk from the time a household chooses to own or rent, when the initial purchase price of the home, P_t, is known. From earlier perspectives where the purchase price is not yet known (e.g., "behind the veil of ignorace" where the individual does not know which location they will live in), risk in the purchase price also contributes to the riskiness of owning. Because the initial prices (wages, rents, and home prices) are also strongly positively correlated (see Appendix Figure A.1), adding such risk reinforces our main conclusion that renting hedges wage income risk and owning exacerbates it.

fits from exposure to rent risk if rent risk hedges wage income risk strongly enough. In that case, the fact that owning reduces exposure to rent risk is a cost, not a benefit, of owning.[8]

3 Data and Empirical Approach

Data.— Our main data source is the IPUMS USA Decennial Census database from 1940 to 2010 (Ruggles et al. 2018). For our purposes, its main strengths are its large sample size, detailed geographic and demographic information, and long time coverage. These enable us to construct measures of earnings, rents, and home prices at the local market level over a 70-year time period. The long time coverage is crucial for learning about the risk facing households over life cycle horizons.[9]

We construct market prices from the household-level data in two main steps: defining markets and measuring the prices in those markets. Our baseline market definition is a Commuting Zone (CZ). A CZ is a county or set of counties chosen so as to approximate a single labor market with a common workforce (Tolbert and Sizer 1996). CZs are widely used in economics as approximations to local markets. They also have the advantage of being amenable to long-run analyses (Eckert et al. 2018), since county boundaries are stable over time, whereas the boundaries of other common definitions of local markets change frequently over time (e.g., MSAs). For households whose county is not uniquely identified, we use IPUMS-supplied allocation factors for each Decennial Census to assign their counties. We drop Alaska and Hawaii because their county population data are not available before they obtained statehood in 1959. This leaves us with 720 out of 744 CZs covering over 99% of the U.S. residential population.

The second step is to measure market prices. This involves choosing samples and constructing market prices from individual- and household-level data. Our main sample is households not living in group quarters. For individual-level variables such as wage income or age, we use the values of the highest earner. To measure market wages, we further restrict the sample to households whose highest earner worked full time and is

[8]The portfolio effect can transform important rules of thumb about the riskiness of owning versus renting. For example, lengthening the horizon may not reduce the riskiness of owning relative to renting if renting provides a valuable hedge against wage income risk, since the value of this hedge tends to increase in the horizon (which works against the more widely-appreciated factor discussed above that an owner's sales price risk is eventually decreasing in the horizon).

[9]To the best of our knowledge, no other source can match the Census's length of coverage of the key variables of interest. Moreover, the quality of the resulting price measures appears to be high. For example, we find that home price changes and the volatility of home price growth rates (both aggregate and location-level) in the Census track closely those in home transaction data such as Zillow or FHFA in periods in which the latter are available; see Appendix D.

between 25 and 60 years. These sample restrictions are aimed at producing clean measures of market prices that are comparable over long periods. The age restriction, for example, reduces the effect of schooling and retirement on measured wages. To make the universe of housing units that report housing costs consistent, we restrict units to non-farm homes that are not used for business purposes, have less than 10 acres of land, and are not condominiums, mobile homes, or trailers.

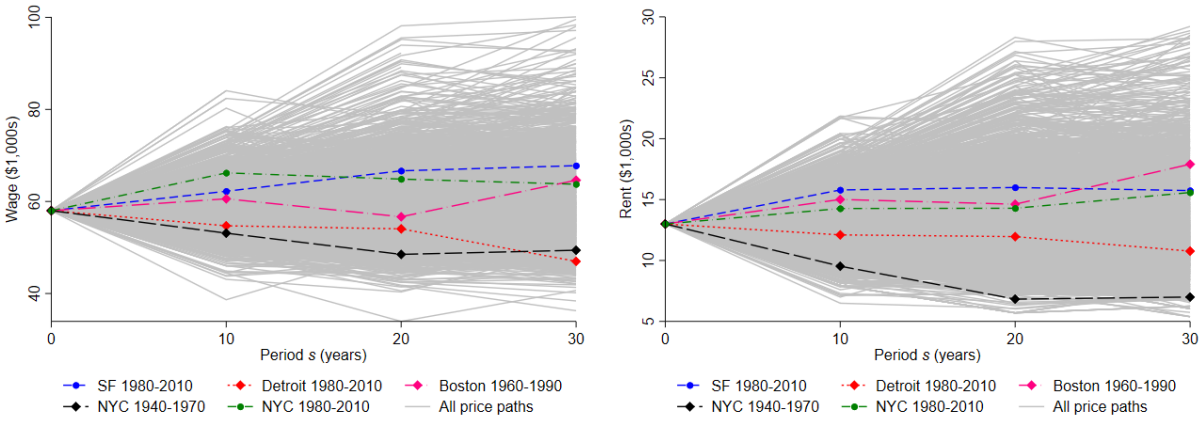
To construct market prices, we follow the common practice of using hedonic regressions to remove the effects of observable differences in quality and quantity. Separately for each Census year, we regress wage income on location fixed effects, year fixed effects (i.e., a constant), and fixed effects for household age and educational attainment. Similarly, we regress rents and home prices on location fixed effects, year fixed effects, and fixed effects for home age, home size, and other home characteristics. The location fixed effects are our measure of quality-adjusted market prices in a given market-decade (equation (A.3) in the appendix). Household-level analyses use household weights, and location-level analyses use population weights. Dollar values are converted to real 2020 dollars using the CPI-All Urban series. Full details are in Appendix A, and Tables 1 and 2 below report summary statistics.

Approach.— To understand the risk properties of different housing exposures, we consider the conceptual experiment of a household that faces exogenous risk in its income and housing costs from risk in wages, rents, and home prices in its labor and housing markets. This experiment focuses on the budget constraint. The question is: What does the volatility in wages and housing prices mean for the risk in net income of owners versus renters?[10]

The main challenge, common to all analyses of risk, is that only the single, realized state of the world is observed. We use the standard approach of modeling risk (i.e., variation across states of the world) using cross-sectional and time-series variation, with a focus on location-specific changes over time: differential changes in the key prices in different locations during the same time period. Specifically, we consider a household that initially faces the wage, rent, and home price of a median full-time working household in 2010:

[10]In reality, households can re-optimize in response to price changes, which makes risk in prices different from risk in net income. Ignoring the effects of re-optimization or, equivalently, assuming perfectly inelastic labor supply and housing demand with respect to changes in these prices, is a convenient approximation that captures the first-order welfare effect of price changes (since by the envelope theorem, the mechanical effect of a price change—i.e., the implied change in household income net of expenditures holding behavior fixed—is a first-order approximation to its welfare effect). The full welfare effect includes higher-order terms in the size of the price change, which reflect the gain from re-optimization. The approximation is better the less elastic are labor supply and housing demand with respect to these prices.

Figure 2: Non-parametric Model of Wage and Housing Cost Risk



(a) wage paths

(b) rent paths

Notes. Illustration of the non-parametric model of price risks showing 30-year price paths of wages and rents starting from the same initial price at time t, \$58,000 for wages and \$13,000 for rents. Price paths only contain location-specific changes (corresponding to price component $X^{loc}_{i,t+s}$ in Appendix A) and no aggregate or within-market idiosyncratic risk. We omit values that are 120% above the 99th percentile from the figure (but not from the analysis) to improve its appearance. Source: Decennial Census and ACS.

\$58,000 per year, \$13,000 per year, and \$186,000, respectively.[11] We assume that the possible future realizations of prices are those that arise from letting this initial price vector evolve at the observed location-specific price changes in each of the 720 locations during different time periods. In other words, the realized path of prices in each location during different time periods is taken to be the path in a corresponding state of the world.

Figure 2 illustrates the distribution of such paths of wages and rents across all locations and for a 30-year housing spell. The risk comes from the fact that the household does not know if its location will boom like San Francisco in the late 20th century, bust like Detroit, or experience the growth of any other location-by-time-period (see also Figure 1). The resulting risk model, discussed in more detail in Appendix B, isolates across-location heterogeneity in the growth of wages, rents, and home prices over time. It is a simple, flexible, non-parametric model of the underlying data generating process. We adopt it as a baseline in Sections 4 and 5, though we test robustness to many other risk models as well. The baseline risk model, in isolating location-specific growth, excludes aggregate

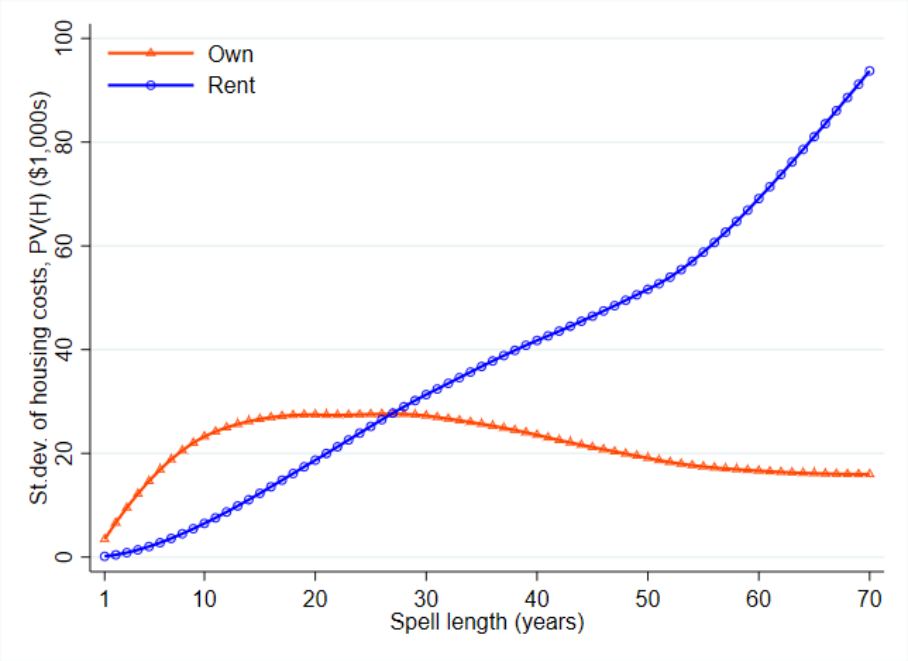
[11]The wage is the median wage among full-time working households in 2010. The rent and home price are the median rent and home value in 2010 among households with earnings within \$1,000 of the median (not the median rent among renters or the median home price among owners), since the goal is to capture the housing costs of a median-earning household (who earns more than the median renter and less than the median owner). Using median rent and home prices instead would increase the risk benefits from renting (because of the median renter has a larger housing budget share) and decrease the risk costs from owning (because the median owner has a smaller housing budget share).

risk and within-market idiosyncratic risk. As discussed in Section 5, including aggregate and idiosyncratic risk tends to reinforce our key conclusions. Focusing on location-specific risk has the additional benefit of avoiding measurement error from imperfect deflating of nominal prices (e.g., Hausman 2003; Crone et al. 2010) or changes in the survey.

4 Descriptive Analysis of Housing Risk

Fact 1: The considerable volatility in housing costs is minimized by owning if the horizon is long, and by renting otherwise.

Figure 3: Volatility in Housing Costs by Length of Spell



Notes. Standard deviation of the present value of housing costs by the length of the spell in the home. Rents and home prices begin at those of a median full-time working household in 2010: $13k per year and $186k, respectively (corresponding to price component $X^{loc}_{i,t+s}$ in Appendix A). Each then grows according to the location-specific path observed in different location-by-time-periods from 1940–2010. The resulting risk process, discussed in more detail in Appendix A, isolates across-location heterogeneity in the evolution of rents and home prices over time. In excluding aggregate and within-market idiosyncratic risk, it tends to understate the riskiness of owning relative to renting. Rents and home prices are quality-adjusted as described in the text before calculating present values and expressed in dollars of 2020 using the Consumer Price Index (CPI) for All Urban Consumers. The discount rate is 2.5%. Source: Decennial Census and ACS.

Figure 3 plots the standard deviation of the present value of housing costs for owners versus renters as a function of the time horizon in the home, using a discount rate of 2.5%, the post-war average real interest rate on long-term government bonds (Jordà et

al., 2019). For renters, housing cost volatility increases roughly linearly in the horizon as risk accumulates. For owners, volatility starts at a higher base, increases rapidly in the horizon up to about ten years as risk accumulates, plateaus for horizons of roughly 10–30 years as greater discounting roughly offsets risk accumulation, and then decreases to a low level at long horizons as discounting dominates risk accumulation. As a result, volatility in housing costs is minimized by owning if the horizon is long (longer than 25–30 years or so in this risk model) and by renting otherwise.[12]

Whether a household owns or rents can have a large impact on the volatility in housing costs as shown in Table 1. For example, even in these calculations with location-specific risk alone, without idiosyncratic or aggregate risk, the standard deviation of the present value of housing costs for a ten-year spell is about $23,300 for owners versus $6,500 for renters (panel C). For a 50-year spell it is $19,100 for owners versus $51,600 for renters. The considerable volatility in housing costs comes from the significant, persistent heterogeneity in location-specific growth rates of home prices and rents. For instance, the standard deviation of location-specific changes of home prices and rents over a decade is $37,800 and $1,600 respectively (panel A). The persistence of these changes is shown in panel B, which reports the standard deviation of annualized growth rates (the log change divided by the square root of the horizon). If location-specific prices followed a random walk, then this volatility would be constant over all horizons from 10 to 70 years, which it generally is for wage and rent growth, while it declines only slightly for home price growth. Finally, as is the case with most other asset classes, home price growth rates are more volatile than the growth rates of the rental cash flows from these assets (Shiller 1981, Campbell et al. 2009, Plazzi et al. 2010).

Implications.— If a household's other risk exposures were independent of its housing costs during a spell, renting would be safer for housing spells of up to 25–30 years or so and owning would be safer for longer spells. More generally, risk in housing costs is significant even when ignoring idiosyncratic and aggregate risk, and a household's exposure to that risk depends crucially on whether it owns or rents its home. The choice of whether to own or rent may be a major determinant of a household's overall risk

[12]For interpreting this "crossover horizon," two considerations are important. First, in terms of risk, moves to correlated markets, which account for a large share of moves (Sinai and Souleles, 2013), are similar to no move at all (at least if idiosyncratic risk is ignored). So owning could lead to lower volatility not only for a 30-year spell in the same location but also perhaps for a 35-year period with one move to a correlated market, say. Second, this analysis excludes idiosyncratic risk. Given the considerable idiosyncratic risk in home prices (Giacoletti, 2021), this tends to lead it to understate the crossover horizon. Of course, there is idiosyncratic risk in rents too (Eichholtz et al., 2021). But even if we add significant idiosyncratic rent risk while continuing to exclude idiosyncratic home price risk, the crossover horizon remains long (e.g., with fully persistent idiosyncratic rent shocks with a standard deviation of 9% of the mean rent, the crossover horizon is 26 years). Adding aggregate risk increases the crossover horizon to about 40 years. We include moves and idiosyncratic and aggregate risk in Section 5.

Table 1: Volatility of Location-Specific Housing Cost and Income Changes

	Spell length, T (years):						
	10	20	30	40	50	60	70
Panel A: Standard deviation of changes ($1,000)							
Rent, $R_{\omega,t+T} - R_t$	1.6	2.4	2.8	3.0	3.3	3.7	4.2
Home price, $P_{\omega,t+T} - P_t$	37.8	51.9	64.5	70.2	72.0	77.5	90.0
Income, $y_{\omega,t+T} - y_t$	3.2	4.6	5.7	6.6	7.4	8.5	9.7
Panel B: Standard deviation of annualized log changes (%)							
Rent, $(\ln R_{\omega,t+T} - \ln R_t)/\sqrt{T}$	3.8%	4.1%	3.8%	3.6%	3.5%	3.5%	3.7%
Home price, $(\ln P_{\omega,t+T} - \ln P_t)/\sqrt{T}$	6.0%	5.4%	5.5%	5.4%	5.0%	4.8%	5.2%
Income, $(\ln y_{\omega,t+T} - \ln y_t)/\sqrt{T}$	1.7%	1.7%	1.7%	1.7%	1.7%	1.8%	1.9%
Panel C: Standard deviation of present values ($1,000)							
Prices:							
Rent, $R_\omega = \sum_{s=1}^{T} \frac{R_{\omega,t+s}}{(1+r_t)^s}$	6.5	18.7	31.3	41.8	51.6	69.1	93.7
Own, $O_\omega = P_t - \frac{P_{\omega,t+T}}{(1+r_t)^T}$	23.3	27.5	27.3	23.6	19.1	16.7	16.0
Income, $Y_\omega = \sum_{s=1}^{T} \frac{y_{\omega,t+s}}{(1+r_t)^s}$	12.8	35.5	58.4	81.8	110.2	143.0	180.5
Net Income:							
Renter, $Y_\omega - R_\omega$	10.6	29.3	49.1	70.2	89.7	108.3	131.7
Owner, $Y_\omega - O_\omega$	31.1	50.2	69.9	87.4	113.6	146.2	183.6
Full housing insurance, $Y_\omega - E(R_\omega)$	12.8	35.5	58.4	81.8	110.2	143.0	180.5

Notes. The table shows the standard deviation of (log) changes of location-specific wages, rents, and home prices and present values (PVs) of renters and owners housing costs and net income for different spell length. Wages and housing costs are quality-adjusted as described in the text. All changes are relative to the same initial price ($R_t = \$13k$, $P_t = \$186k$, $y_t = \$58k$) and the real discount rate used for computing PVs is the average post-war yield on long-run Government bonds, $r_t = 2.5\%$. Panel B shows changes in quality-adjusted log prices (location-year fixed effects $\hat{\lambda}_{ct}$ of the hedonic regression, equation (A.3) in the Appendix), normalized by the square-root of the spell length to make them comparable (see Appendix Figure A.2). If log price changes are permanent (i.e., follow a random walk), then the standard deviation of annualized log changes are constant irrespective of spell length. Panel A converts these quality-adjusted growth rates to dollar changes (corresponding to price component X_{it}^{loc} in equation (A.6) of the appendix). Panels A and B pool decadal data from Census years 1940 to 2010 so that the standard deviation of 10-year changes is based on 5,040 CZ-year observations (720 CZs times 7 10-year changes 1940–1950 to 2000–2010), 2,880 observations for 40-year changes, and 720 observations for 70-year changes. Panel C interpolates between census years to annual data spanning 1940 to 2010. The standard deviation of 10-year PVs is based on 43,920 CZ-year observations (720 CZs times 61 PVs for 1941–1950 to 2001–2010), 22,320 observations for 40-year PVs, and 720 observations for 70-year PVs. All calculations use population weights for the 720 CZs. Source: Decennial Census and ACS.

exposure.

Fact 2: Rent growth and home price growth are strongly positively correlated with wage growth over all horizons.

The most important financial risk for many households is wage income risk. A large share of this uninsured risk is location specific. For instance, over a 30-year period, the

location-specific volatility of the present value of wage income is roughly $60,000, about twice that of housing costs. This is despite that rent growth is twice as volatile as wage growth and home price growth is three times as volatile (Table 1). The fact that wages are much larger than housing costs more than offsets the fact that wage growth is less volatile than housing cost growth, because consumption risk is ultimately measured in dollars.

The net effect of income and housing cost risk on consumption risk depends crucially on the interaction of these two important risk exposures. Figure 4 shows the joint distributions across locations of rent growth and wage growth (left panels) and home price growth and wage growth (right panels); see also Figure 1. Table 2 shows correlations of dollar changes (panel A), growth rates (panel B), and present values (panel C) of housing costs with wage income. We find that for all horizons from 10–70 years and all specifications, rent changes and home price changes are strongly positively correlated with wage changes. Ten percent faster wage growth over a decade is associated with 12% faster rent growth and 20% faster home price growth. Significant shares of the variation in rents and home prices are associated with variation in wages, with R^2 statistics ranging from 33–60% for rents and 18–54% for home prices depending on the horizon, time period, and specification.[13] Heterogeneity and robustness analyses are in Appendix D.

Implications.— Renting hedges wage risk and owning exacerbates it. A renter's housing costs tend to be lower when wages are lower and higher when wages are higher, which mitigates wage income risk. An owner's housing costs tend to be higher when wages are lower and lower when wages are higher, which exacerbates wage income risk. So for working households, renting is safer than it would appear to be from its contribution to risk in the cost-of-living alone, and owning is riskier. More generally, a household's exposure to wage income risk is a crucial determinant of the riskiness of renting versus owning housing.

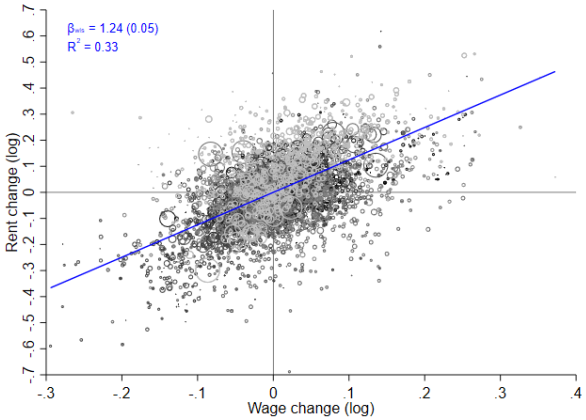
Fact 3: For working households, renting reduces net income risk and owning increases it, regardless of the time horizon in the home.

Figure 5 plots the standard deviation of the present value of net income (income minus housing costs) of a working household as a function of the time horizon in the home under three housing scenarios: own, rent, and full housing insurance. Table 1, panel C provides associated statistics and Appendix D contains heterogeneity and robustness

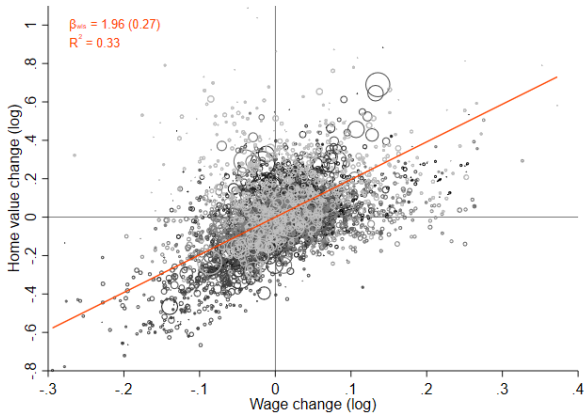
[13] In Appendix D we provide evidence of strong positive relationships between national-level changes in wages and housing prices as well, especially for rents. We also find similarly strong relationships between housing prices and national income (GDP per capita) across countries and over long time periods.

Figure 4: Relationships Between Wage, Rent, and Home Price Growth

A. 10-year growth

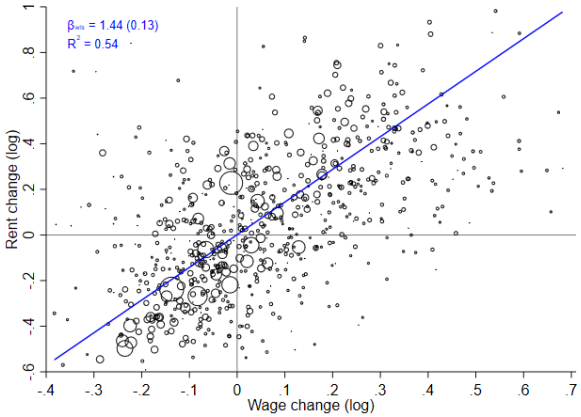


(a) rent vs. wage growth

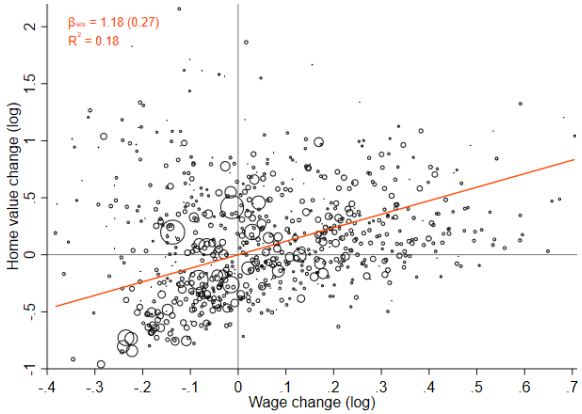


(b) home price vs. wage growth

C. 70-year growth



(c) rent vs. wage growth



(d) home price vs. wage growth

Notes. Changes in average location-level full-time log wages, log rents, and log home values from 1940–2010 by Commuting Zone (CZ) relative to national-level log changes (corresponding to $\Delta_s \ln x^{loc}_{c,t+s}$ with $s = 10$ and 70 in Appendix A). All prices are quality-adjusted, using the location-year fixed effects of the hedonic regression of log prices on household and home characteristics as described in the text. Panel A pools 10-year log changes (5,040 observation from 720 CZs times 7 decades); panel B shows the single cross-section of 70-year log changes across 720 CZs. Circle areas are proportional to the location's population share. Observations in darker shades are from more recent decades. β_{wls} and R^2 are obtained from regressing log changes in housing costs on log changes in wages, weighted by population. Robust standard errors in brackets are clustered by state. Prices are adjusted to dollars of 2020 using the Consumer Price Index (CPI) for All Urban Consumers. Source: Decennial Census and ACS.

Table 2: Correlation of Location-Specific Changes in Housing Costs and Income

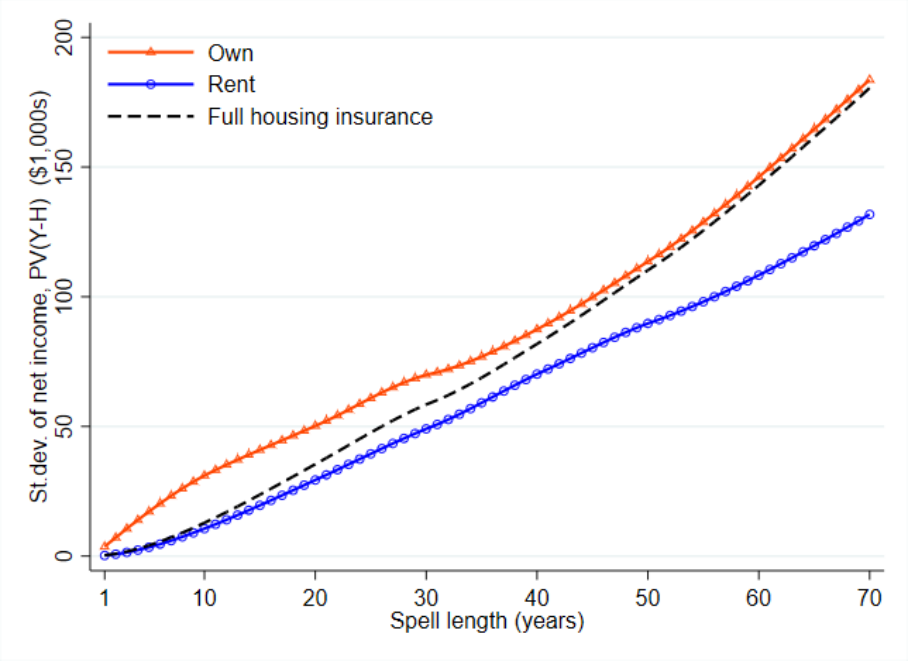
	Spell length, T (years):						
	10	20	30	40	50	60	70
Panel A: Dollar changes							
Rent, $Corr(R_{\omega,t+T} - R_t, y_{\omega,t+T} - y_t)$	0.57	0.55	0.58	0.54	0.57	0.66	0.71
Home price, $Corr(P_{\omega,t+T} - P_t, y_{\omega,t+T} - y_t)$	0.54	0.38	0.39	0.32	0.27	0.28	0.30
Panel B: Log changes							
Rent, $Corr(\ln R_{\omega,t+T} - \ln R_t, \ln y_{\omega,t+T} - \ln y_t)$	0.57	0.55	0.58	0.56	0.59	0.68	0.74
Home price, $Corr(\ln P_{\omega,t+T} - \ln P_t, \ln y_{\omega,t+T} - \ln y_t)$	0.57	0.44	0.46	0.41	0.36	0.39	0.43
Panel C: Present values							
Rent, $Corr(R_\omega, Y_\omega)$	0.57	0.56	0.54	0.51	0.59	0.68	0.71
Own, $Corr(O_\omega, Y_\omega)$	-0.43	-0.26	-0.23	-0.10	-0.09	-0.14	-0.16
Panel D: Coefficients of reverse regressions of present values, $Y_\omega = \alpha + \beta_{Y\|H} H_\omega + \epsilon_\omega$							
Rent, $\beta_{Y\|R}$	1.12	1.07	1.01	1.01	1.27	1.41	1.37
	(0.08)	(0.10)	(0.13)	(0.17)	(0.17)	(0.15)	(0.13)
Own, $\beta_{Y\|O}$	-0.24	-0.34	-0.49	-0.36	-0.53	-1.16	-1.77
	(0.04)	(0.14)	(0.24)	(0.32)	(0.50)	(0.73)	(0.94)

Notes. The table shows the correlation of log changes of location-specific housing costs (rents and home prices) and wages (full-time wage and salary income) of renters and owners housing costs and net income for different spell length. Wages and housing costs are quality-adjusted as described in the text. All changes are relative to the same initial price ($R_t = \$13k$, $P_t = \$186k$, $y_t = \$58k$) and the real discount rate used for computing present values (PVs) is the average post-war yield on long-run Government bonds, $r_t = 2.5\%$. Panel B shows changes in quality-adjusted log prices (location-year fixed effects $\hat{\lambda}_{ct}$ of the hedonic regression, equation (A.3) in the Appendix). Panel A converts these quality-adjusted growth rates to dollar changes (corresponding to price component X_{it}^{loc} in equation (A.6) of the appendix). Panels A and B pool decadal data from Census years 1940 to 2010 so that the correlation coefficient of 10-year changes is based on 5,040 CZ-year observations (720 CZs times 7 10-year changes 1940–1950 to 2000–2010), 2,880 observations for 40-year changes, and 720 observations for 70-year changes. Panel C interpolates between census years to annual data spanning 1940 to 2010. The standard deviation of 10-year PVs is based on 43,920 CZ-year observations (720 CZs times 61 PVs for 1941–1950 to 2001–2010), 22,320 observations for 40-year PVs, and 720 observations for 70-year PVs. All calculations use population weights for the 720 CZs. Robust standard errors in brackets of panel D are clustered by state. Source: Decennial Census and ACS.

analyses. While insurance markets for housing risk are currently not available to households, we can imagine hypothetical futures markets where households could achieve full housing cost insurance either by pre-selling their home at a price known at the time of purchase, or by pre-paying all future rents for the entire housing spell at prices already known in the initial period. In that case, net income risk is equal to income risk.

Figure 5 and Table 1 show that regardless of the time in the home, the standard deviation of a working household's net income is minimized by renting, is greater with full housing insurance, and is maximized by owning. The advantage of renting over full housing insurance increases in the time horizon as risk accumulates. The disadvantage

Figure 5: Risk in Net Income by Length of Spell



Notes. Standard deviation of the present value of net income by the length of the spell in the home. Net income is full-time wage income net of housing costs. Prices begin at those of a median full-time working household in 2010 ($58,000 wage income per year, $13,000 rent expenses per year, and a home price of $186,000, corresponding to price component $X^{loc}_{i,t+s}$ in Appendix A) and then grow according to the location-specific path observed in different location-by-time-periods from 1940–2010. The resulting risk process, discussed in more detail in Appendix A, isolates across-location heterogeneity in the evolution of rents and home prices over time. In excluding aggregate and within-market idiosyncratic risk, it tends to understate the riskiness of owning relative to renting. The discount rate is 2.5%. Source: Decennial Census and ACS.

of owning relative to full housing insurance first increases and then decreases in the horizon as risk accumulation first dominates and then gets dominated by discounting.

How could a renter's exposure to the substantial risk in rents *reduce* net income risk relative to full housing insurance? Recall from Section 2 that exposure to housing risk decreases the variance of net income if the hedge of income risk dominates the partial effect of housing risk itself, which occurs if the slope of the reverse regression of income on housing costs exceeds one half. Table 2, panel D shows that for all horizons from 10–70 years, this slope significantly exceeds one half for renters (e.g., $\widehat{\beta}_{Y|R} = 1.37$ over a 70-year horizon) and is negative for owners ($\widehat{\beta}_{Y|O} = -1.77$). Indeed, a renter's housing costs covary positively not only with gross income ($\beta_{Y|R} > 0$, as shown in Fact 2) but even with net income ($\beta_{Y|R} > 1$). As a result, renters' housing costs tend to be higher when net income is higher and lower when net income is lower, which reduces net income risk.[14]

[14]Renting would reduce net income risk even for households with much less exposure to wage income risk than the typical working household. We estimate that even for a household whose exposure to wage

Decomposing the slope for renters into $\beta_{Y|R} = Corr(Y_\omega, R_\omega) \times SD(Y_\omega)/SD(R_\omega)$, the hedge of wage income risk dominates the partial effect of rents being risky because (i) rent risk is a good hedge against wage income risk ($Corr(Y_\omega, R_\omega) \approx 0.6$ on average over all horizons), and (ii) wage income is significantly more volatile than rents ($SD(Y_\omega)/SD(R_\omega) \approx 2.0$). In second-best contexts in which people face multiple risks, greater exposure to one risk does not imply greater exposure to risk on net. Given the substantial risk in wages for working households, renting's hedge of wage income risk (a moderate reduction in a large wedge) dominates its contribution to cost-of-living risk (moderate increase in a smaller wedge).

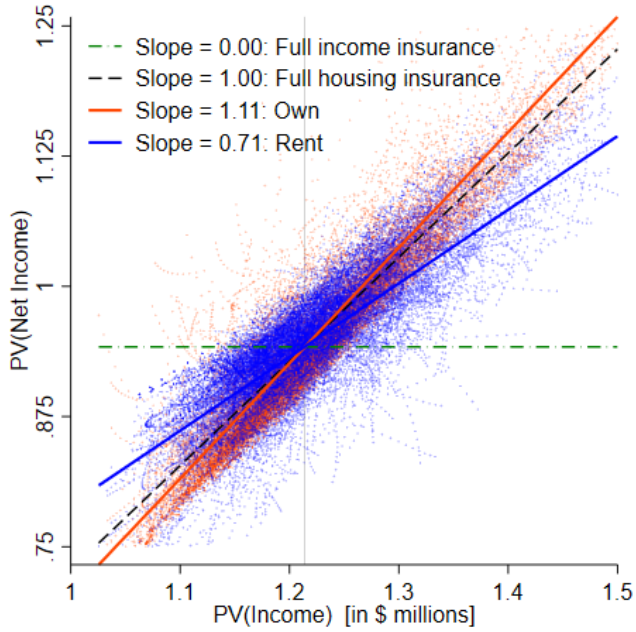
Figure 6 illustrates graphically, using a 30-year horizon, how the housing tenure choice affects the extent to which income shocks pass through to to net income. If a household is fully insured against housing cost risk, then shocks to income translate one-for-one into shocks to net income, resulting in a slope of one. On the other extreme, with full income insurance, shocks to income do not affect net income (a slope of zero). Because a renter's housing costs covary positively with income ($\widehat{\beta}_{Y|R} = 1.01$ for a 30-year horizon), renting reduces the pass-through of income shocks to net income shocks, dampening them by 29% (from a slope of one with full housing insurance to 0.71). Because an owner's housing costs covary negatively with income ($\widehat{\beta}_{Y|O} = -0.49$ for a 30-year horizon), owning amplifies income shocks by 11% (a slope of 1.11).

This connects the three main findings of the descriptive analysis. Renting reduces net income risk even relative to full housing insurance (Fact 3) because its valuable hedge against wage income risk (Fact 2) dominates its moderate contribution to cost-of-living risk (Fact 1), which in turn is largely because wage income risk is greater than rent risk. Owning increases net income risk relative to full housing insurance (Fact 3) because of reinforcing effects of increasing cost-of-living risk, especially if the horizon is not very long (Fact 1), and of exacerbating wage income risk (Fact 2). As a result, the net income of working households tends to be considerably more volatile if they own than if they rent.

Implications.— Regardless of the time horizon in the home, renting reduces a typical working household's net income risk and owning increases it. The variance of net income is minimized by renting, is greater with full housing insurance, and is maximized by owning. Hence, for working households, exposure to rent risk likely is valuable, not

income risk was 35% that of a typical household (i.e., whose wage was a weighted average of the realized wage and the mean wage, with a 35% share on the realized wage), renting would reduce net income risk relative to full housing insurance for all horizons (and would reduce risk relative to owning by even more). Renting reduces net income risk relative to owning for all horizons even if we add significant idiosyncratic rent risk (e.g., fully persistent idiosyncratic rent shocks with a standard deviation of 9% of the mean rent) while continuing to exclude idiosyncratic home price risk.

Figure 6: Translation of Income Risk into Net Income Risk by Housing Position



Notes. Present value of net income (income net of housing costs) as a function of the present value of gross (wage) income for a household that lives and works in the same location for thirty years under three scenarios: renting, owning, and having full, actuarially fair housing insurance. An observation is a CZ-by-30-year present value pair (corresponding to price component $X_{i,t+s}^{loc}$ in Appendix A). Wages, rents, and home prices are quality-adjusted as described in the text before calculating present values. Prices begin at that of a median full-time working household in 2010 ($58,000 per year for wages, $13,000 for rents, and $186,000 for home prices) and then grow according to the location-specific path observed in different location-by-time-periods from 1940–2010. Price paths exclude aggregate and within-market idiosyncratic risk. Regression slopes are from a population-weighted regression of 30-year present values of net income on 30-year present values of gross income. Source: Decennial Census and ACS.

costly. Owning not only involves sale price risk, it also eliminates the valuable hedge against wage income risk that rent risk provides. So while the particular time and place in which a household lives has a large impact on its net income, this risk is much greater for owners than renters.

Robustness.— Facts 1–3 are highly robust, including to using different sample periods (e.g., excluding the Great Recession, the recent housing boom and bust cycle, or the post-war economic boom), different market definitions (e.g., counties instead of CZs), and different measures of income and housing costs (e.g., total family income or gross instead of contract rents). They are robust to dropping households that migrated between CZs and to controlling for changes in population and housing quality within locations over time. We find similar results if we restrict the analysis to the modal 5-room single family home, and similar patterns hold across the distribution of wages and housing costs (e.g., at the 25^{th} or 75^{th} percentiles). Placebo-type tests suggest that the results are not driven by any spurious relation or omitted variable. For example, we find that factors that

are largely determined at broader geographic levels are not associated with location-specific wage changes, such as Social Security income (a federal program) or electricity demand (supplied competitively across many CZs). Full details of these and additional robustness checks are in Appendix D.

5 Risk Cost of Owning Versus Renting

In this section, we go beyond the preceding summaries of the raw data to quantify the ex ante welfare costs of the risk aspect of owning versus renting.

5.1 Estimates

Approach.— *Risk.—* Our baseline risk model extends that described in Section 3, for a single housing spell, to a full life cycle, including allowing for moves to different locations. We consider a household that faces risk in its income and housing costs arising from risk in wage income, rents, and home prices. The household starts adult life at age 25, retires at age 65, and dies at age 95. It begins adulthood in a location whose wage, rent, and home price match those of median full-time working households in 2010: \$58,000 per year, \$13,000 per year, and \$186,000, respectively. Subsequent evolution of these prices is drawn from the observed location-specific paths in each of the 720 locations from 1940–2010, similar to the price paths in Figure 2 but for a 70-year period, the same length as the household's adult life. The household moves to a new location according to an exogenous schedule. Conditional on moving M years into adulthood, the baseline probability of moving to a particular location at that time is proportional to the corresponding CZ's population M years after 1940, though we consider moves to correlated markets as well.[15,16] Prices in locations other than the initial location follow

[15] Sinai and Souleles (2013) find that households tend to move to correlated housing markets. We adopt uncorrelated moves as a baseline because it tends to work against our conclusions, by both reducing the value of a renter's hedge against wage income risk (since uncorrelated moves diversify location-specific wage income risk to a greater extent) and increasing the value of an owner's hedge against wage income risk from the purchase price of their home (see the discussion on page 26). Because correlated moves are somewhat akin to non-moves from a risk perspective, Sinai and Souleles's (2013) finding leads us to focus our discussion somewhat more on the results for less-frequent moves.

[16] The assumption of exogenous moves is made for simplicity. In reality, endogenous moves can affect the risk costs of renting and owning. One effect arises from the fact that owning makes moving more costly, due to the significant transaction costs associated with home sales and purchases and non-transferable mortgages (e.g., Quigley, 1987; Fonseca and Liu, 2023; Liebersohn and Rothstein, 2023). This is a key way in which renters benefit more from the option to move: It is cheaper for renters to move in response to changing circumstances. An opposing effect is that the option to move can enable owners to selectively "lock in" capital gains (and avoid "locking in" losses). In principle, this effect could make a household better off owning than having full housing insurance. But as Sinai and Souleles (2013) note, that households tend to move to correlated markets both reduces such an option value and is suggestive

the observed price paths (without aggregate and idiosyncratic risk) in the 720 locations from 1940–2010 re-scaled so that the starting level is that observed in the 2010 cross-section.

In each of the 40 years of working life, from age 25 to 64, income equals median annual earnings in the household's current location. In each of the 30 years of retirement, from age 65 to age 95, income equals 50 percent of the household's mean annual earnings during working life. If it rents, the household pays the median rent in its location. If it owns, it purchases a home for the median home price in its location in the year before it arrives in the location and sells it in the year in which it leaves or dies. Ex ante, before the household knows how the prices in its initial location will evolve and in which locations it will live thereafter, it is exposed to risk in its net income stream from differences across locations in the evolution of wages, rents, and home prices during the 1940–2010 period. See Appendix A for more detail.

Consumption.— We use the following three models of consumption:

- Perfect within-lifetime consumption smoothing (PS): $c_{\omega,t} = \overline{c}_{\omega} \equiv \frac{Y_{\omega} - H_{\omega}}{\sum_{s=1}^{70}(1+r_t)^{-s}} \; \forall \; \omega, t,$
- Hand-to-mouth (HTM): $c_{\omega,t} = y_{\omega,t} - h_{\omega,t} \; \forall \; \omega, t,$
- Life cycle model (LC): $c_{\omega,t} = c^{\star}_{\omega,t} \; \forall \; \omega, t,$

where $c_{\omega,t}$ is consumption in state of the world ω and period t, and, as before, $y_{\omega,t}$ is labor income, $h_{\omega,t}$ is the net housing cost flow, and Y_{ω} and H_{ω} are the present values of lifetime income and housing costs in state ω, respectively. $c^{\star}_{\omega,t}$ denotes optimal consumption in the life cycle model.

The PS and HTM models share the same key strengths: transparency and flexibility in accommodating any possible model of the risk process. The PS model features perfect consumption smoothing over time because households have access to perfect credit markets. As such, by minimizing the welfare cost of a given amount of risk in the present value of net income, it should put a lower bound on the risk cost or benefit of different housing exposures. The HTM model is the opposite extreme in that it features no consumption smoothing over time. As such, it likely is a poor model of behavior in the face of short-term, transitory shocks, but a better model of behavior in the face of the kind of longer-term, more persistent risks that are crucial for lifetime welfare. It can be a useful benchmark in part because it likely overstates the welfare cost of fluctuations in

that moves tend to be driven by other factors. Moreover, moving costs, estimates of which tend to be quite high (e.g., Kennan and Walker, 2011), likely significantly limit the option value to move.

net income over time, so the associated risk costs from the HTM model likely are upper bounds. The LC model has the benefit of allowing for partial consumption smoothing within lifetime states of the world based on information revealed up to that point, at the cost of requiring strong assumptions for implementation, including about the underlying risk process. In the main text, we focus mostly on the PS model for simplicity and to be conservative, as it tends to understate the risk benefit of renting relative to owning (by understating both the risk benefit of renting and the risk cost of owning relative to full housing insurance).[17]

Utility.— We assume that the flow utility function is constant relative risk aversion, $u(c) = \frac{c^{1-\gamma}}{1-\gamma}$, with a baseline of $\gamma = 2$. We test robustness to alternatives. As before, the discount rate is constant and equal to the real interest rate, $r_t = 2.5\%$.

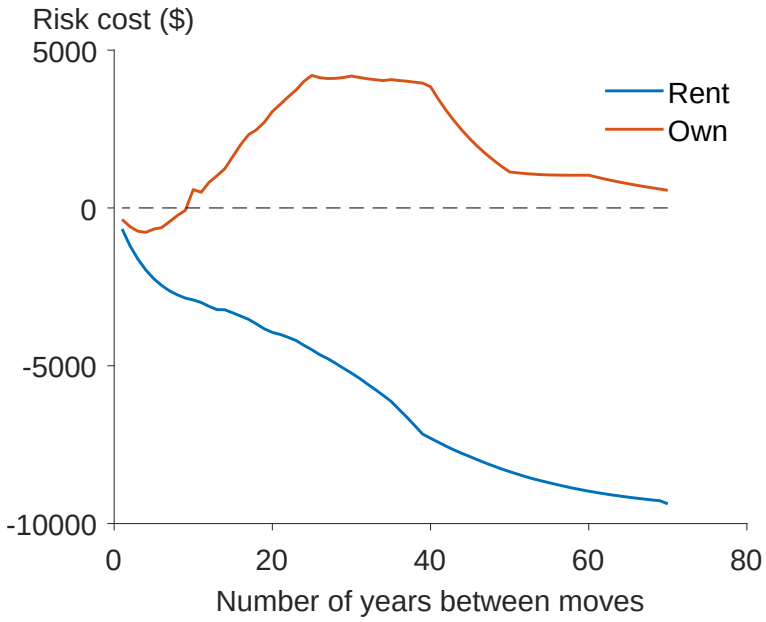
Risk cost measure.— We calculate the lifetime risk cost of always renting versus always owning for households that move exogenously at various frequencies. Our main measure of risk cost is the equivalent variation EV of going from the housing exposure of interest, $h_{\omega,t+s}$, to a counterfactual alternative in which housing costs are fully insured at actuarially fair prices at the beginning of adult life, $E\left(h_{\omega,t+s} \mid t\right)$ for all future periods s. We report the present value of the uniform increment to income, $PV(EV)$, in all time periods under the housing exposure of interest that makes the household as well off ex ante as it would be if housing costs were fully insured at actuarially fair prices (same expected utility at the beginning of adulthood):

$$v\left(\left\{y_{\omega,t+s} + EV\ ,\ h_{\omega,t+s}\right\}_{s=1}^{70}\right) = v\left(\left\{y_{\omega,t+s}\ ,\ E\left(h_{\omega,t+s} \mid t\right)\right\}_{s=1}^{70}\right). \tag{3}$$

v is the indirect utility or value function of the corresponding consumption model. The full-insurance counterfactual for renters involves the household paying at each date the average rent at that same date. So, in the extension where we allow for aggregate risk, average rent changes over time and the full-insurance counterfactual for renters features the same changes in the certain rent over time. This means that the only thing that changes between the renter's status quo exposure and their counterfactual full-insurance exposure is risk, not changes in expected rents over time. This applies to both the baseline without aggregate risk and to all extensions. For owners, at each date the household pays the same annuitized value of the expected present value of lifetime net housing costs. Hence, owners' housing costs are perfectly smoothed over the life cycle

[17] Were it not for the correlation between rents and wages, the PS model would exclude a potentially-important disadvantage of renting relative to full housing insurance in situations with binding borrowing constraints: Variation in rents over time can cause costly variation in consumption over time (Sinai, 2011). Given the strong correlation between rents and wages, however, variation in rents over time tends to reduce variation in net income over time by offsetting the variation in wages.

Figure 7: Risk Cost Relative to Full Housing Insurance as a Function of Time Between Moves



Notes. Lifetime risk costs of owning and renting, respectively, as a function of the number of years between moves during the life cycle. Risk cost is the equivalent variation of going from the housing exposure to full, actuarially fair housing insurance. See Section 5 for details. Risk only includes location-specific changes without aggregate and within-market idiosyncratic price changes. Source: Decennial Census and ACS.

in both the status quo and full-insurance counterfactual. We make this assumption, which tends to understate the utility cost of an owner's housing cost volatility, to be conservative with respect to the conclusion that owning tends to be riskier than renting. For both renters and owners, the expected present value of housing costs is the same under the risky and certain housing cost counterfactuals, so the equivalent variation of going from risky to certain housing costs is a measure of the pure risk cost of the risky housing exposure.

Results.— Figure 7 plots, for the PS consumption model, the risk costs of renting and owning, both relative to full housing insurance, as functions of the time between moves. The risk cost of owning is slightly negative for spells between moves of up to eight years, increases to around $4,000 for spells of 20–40 years, and decreases to around $1,000 for spells of 50 years or more.[18] By contrast, renting has a negative risk cost (i.e., positive

[18]Owning confers a small risk benefit over full housing insurance for short spells between uncorrelated moves because (i) we have excluded idiosyncratic risk, which is quite large for owners with short spells (Giacoletti, 2021), and (ii) with exogenous moves, the purchase price hedges risk in wages and in the future sale price in the new location. Although this hedge can be valuable with frequent uncorrelated moves, it is less valuable with moves to correlated markets (as discussed on page 26), which account for a large share of moves empirically (Sinai and Souleles, 2013). It is also less valuable than the hedge provided by renting. Whereas an owner's purchase price in the new location hedges risk in the wages that are *anticipated* at the

risk benefit) relative to full housing insurance regardless of moving frequency. This risk benefit increases roughly linearly in the time between moves as risk accumulates, reaching a maximum risk benefit over full housing insurance of over $9,000 if the household never moves. As a result, renting provides a substantial risk benefit over owning for all moving frequencies, increasing roughly linearly in the time between moves up to about 40 years before reaching a plateau at about $10,000 (shown in Figure 8's "wage beta = 1" curve).

The conclusion that renting is safer than both owning and full housing insurance is highly robust to a wide range of changes in the risk process, the model of consumption, and other changes to the model. The risk benefit of renting is increasing in the coefficient of relative risk aversion and decreasing in the extent of discounting. It is greater if we include idiosyncratic and aggregate risk, and it remains substantial even if we include considerable idiosyncratic rent risk while continuing to exclude idiosyncratic home price risk. It is highly robust to the treatment of retirement (both the timing and the replacement rate) and the modeling of owner's carrying costs.

5.2 Mechanisms

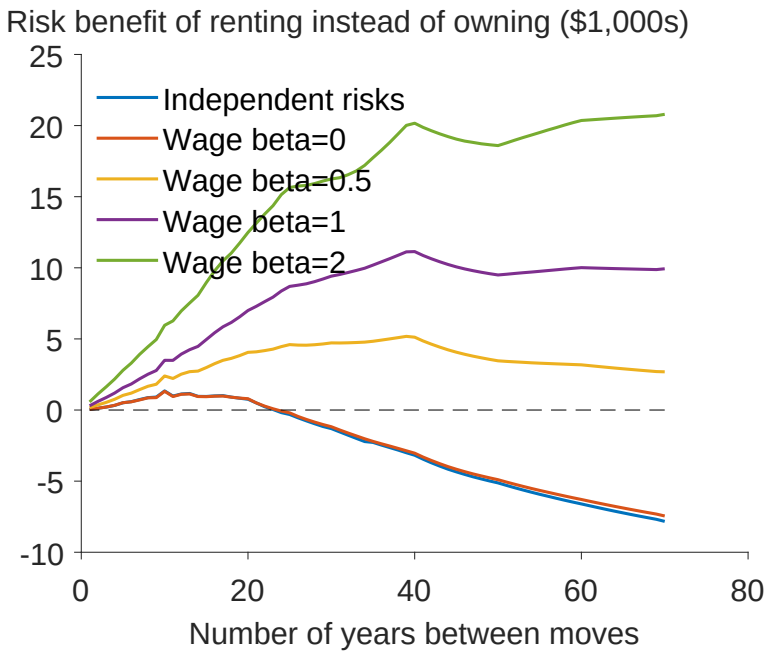
Proximate driver: Correlated wage risk.— Figure 8 plots, for the PS consumption model, the lifetime risk benefit of renting over owning as a function of the time between moves for different levels of wage income risk. The measure of wage risk is "wage beta" β_w, where the wage is $w_{\omega,t}(\beta_w) = \beta_w w_{\omega,t} + (1 - \beta_w)E(w_{\omega,t} \mid t)$. The wages of a household with a wage beta of two, for example, are on average the same as those of a typical household but twice as volatile around the mean. A beta of zero means the household faces no wage risk. A beta of one is the baseline risk model.[19]

The results reveal the crucial role of the interaction with wage risk. The net risk benefit of renting scales roughly linearly in wage beta for values from 0.5–2. So whereas with the baseline wage risk exposure, a household with spells between moves of 20 years or more enjoys a net risk benefit from renting of about $10,000, with a wage beta of two the benefit is about $20,000, and with a wage beta of 0.5 the benefit is about $3,000–$5,000. The interaction with wage risk transforms rent risk from being quite costly (welfare cost to a household that never moves with independent wage risk of $7,500) to quite valuable (welfare benefit of $9,400). This is because renting provides a highly valuable

time of the move, renting hedges (and an owner's sale price tends to exacerbate) risk in the wages that are actually *realized* ex post.

[19]Wage beta affects not only the extent to which wages covary with housing costs but also the level of wage volatility. An alternative would vary the correlation with housing costs while holding constant the level of wage volatility. In practice, we find that the risk costs are virtually identical either way.

Figure 8: Risk Benefit of Renting Over Owning as a Function of Wage Risk Exposure



Notes. Lifetime risk benefit of renting over owning as a function of the number of years between moves during the life cycle with different levels of exposure to wage income risk. The measure of exposure to wage risk is "wage beta" β_y, where the wage is $y_{\omega,t}(\beta_y) = \beta_y y_{\omega,t} + (1 - \beta_y)E(y_{\omega,t} \mid t)$. So the wages of a household with a wage beta of two are on average the same as those of a typical household but are twice as volatile around the mean. A beta of zero means the household faces no wage risk. A beta of one is the baseline risk model. The results with independent wage risk are almost identical to those without wage risk (wage beta of 0) (so they can be hard to distinguish). See Section 5 for details. Risk only includes location-specific changes without aggregate and within-market idiosyncratic price changes. Source: Decennial Census and ACS.

hedge against earnings risk (worth about $16,900) that dominates the risk cost of its contribution to cost-of-living risk (cost of about $7,500).

The hedge against wage income risk is more beneficial than the contribution to cost-of-living risk is costly because wage income risk is greater than cost-of-living risk. In terms of the economics of the second best, the modest hedge against wage income risk from renting is a modest decrease in a large wedge (tall, skinny trapezoid), whereas the partial effect of rent risk is a modest increase in a smaller wedge (somewhat wider, significantly shorter trapezoid;). This is in keeping with Fact 3 in the descriptive analysis that renting has a valuable portfolio effect from hedging income risk that dominates its partial effect of increasing cost-of-living risk, and so reduces net risk exposure even relative to full insurance against housing cost risk. So for a household with otherwise-hard-to-hedge exposure to wage income risk, exposure to rent risk is valuable, not costly. Indeed, while the rent risk exposure of a typical renter is quite valuable, in the baseline specification greater exposure to rent risk would be even better. For example, we estimate that for a household that moves once ten years into adulthood, the optimal exposure to rent risk

is about 80% greater than that of a typical renter (i.e., the optimal "rent beta" is 1.8). Rent risk becomes costly relative to full housing insurance for this household only for exposures at least four times that of a typical renter (rent beta of four).

Implications.— By transforming rent risk from costly to valuable, the interaction with wage risk also transforms important rules of thumb about the riskiness of renting versus owning. One is that the risk benefit of owning relative to renting increases in the time between moves, due to greater discounting of an owner's sale price and accumulation of a renter's rent risk (Ortalo-Magné and Rady, 2002; Sinai and Souleles, 2005). The interaction with wage risk is a powerful force in the opposite direction. With wage income risk, rent risk is a benefit, not a cost, of renting. This benefit accumulates as the time between moves increases.[20] Intuitively, with longer spells between moves, a household has greater exposure to location-specific risk in wage income; its location exposures are less diversified. This increases the value of a renter's hedge against wage income risk: Renting provides "location insurance."[21]

Another rule of thumb that the interaction with wage income risk transforms is that moves to positively-correlated housing markets decrease the risk cost of owning. The idea is that with correlated moves, owning in one location hedges risk in the housing costs one will face in future locations (Ortalo-Magné and Rady, 2002; Han, 2008; Sinai and Souleles, 2013). Here, too, the interaction with wage income risk is a powerful force in the opposite direction. With wage income risk, the initial purchase price in a new location hedges risk in wage income in that location (as discussed in footnote 18). Concretely, if circumstances lead one to move to a location with low wages, a silver lining is that on average home prices (and rents) will be low as well. Correlated moves diminish this hedge by making the net outlay from selling one's current home and buying one's next home smaller on average and less variable. We find that going from uncorrelated to correlated moves (including "perfectly correlated" moves to the location whose home price is closest to that in the current location) tends to increase the risk cost of owning, by reducing its hedge against wage income risk from the initial purchase price, and increase the risk benefit of renting, by increasing the value of its hedge against the (greater due

[20] Opposing this is the well-known force that a longer horizon eventually decreases an owner's sale price risk. This force is understated by our baseline model since it excludes idiosyncratic and aggregate risk (which we do to be conservative, since including them reinforces our conclusions). Realistic idiosyncratic risk significantly increases the risk of owning for short spells and so leads to non-monotonicity where the risk benefit of renting over owning decreases over a small range of horizons while increasing elsewhere (where the growing value of rent risk dominates). See Appendix XXX.

[21] That the risk benefit of renting over full housing insurance grows roughly linearly in the time between moves is another manifestation, in addition to the results on wage beta, of the risk benefit of renting being roughly proportional to the household's exposure to wage income risk over a wide range.

to being less-diversified) wage income risk.[22,23]

Yet while the interaction with wage income risk is a key driver of the risk benefit of renting, even with no such interaction, owning would confer a sizeable risk benefit over renting only for households with long spells between moves. For example, for a household without any wage income risk exposure, owning is riskier than renting with spells between moves of up to 16 years, and confers a lifetime net risk benefit of at least $3,000 only if spells exceed 30 years. With correlated moves or idiosyncratic or aggregate risk, the circumstances in which owning confers a risk benefit over renting are rarer still.

Summary.— The key proximate driver of the conclusion that for working households renting is safer than owning is the large positive covariance of rents and wages. Because of this covariance, renting hedges wage risk. Were it not for this hedge, a renter's housing exposure would have a straightforward effect on risk relative to full housing insurance: It would increase risk due to its contribution to cost-of-living risk. But because of this hedge, a renter's housing exposure effectively bundles two assets: one that pays off more when the cost of living is lower, which exacerbates cost-of-living risk, and one that pays off more when wages are lower, which hedges wage risk. The overall effect on risk exposure therefore reflects opposing pro- and anti-insurance effects. In principle, either could dominate. In practice, the insurance of wage risk dominates, so renting decreases risk exposure even relative to full housing insurance. As a result, for working households renting tends to be safer than owning regardless of the time horizon in the home or location.

Ultimate driver: Local labor demand shocks.— Why are location-specific changes in housing prices and wages so strongly positively correlated? Although this likely reflects multiple factors and a full analysis is beyond the scope of this paper, a variety of evidence

[22]Compared to uncorrelated moves, "perfectly correlated" moves increase the estimated risk benefit of renting over owning by 12% on average across times between moves. That a move to a correlated market is in a sense similar to no move at all reveals the connection between these modifications to different rules of thumb. Compared to a move to an uncorrelated market, a move to a correlated market is similar to an increase in the time horizon in the original market, which increases the value of a renter's wage risk hedge and decreases the risk consequences of an owner's sale and purchase at the time of the move.

[23]A variety of evidence is consistent with many households viewing rent risk as costly rather than beneficial and using these rules of thumb in their housing tenure choices (e.g., Sinai and Souleles, 2005; Han, 2010; Sinai and Souleles, 2013). Part of the explanation may be heterogeneity: Rent risk *is* costly to households with minimal exposure to wage income risk, and a lengthening of the effective horizon, including from correlated moves, *does* reduce the riskiness of owning relative to renting for certain households. Another part of the explanation may be that some households misperceive the riskiness of exposure to rents or of owning versus renting more generally, perhaps in part by neglecting the interaction with wage income risk. Such misperceptions would be in keeping with survey evidence from Adelino et al. (2018) that 71% of U.S. households view housing as a "safe" investment, despite the considerable volatility of home prices and the high frequency with which many households move. We return to this issue in Section 6.

suggests that a key factor is the combination of (i) significant and persistent location-specific shocks to labor demand, (ii) relatively inelastic supply of housing and other location-specific capital in many locations, and (iii) spatial equilibrium forces toward different locations being similarly attractive to potential residents.

The large, persistent differences in wage growth across locations and time periods are suggestive that labor productivity risk is substantial, especially at the local level. This interpretation is consistent with a large literature on the local labor market effects of shocks to the demand for tradable goods (e.g., Barro and Sala-i Martin, 1991; Blanchard and Katz, 1992; Autor et al., 2013; Giannone, 2018; Howard and Liebersohn, 2018; Hornbeck and Moretti, 2019). An important finding of this literature is that in recent decades there has been little if any tendency for convergence in nominal wages across locations (e.g., see Moretti, 2011, for a review). Such persistence is a key driver of our finding, which builds on the findings of this literature, that location-specific shocks can have a large effect on lifetime earnings.

Labor demand shocks affect housing prices both by affecting the cost of producing housing (since labor is an input to housing production) and, likely of greater importance in explaining the relationship between location-specific changes in wages and housing prices, by shifting the demand for housing. The strength of the association between location-specific changes in wages and housing prices suggests that wage changes shift the demand for housing substantially and that the supply of housing is relatively inelastic. The inelasticity of housing supply in response to negative demand shocks is a natural consequence of the durability of housing (Glaeser and Gyourko, 2005). The inelasticity in response to positive demand shocks in the locations with the greatest wage growth over this period is largely a consequence of the increased difficulty of building in these locations between the 1960s and the 1990s (Glaeser and Gyourko, 2018).

The apparently large impact of wage changes on the demand for housing likely primarily reflects a combination of income effects (wage growth enriches working households, which increases their demand for normal goods, including housing) and, especially, spatial equilibrium-type responses to location-specific wage changes. Locations with high wages tend to have high housing costs, and the positive relationship between wages and housing costs is even stronger when isolating location-specific changes in these prices relative to the national average. This is consistent with spatial equilibrium forces ensuring that, other things equal, places that experience faster growth in wages experience faster growth in the prices of non-tradables like housing. The strong positive correlation between the growth of a location's wages and rents relative to the national average provides additional support for the conclusion that local labor demand risk is the major

driver of location-specific risk, since risk in local consumer amenities, by contrast, would tend to lead to rents covarying negatively with wages (Roback, 1982).

The combination of significant local labor demand shocks and relatively inelastic supply of housing generates considerable location-specific risk in wages, which in turn translates into significant risk in lifetime earnings. An obvious way to hedge this "location risk" is to move frequently, which diversifies one's location exposures over the life cycle. But the large costs of moving make such a strategy extremely costly (e.g., Kennan and Walker, 2011). An alternative way to hedge location risk that avoids such costs is to rent one's housing. The strong positive relationship between growth in wages and rents across locations means that renting provides a powerful hedge against location-specific risk in wages that does not require frequent moving.

6 Implications

Is renting better than owning? *If* the key difference between renting and owning were the effect on risk, then renting would be significantly better than owning for most households. Our estimates suggest that for a typical household, the lifetime risk benefit of renting over owning is on the order of $10,000. But risk is not the only difference between renting and owning. Other differences include any direct preference for owning, moral hazard and other transaction costs between renters and landlords (including the risk of non-renewal of a rental contract), and considerations related to mortgages and owned housing as a store of wealth and saving commitment. Given the many differences between renting and owning, that renting is safer than owning for many households just means that one important component of the overall welfare effect of owner-occupied housing is not the benefit commonly thought but a cost.

How generalizable are the results? The analysis highlights the importance of two factors that likely vary significantly across households both within and across economies: exposure to wage risk and time horizon in a home or location. The smaller the exposure to wage risk, the smaller the risk benefit from renting. Still, our results suggest that, other things equal, exposure to wage risk would have to be significantly smaller than that of a typical working-age household in the U.S. to make owning safer than renting, and even then its risk advantage would tend to be small. The key factor is that wage risk is a major risk—for many households the biggest risk they face—so even a modest hedge against it can be quite valuable and a modest exacerbation of it can be quite costly.

The time horizon in a home or location is also an important determinant of the risk cost of renting versus owning, though it has a more subtle effect than previously thought. In

addition to the well-known effect due to time discounting that a longer horizon eventually reduces the sale price risk of owning, there is an opposing effect due to the interaction with wage risk that a longer horizon increases the value of the wage risk hedge from renting. Our results suggest that the net effect of these and other forces is such that renting is safer than owning for working households regardless of the horizon and for retired households with horizons of up to roughly 25–30 years or so.

Finally, while we mostly focus on location-specific risk in this paper because it allows us to deal with many concerns one might have when dealing with aggregate income risk (e.g., changes in interest rates), in Appendix A.3 we document similarly strong relationships for aggregate income and housing cost changes, both with national-level U.S. data commonly used in macroeconomics and with the cross-country panel data recently compiled by Jordà et al. (2019). Hence, our results also apply to aggregate risk, with potentially important implications for asset pricing that remain to be explored.

What are the policy implications? Our results have implications for two main sets of policies. The first is the many policies that restrict the supply of housing, especially in booming coastal locations (e.g., Glaeser and Gyourko, 2018). Such restrictions, by reducing the elasticity of housing supply, increase the extent to which labor demand shocks capitalize into wages and housing costs, which increases risk in earnings and housing costs. Such restrictions also increase the risk consequences of housing tenure choices, by increasing both the value of a renter's hedge against wage income risk and the cost of an owner's exposure to home price risk. Our results suggest that the risk cost of such policies could be substantial. Better understanding these costs is an important topic for future research.

The second set of policies our results have implications for are the many policies that implicitly subsidize owner-occupied relative to rental housing. Although a full welfare analysis of these policies is beyond the scope of this paper, here we report simple calculations aimed at providing a rough sense of the likely magnitudes. The key idea is that any tendency of households to underestimate the risk (or other costs) of renting relative to owning magnifies the efficiency cost of policies that explicitly or implicitly subsidize owning relative to renting beyond what would internalize any externalities.

A variety of evidence suggests that many households underestimate the private costs of owning relative to renting, and that these misperceptions affect choices about whether to own or rent. Those who teach economics are likely familiar with the widely-held view that renting is "throwing money away" in a way that owning is not, which reflects a failure to understand the opportunity cost of living in a home you own (the foregone

opportunity to rent it to someone else). More germane to our findings, many people appear to underestimate the relative riskiness of owning. In a nationally representative survey of more than 50,000 households, Adelino et al. (2018) find that 71% of U.S. households believe that housing is a "safe" investment, significantly higher than the 55% who believe the same about government and corporate bonds. Moreover, both current housing tenure status and future intentions to buy versus rent are strongly correlated with perceptions of house price risk. In light of our finding that for many households owning is riskier than renting, these survey results suggest that many people are underestimating the riskiness of owning relative to renting and that these misperceptions affect choices about whether to own or rent. If so, such misperceptions tend to reduce welfare both directly, by leading people to make choices that are privately sub-optimal, and indirectly, by magnifying the efficiency cost of policies that subsidize owner-occupied relative to rental housing.

A simple measure of the marginal excess burden of the net subsidy on owner-occupied housing (Harberger 1964; Auerbach 1985; Auerbach and Hines Jr 2002) is:[24]

$$MEB_{sub} = \left(\frac{\sigma_{sub}}{2} + \sigma_{int}\right) \times \Delta Q_{sub}, \tag{4}$$

where σ_{int} is the internality wedge (i.e., the subsidy that would increase the quantity of owner-occupied housing as much as misperceptions do), σ_{sub} is the implicit net subsidy on owner-occupied relative to rental housing beyond that justified by externalities, and ΔQ_{sub} is the increase in the quantity of owner-occupied housing from this implicit net subsidy.[25] As is well known, pre-existing distortions magnify the marginal excess burden of additional distortions that reinforce them; the marginal excess burden is not the usual triangle but the potentially much larger trapezoid. That many people appear to underestimate the cost of owner-occupied relative to rental housing magnifies the marginal excess burden of implicit housing subsidies. (Similarly, the subsidies increase the private welfare costs of misperceptions.)

Internality wedge σ_{int}.— As a conservative benchmark, we ignore any bias from failures to appreciate the opportunity cost of owning and suppose that the only bias is that people underestimate the riskiness of owning relative to renting. We suppose that people understand that owning, like renting, is risky, but fail to account for the correlations between wages and housing costs. In that case, our results suggest that a household that

[24] See Appendix Figure A.17 for a graphical illustration.

[25] Equation (4) is the marginal excess burden if there are no spillovers to other distorted markets and if the difference between the social marginal cost and social marginal benefit of owner-occupied housing is linear in the quantity of owner-occupied housing (as would be the case if both of those schedules are linear). If these schedules are non-linear, then equation (4) is a second-order approximation to the marginal excess burden.

never moved would underestimate the *lifetime* (risk) cost of owning relative to renting by more than \$16,000 (the difference between "wage beta = 1" and "independent risk" in Figure 8): $\sigma_{int} = \$16,000$.

Net subsidy σ_{sub}.— A variety of policies subsidize owner-occupied relative to rental housing. The effective subsidy depends on many particulars of a household's situation, including its marginal income tax rate, whether it itemizes deductions, and the extent to which it will be a net seller of housing before death. As a rough approximation, we assume that for a typical household the net subsidy beyond what would internalize externalities is the product of its marginal income tax rate τ and its (possibly implicit) rent R, $\sigma_{sub} = \tau \times R$. This is a rough approximation to what is likely the most important driver of the net subsidy for owner-occupied housing: that the implicit rent of owners is untaxed, whereas the explicit rental income of landlords is taxed.[26] We assume a benchmark marginal tax rate on income of 25 percent and an annual rent of \$14,300, which yields an annual subsidy on owner-occupied housing of roughly \$3,600 per year.[27] We assume that the net subsidy *per lifetime* of living in owner-occupied housing (to match the units of σ_{int}) is \$95,400, the value of a 50-year annuity of \$3,600 per year with a 2.5% annual discount rate: $\sigma_{sub} = \$95,400$.

Quantity response ΔQ_{sub}.— For our benchmark quantity response, we assume that the subsidy increases the fraction of its lifetime that a typical household owns by 28.2 percentage points, roughly the difference in the homeownership rate between the U.S. and Switzerland, which has small homeowner subsidies (64.6% and 36.4% in 2019; U.S. Census Bureau 1980–2021; Swiss Federal Statistical Office 2000–2020): $\Delta Q_{sub} = 0.282$.

Quantity response ΔQ_{int}.— Assuming linear demand for owner-occupied housing, this price responsiveness implies that misperceptions increase the fraction of time the typical household spends owning by about $\Delta Q_{int} = \Delta Q_{sub} \times \frac{\sigma_{int}}{\sigma_{sub}} = 4.7$ percentage points of its lifetime: $\Delta Q_{int} = 0.047$. (For robustness, we also consider quantity responses half as large.)

[26] The "tax expenditure" associated with the lack of taxation of owner occupiers' implicit rent totaled \$126b in current dollars in 2020 (Office of Tax Analysis 2020), making it the second-largest tax expenditure after that for employer contributions for health insurance. Many other policies also implicitly subsidize owner-occupied relative to rental housing, including implicit taxes on high-density housing (and so subsidies on low-density housing, which is more likely to be owner-occupied) such as zoning restrictions and government-provided infrastructure. While it is impossible to capture the full richness of the tax treatment of housing in a simple calculation, our aim is to provide rough, back-of-the-envelope estimates of the likely magnitude for a typical household.

[27] This annual rent comes from applying the average net annual rental yield on housing of 4.2 percent (Demers and Eisfeldt, 2022) to the median home price in 2020 of \$340,000 (U.S. Department of Housing and Urban Development, 1963–2021). This estimate of the annual subsidy is broadly in keeping with those of Gyourko and Sinai (2003), who estimate per-owner tax subsidies on the order of \$1,000–\$6,000.

Table 3: Efficiency Costs of Ownership Subsidies and Housing Risk Misperceptions

Panel A: Inputs for MEB formula

Implicit subsidies ($s per lifetime)		Quantity response (in lifetimes)		
			η^{main}	η^{low}
Misperceptions, σ_{int}	16,000			
Net subsidy, σ_{sub}	95,400	$\Delta Q_{sub} = \eta\sigma_{sub}$	0.282	0.141
		$\Delta Q_{int} = \eta\sigma_{int}$	0.047	0.024

Panel B: Efficiency cost of excess homeownership subsidies ($s per lifetime)

	With misperception		Without misperception		Effect of misperception	
Formula	η^{low}	η^{main}	η^{main}	η^{low}	η^{main}	η^{low}
$\left(\frac{\sigma_{sub}}{2} + \sigma_{int}\right)\Delta Q_{sub}$	17,963	8,982	13,451	6,726	4,512	2,256

Panel C: Efficiency cost of misperceptions on top of excess subsidies ($s per lifetime)

	With subsidies		Without subsidies		Effect of subsidies	
Formula	η^{main}	η^{low}	η^{main}	η^{low}	η^{main}	η^{low}
$\left(\frac{\sigma_{int}}{2} + \sigma_{sub}\right)\Delta Q_{int}$	4,890	2,445	378	189	4,512	2,256

Notes. σ_{int} is the internality wedge, i.e., the subsidy that would increase the quantity of owner-occupied housing by as much as misperceptions would under the assumptions discussed in Section 6. σ_{sub} is the implicit net subsidy on owner-occupied relative to rental housing beyond that justified by externalities. η^{main} (η^{low}) is the slope of demand for owner-occupied housing with respective to the subsidy rate assuming the baseline (low) price responsiveness. ΔQ_{int} (ΔQ_{sub}) is the increase in owner-occupied housing as a result of misperceptions (subsidies).

Results.— Table 3 shows the results. The estimates suggest that misperceptions increase the efficiency cost of ownership subsidies by about one-third, roughly from $13,000 to $18,000 per household over its lifetime.[28] Misperceptions themselves have a welfare cost of over $4,500 per household over its lifetime. These calculations, despite their many caveats, are suggestive that misperceptions about the costs of renting relative to owning have large costs, both direct, by leading people to make choices that are privately suboptimal, and indirect, by magnifying the efficiency cost of policies that subsidize owner-occupied relative to rental housing beyond the associated externalities.

7 Conclusion

For many U.S. households, renting is safer than owning, often significantly so. The key driver is wage income risk. Because of the strong positive correlations between changes in housing prices and changes in wages, renting hedges wage income risk and owning

[28]While the absolute size of the efficiency cost depends on (is proportional to) the quantity response ΔQ_{sub} (the most uncertain ingredient), the implicit subsidies alone determine the proportional increase due to misperceptions: $MEB_{sub}/MEB_{sub|\sigma_{int}=0} = 1 + 2\sigma_{int}/\sigma_{sub} \approx 1.34$.

exacerbates it. The interaction with wage income risk is so strong, in fact, that for many working households renting would reduce risk even relative to full housing insurance. The greater risk exposure from owning is an important cost of owning and the many major policies that encourage it.

Our findings have implications for economic research and policy. In terms of research, the strong correlations between wages and housing prices mean that analyses of earnings risk should account for housing cost risk and vice versa. Failing to account for these linkages could lead to misleading conclusions about the costs of different risk exposures and the performance of different models. In terms of policy, our results, together with evidence that many households underestimate the relative riskiness of owning, suggest that the efficiency cost of many housing-related policies are larger than previously appreciated. Our results also suggest that the many policies that restrict the supply of housing could have large costs in terms of increasing risk in earnings and housing costs. Better understanding these issues is a high priority for future research.

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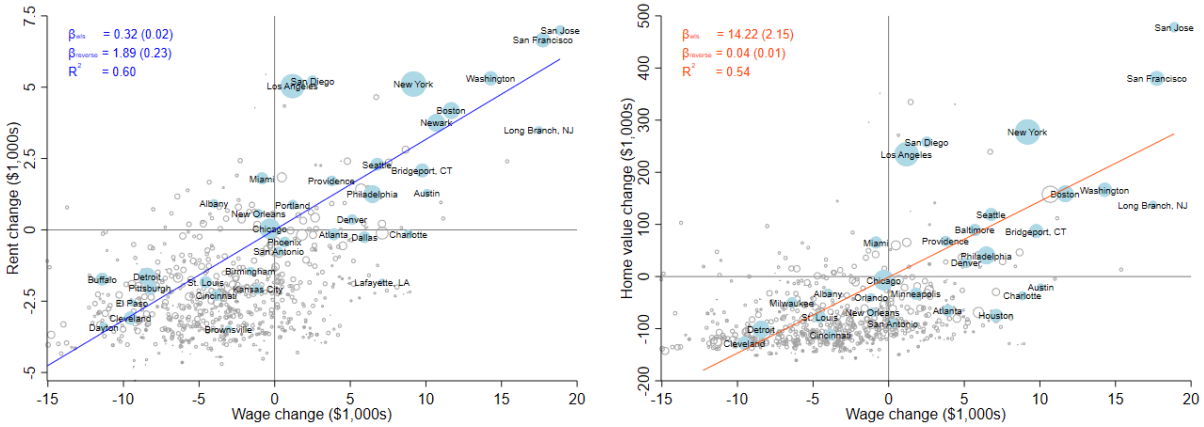
Copy of Figures and Tables from Main Text

Figure 1: Relationships Between Wages, Rents, and Home Prices, 1960–2010

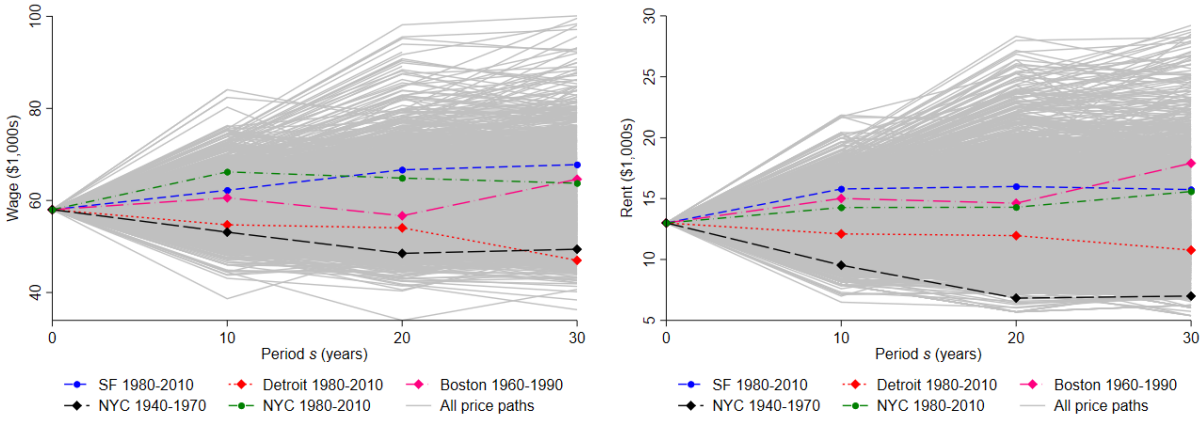


(a) rent vs. wage changes

(b) home price vs. wage changes

Notes. Changes in average location-level full-time wages, rents, and home values from 1960–2010 by Commuting Zone (CZ) relative to national-level changes. All prices are quality-adjusted, using the location-year fixed effects of the hedonic regression of prices on household and home characteristics as described in the text and shown in equation A.4 in Appendix A. Circle areas are proportional to the location's population share. β_{wls} and R^2 are obtained from regressing changes in housing costs on changes in wages, weighted by population. $\beta_{reverse}$ is obtained from the corresponding population-weighted reverse regression of wage changes on housing cost changes. Robust standard errors in brackets are clustered by state. Prices are adjusted to dollars of 2020 using the Consumer Price Index (CPI) for All Urban Consumers. Source: Decennial Census and ACS.

Figure 2: Non-parametric Model of Wage and Housing Cost Risk

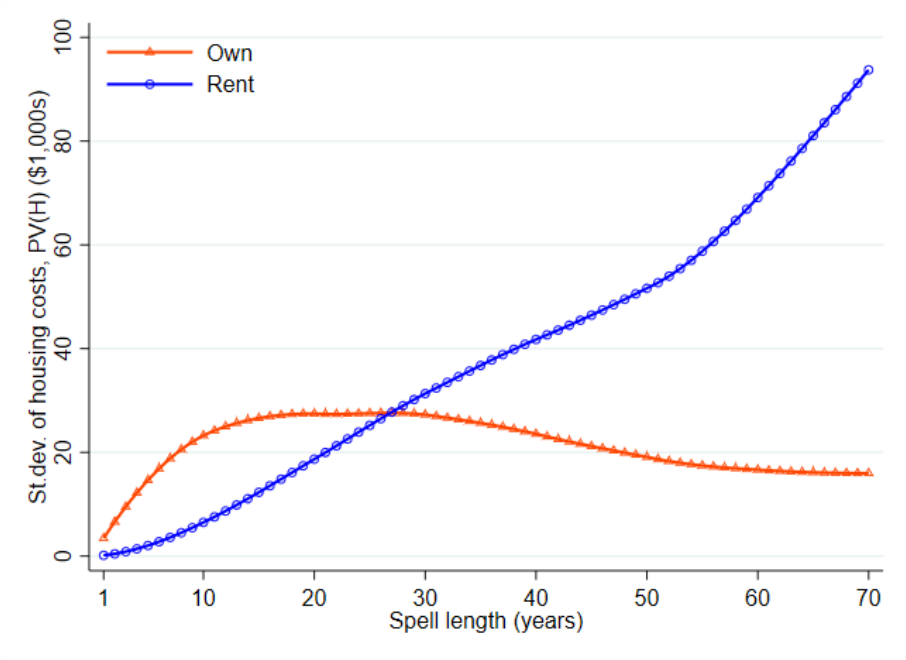


(a) wage paths

(b) rent paths

Notes. Illustration of the non-parametric model of price risks showing 30-year price paths of wages and rents starting from the same initial price at time t, \$58,000 for wages and \$13,000 for rents. Price paths only contain location-specific changes (corresponding to price component $X^{loc}_{i,t+s}$ in Appendix A) and no aggregate or within-market idiosyncratic risk. We omit values that are 120% above the 99th percentile from the figure (but not from the analysis) to improve its appearance. Source: Decennial Census and ACS.

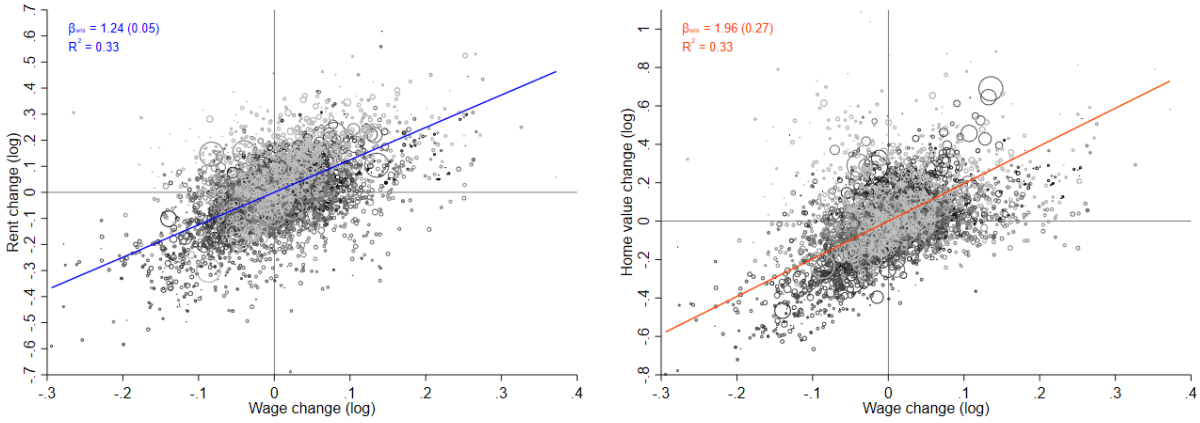
Figure 3: Volatility in Housing Costs by Length of Spell



Notes. Standard deviation of the present value of housing costs by the length of the spell in the home. Rents and home prices begin at those of a median full-time working household in 2010: $13,000 per year and $186,000, respectively (corresponding to price component $X^{loc}_{i,t+s}$ in Appendix A). Each then grows according to the location-specific path observed in different location-by-time-periods from 1940–2010. The resulting risk process, discussed in more detail in Appendix A, isolates across-location heterogeneity in the evolution of rents and home prices over time. In excluding aggregate and within-market idiosyncratic risk, it tends to understate the riskiness of owning relative to renting. Rents and home prices are quality-adjusted as described in the text before calculating present values and expressed in dollars of 2020 using the Consumer Price Index (CPI) for All Urban Consumers. The discount rate is 2.5%. Source: Decennial Census and ACS.

Figure 4: Relationships Between Wage, Rent, and Home Price Growth

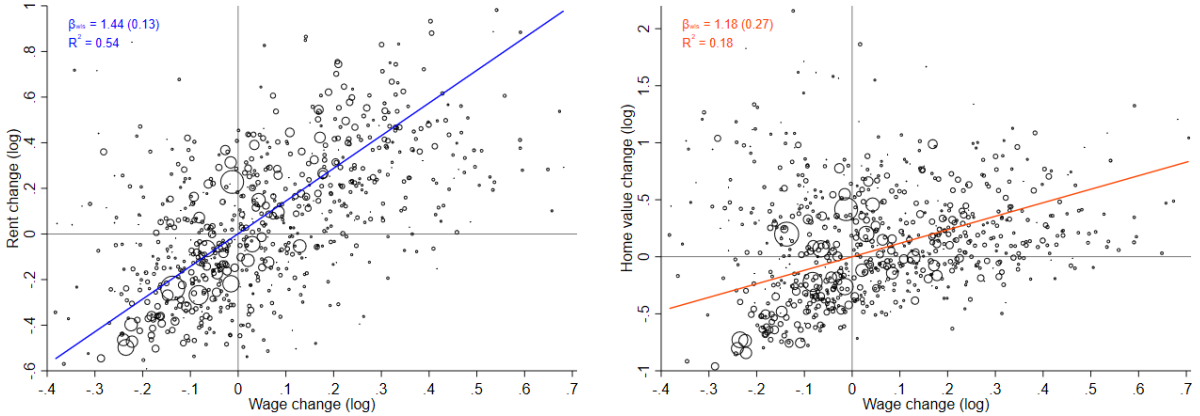
A. 10-year growth



(a) rent vs. wage growth

(b) home price vs. wage growth

C. 70-year growth

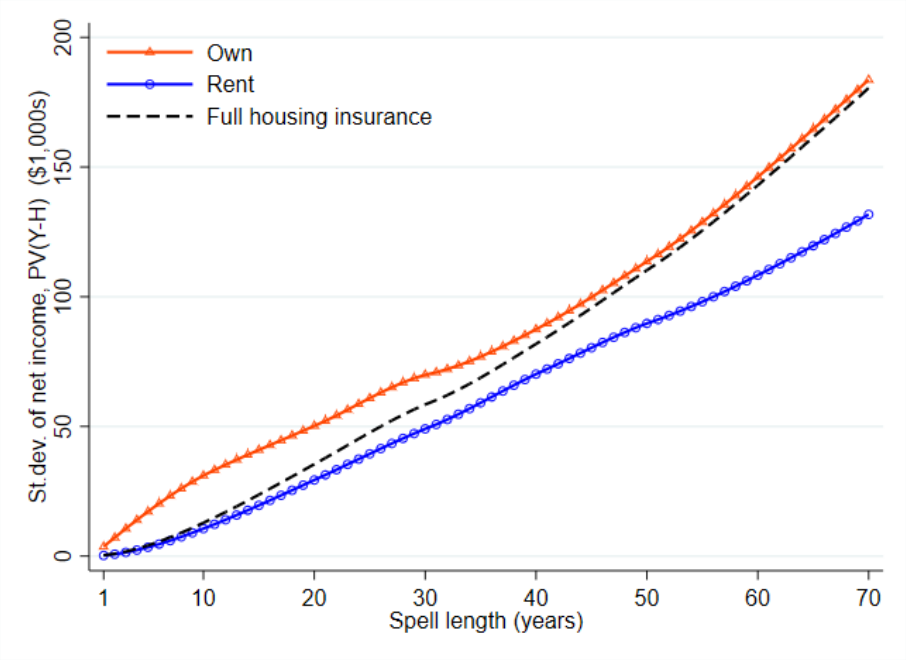


(c) rent vs. wage growth

(d) home price vs. wage growth

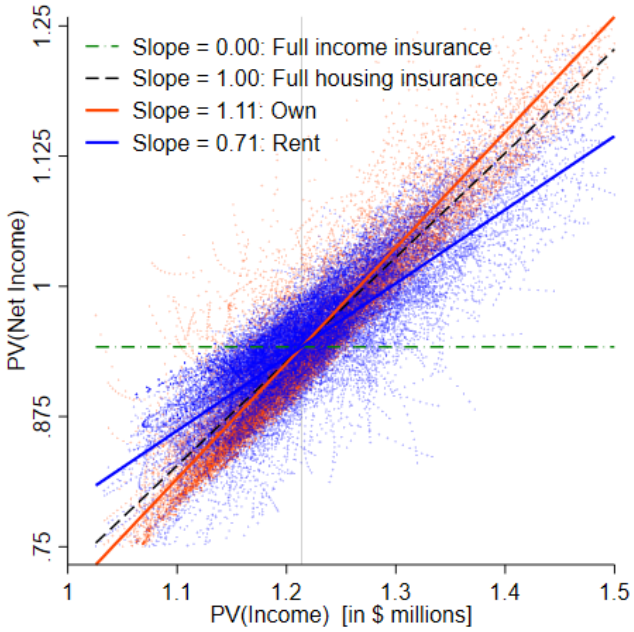
Notes. Changes in average location-level full-time log wages, log rents, and log home values from 1940–2010 by Commuting Zone (CZ) relative to national-level log changes (corresponding to $\Delta_s \ln x^{loc}_{c,t+s}$ with $s = 10$ and 70 in Appendix A). All prices are quality-adjusted, using the location-year fixed effects of the hedonic regression of log prices on household and home characteristics as described in the text. Panel A pools 10-year log changes (5,040 observation from 720 CZs times 7 decades); panel B shows the single cross-section of 70-year log changes across 720 CZs. Circle areas are proportional to the location's population share. Observations in darker shades are from more recent decades. β_{wls} and R^2 are obtained from regressing log changes in housing costs on log changes in wages, weighted by population. Robust standard errors in brackets are clustered by state. Prices are adjusted to dollars of 2020 using the Consumer Price Index (CPI) for All Urban Consumers. Source: Decennial Census and ACS.

Figure 5: Risk in Net Income by Length of Spell



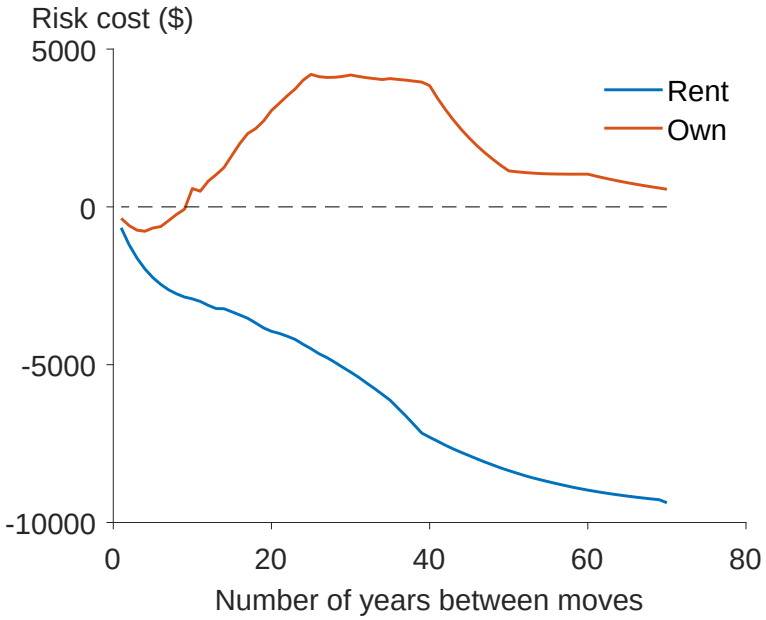
Notes. Standard deviation of the present value of net income by the length of the spell in the home. Net income is full-time wage income net of housing costs. Prices begin at those of a median full-time working household in 2010 ($58,000 wage income per year, $13,000 rent expenses per year, and a home price of $186,000, corresponding to price component $X_{i,t+s}^{loc}$ in Appendix A) and then grow according to the location-specific path observed in different location-by-time-periods from 1940–2010. The resulting risk process, discussed in more detail in Appendix A, isolates across-location heterogeneity in the evolution of rents and home prices over time. In excluding aggregate and within-market idiosyncratic risk, it tends to understate the riskiness of owning relative to renting. The discount rate is 2.5%. Source: Decennial Census and ACS.

Figure 6: Translation of Income Risk into Net Income Risk by Housing Position



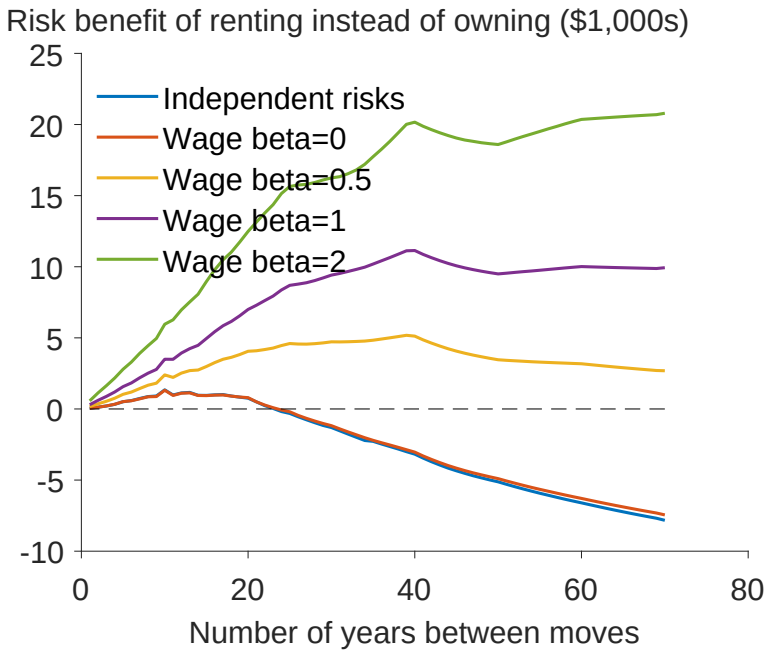
Notes. Present value of net income (income net of housing costs) as a function of the present value of gross (wage) income for a household that lives and works in the same location for thirty years under three scenarios: renting, owning, and having full, actuarially fair housing insurance. An observation is a CZ-by-30-year present value pair (corresponding to price component $X_{i,t+s}^{loc}$ in Appendix A). Wages, rents, and home prices are quality-adjusted as described in the text before calculating present values. Prices begin at that of a median full-time working household in 2010 ($58,000 per year for wages, $13,000 for rents, and $186,000 for home prices) and then grow according to the location-specific path observed in different location-by-time-periods from 1940–2010. Price paths exclude aggregate and within-market idiosyncratic risk. Regression slopes are from a population-weighted regression of 30-year present values of net income on 30-year present values of gross income. Source: Decennial Census and ACS.

Figure 7: Risk Cost Relative to Full Housing Insurance as a Function of Time Between Moves



Notes. Lifetime risk costs of owning and renting, respectively, as a function of the number of years between moves during the life cycle. Risk cost is the equivalent variation of going from the housing exposure to full, actuarially fair housing insurance. See Section 5 for details. Risk only includes location-specific changes without aggregate and within-market idiosyncratic price changes. Source: Decennial Census and ACS.

Figure 8: Risk Benefit of Renting Over Owning as a Function of Wage Risk Exposure



Notes. Lifetime risk benefit of renting over owning as a function of the number of years between moves during the life cycle with different levels of exposure to wage income risk. The measure of exposure to wage income risk is "wage beta" β_y, where the wage is $y_{\omega,t}(\beta_y) = \beta_y y_{\omega,t} + (1-\beta_y)E(y_{\omega,t} \mid t)$. So the wages of a household with a wage beta of two are on average the same as those of a typical household but are twice as volatile around the mean. A beta of zero means the household faces no wage risk. A beta of one is the baseline risk model. The results with independent wage risk are almost identical to those without wage risk (wage beta of 0) (so they can be hard to distinguish). See Section 5 for details. Risk only includes location-specific changes without aggregate and within-market idiosyncratic price changes. Source: Decennial Census and ACS.

Table 1: Volatility of Location-Specific Housing Cost and Income Changes

	Spell length, T (years):						
	10	20	30	40	50	60	70
Panel A: Standard deviation of changes ($1,000)							
Rent, $R_{\omega,t+T} - R_t$	1.6	2.4	2.8	3.0	3.3	3.7	4.2
Home price, $P_{\omega,t+T} - P_t$	37.8	51.9	64.5	70.2	72.0	77.5	90.0
Income, $y_{\omega,t+T} - y_t$	3.2	4.6	5.7	6.6	7.4	8.5	9.7
Panel B: Standard deviation of annualized log changes (%)							
Rent, $(\ln R_{\omega,t+T} - \ln R_t)/\sqrt{T}$	3.8%	4.1%	3.8%	3.6%	3.5%	3.5%	3.7%
Home price, $(\ln P_{\omega,t+T} - \ln P_t)/\sqrt{T}$	6.0%	5.4%	5.5%	5.4%	5.0%	4.8%	5.2%
Income, $(\ln y_{\omega,t+T} - \ln y_t)/\sqrt{T}$	1.7%	1.7%	1.7%	1.7%	1.7%	1.8%	1.9%
Panel C: Standard deviation of present values ($1,000)							
Prices:							
Rent, $R_\omega = \sum_{s=1}^{T} \frac{R_{\omega,t+s}}{(1+r_t)^s}$	6.5	18.7	31.3	41.8	51.6	69.1	93.7
Own, $O_\omega = P_t - \frac{P_{\omega,t+T}}{(1+r_t)^T}$	23.3	27.5	27.3	23.6	19.1	16.7	16.0
Income, $Y_\omega = \sum_{s=1}^{T} \frac{y_{\omega,t+s}}{(1+r_t)^s}$	12.8	35.5	58.4	81.8	110.2	143.0	180.5
Net Income:							
Renter, $Y_\omega - R_\omega$	10.6	29.3	49.1	70.2	89.7	108.3	131.7
Owner, $Y_\omega - O_\omega$	31.1	50.2	69.9	87.4	113.6	146.2	183.6
Full housing insurance, $Y_\omega - E(R_\omega)$	12.8	35.5	58.4	81.8	110.2	143.0	180.5

Notes. The table shows the standard deviation of (log) changes of location-specific wages, rents, and home prices and present values (PVs) of renters and owners housing costs and net income for different spell length. Wages and housing costs are quality-adjusted as described in the text. All changes are relative to the same initial price ($R_t = \$13k$, $P_t = \$186k$, $y_t = \$58k$) and the real discount rate used for computing PVs is the average post-war yield on long-run Government bonds, $r_t = 2.5\%$. Panel B shows changes in quality-adjusted log prices (location-year fixed effects $\hat{\lambda}_{ct}$ of the hedonic regression, equation (A.3) in the Appendix), normalized by the square-root of the spell length to make them comparable (see Appendix Figure A.2). If log price changes are permanent (i.e., follow a random walk), then the standard deviation of annualized log changes are constant irrespective of spell length. Panel A converts these quality-adjusted growth rates to dollar changes (corresponding to price component X_{it}^{loc} in equation (A.6) of the appendix). Panels A and B pool decadal data from Census years 1940 to 2010 so that the standard deviation of 10-year changes is based on 5,040 CZ-year observations (720 CZs times 7 10-year changes 1940–1950 to 2000–2010), 2,880 observations for 40-year changes, and 720 observations for 70-year changes. Panel C interpolates between census years to annual data spanning 1940 to 2010. The standard deviation of 10-year PVs is based on 43,920 CZ-year observations (720 CZs times 61 PVs for 1941–1950 to 2001–2010), 22,320 observations for 40-year PVs, and 720 observations for 70-year PVs. All calculations use population weights for the 720 CZs. Source: Decennial Census and ACS.

Table 2: Correlation of Location-Specific Changes in Housing Costs and Income

	Spell length, T (years):						
	10	20	30	40	50	60	70
Panel A: Dollar changes							
Rent, $Corr(R_{\omega,t+T} - R_t, y_{\omega,t+T} - y_t)$	0.57	0.55	0.58	0.54	0.57	0.66	0.71
Home price, $Corr(P_{\omega,t+T} - P_t, y_{\omega,t+T} - y_t)$	0.54	0.38	0.39	0.32	0.27	0.28	0.30
Panel B: Log changes							
Rent, $Corr(\ln R_{\omega,t+T} - \ln R_t, \ln y_{\omega,t+T} - \ln y_t)$	0.57	0.55	0.58	0.56	0.59	0.68	0.74
Home price, $Corr(\ln P_{\omega,t+T} - \ln P_t, \ln y_{\omega,t+T} - \ln y_t)$	0.57	0.44	0.46	0.41	0.36	0.39	0.43
Panel C: Present values							
Rent, $Corr(R_\omega, Y_\omega)$	0.57	0.56	0.54	0.51	0.59	0.68	0.71
Own, $Corr(O_\omega, Y_\omega)$	-0.43	-0.26	-0.23	-0.10	-0.09	-0.14	-0.16
Panel D: Coefficients of reverse regressions of present values, $Y_\omega = \alpha + \beta_{Y\|H} H_\omega + \epsilon_\omega$							
Rent, $\beta_{Y\|R}$	1.12	1.07	1.01	1.01	1.27	1.41	1.37
	(0.08)	(0.10)	(0.13)	(0.17)	(0.17)	(0.15)	(0.13)
Own, $\beta_{Y\|O}$	-0.24	-0.34	-0.49	-0.36	-0.53	-1.16	-1.77
	(0.04)	(0.14)	(0.24)	(0.32)	(0.50)	(0.73)	(0.94)

Notes. The table shows the correlation of log changes of location-specific housing costs (rents and home prices) and wages (full-time wage and salary income) of renters and owners housing costs and net income for different spell length. Wages and housing costs are quality-adjusted as described in the text. All changes are relative to the same initial price ($R_t = \$13k$, $P_t = \$186k$, $y_t = \$58k$) and the real discount rate used for computing present values (PVs) is the average post-war yield on long-run Government bonds, $r_t = 2.5\%$. Panel B shows changes in quality-adjusted log prices (location-year fixed effects $\hat{\lambda}_{ct}$ of the hedonic regression, equation (A.3) in the Appendix). Panel A converts these quality-adjusted growth rates to dollar changes (corresponding to price component X_{it}^{loc} in equation (A.6) of the appendix). Panels A and B pool decadal data from Census years 1940 to 2010 so that the correlation coefficient of 10-year changes is based on 5,040 CZ-year observations (720 CZs times 7 10-year changes 1940–1950 to 2000–2010), 2,880 observations for 40-year changes, and 720 observations for 70-year changes. Panel C interpolates between census years to annual data spanning 1940 to 2010. The standard deviation of 10-year PVs is based on 43,920 CZ-year observations (720 CZs times 61 PVs for 1941–1950 to 2001–2010), 22,320 observations for 40-year PVs, and 720 observations for 70-year PVs. All calculations use population weights for the 720 CZs. Robust standard errors in brackets of panel D are clustered by state. Source: Decennial Census and ACS.

Table 3: Efficiency Costs of Subsidies and Misperceptions

Panel A: Ingredients

Implicit subsidies ($s per lifetime)	
Misperceptions, σ_{int}	16,000
Net subsidy, σ_{sub}	90,000

Quantity response (in lifetimes)		
	η^{low}	η^{main}
$\Delta Q_{sub} = \eta\sigma_{sub}$	0.13	0.25
$\Delta Q_{int} = \eta\sigma_{int}$	0.02	0.04

Panel B: Efficiency cost of subsidies ($s per lifetime)

	With misperception		Without misperception		Effect of misperception	
Formula	η^{low}	η^{main}	η^{low}	η^{main}	η^{low}	η^{main}
$\left(\sigma_{int} + \frac{\sigma_{sub}}{2}\right) \Delta Q_{sub}$	7,625	15,250	5,625	11,250	2,000	4,000

Panel C: Efficiency cost of misperceptions ($s per lifetime)

	With subsidies		Without subsidies		Effect of subsidies	
Formula	η^{low}	η^{main}	η^{low}	η^{main}	η^{low}	η^{main}
$\left(\sigma_{sub} + \frac{\sigma_{int}}{2}\right) \Delta Q_{int}$	2,178	4,356	178	356	2,000	4,000

Notes. σ_{int} is the internality wedge, i.e., the subsidy that would increase the quantity of owner-occupied housing by as much as misperceptions would under the assumptions discussed in Section 6. σ_{sub} is the implicit net subsidy on owner-occupied relative to rental housing beyond that justified by externalities. η^{main} (η^{low}) is the slope of demand for owner-occupied housing with respective to the subsidy rate assuming the baseline (low) price responsiveness. ΔQ_{int} (ΔQ_{sub}) is the increase in owner-occupied housing as a result of misperceptions (subsidies).

Online Appendix: Renting Insures Wage Risk

Lorenz Kueng Lee M. Lockwood Pinchuan Ong

A Data Appendix

A.1 Location-Level IPUMS USA Data

Microdata to analyze location-level risk is from IPUMS USA (Integrated Public Use Microdata Series), https://usa.ipums.org/usa (Ruggles et al. 2018). We use a 5% random sample, stratified by county, of the IPUMS USA 1940 100% "full count data" (to speed up the computations), the 1960 5% sample, the 1970 1% "metro form 1" and 1% "metro form 2" samples, the 1980 5% "state" sample, the 1990 5% "state" sample, the 2000 5% sample, and the 2012 5-year ACS 5% sample. The 1950 1% and the 2010 10% sample lack data on rent and home values. For 2010 we therefore use the 2012 5-year ACS—which is centered at 2010, covering years 2008–2012—and we linearly interpolate location-level values in 1950 with the geometric average of the values in 1940 and 1960. Our results are robust to dropping 1950 from our sample or to using the shorter period from 1960–2010. When using data at annual frequencies (e.g., when calculating present values), we also linearly interpolate with the geometric average between Census years. To compute population weights, we use accurate county- or CZ-level population counts from the full Census instead of estimates from the IPUMS samples.

A.1.1 Defining Local Labor Markets

The most detailed geographic identifier available in the microdata is a Public Use Microdata Area (PUMA) or a County Group, which maps either to a single county or a group of counties, and this assignment changes over time. To maintain confidentiality, only counties with a population larger than 100,000 (or 250,000 in 1970) are identified. The remaining counties are aggregated to county groups until these groups exceed this population threshold. We use separate allocation factors for each Decennial Census provided by IPUMS to allocate households to counties in cases where only county groups are identified or where PUMAS are larger than counties and potentially cross CZs. For

example, if a PUMA consists of two counties, A and B, with population fractions 10% and 90% respectively, we allocate households in that PUMA to both counties, and multiply their household sampling weights by 10% in county A and 90% in county B.

Although county boundaries are fairly stable during our sample period, some changes do occur. To take these into account, we use the crosswalk files from Eckert et al. (2018) which map counties in each decade to stable county identifiers based on county areas in 1990. We then assign each county to a unique CZ using the definition of CZs in 1990.[29]

A.1.2 Sample Selection

We drop housing units that are farm houses (variable FARM) and group quarters (GQ). We drop Alaska and Hawaii because their county population data are not available before they obtained statehood in 1959. We also drop two small CZs with missing prices in 1940. These two CZs contain Wayne county, UT (CZ population of 2,778 in 2010) and Norton county, KS (CZ population of 5,671 in 2010).

To make the universe of housing units consistent across decennial censuses, we use the steps recommended by IPUMS USA: "By using other variables such as COMMUSE, ACREPROP, FARM, and CONDO to eliminate certain units, users can create a fairly comparable housing value universe for 1960 on."[30] Specifically, we set housing costs (rent or home value) to missing for housing units that are commercially used (containing businesses or medical/dental offices; COMMUSE) or have 10 acres or more of land (ACREPROP and ACREHOUS). Because the 1970 Census also excludes condos and cooperatives (CONDO) and mobile homes, trailers, and boats (UNITSSTR), we also set housing costs of those units to missing in all years. Before 1970, the Census also excluded vacant-for-sale and vacant-for-rent housing units (GQ), which we therefore also code as having missing housing costs.Finally, we set housing costs to missing for units that are occupied under "no cash rent" contracts (e.g., rent-for-work arrangements; OWNERSHPD) and where renters do not report contract rents paid (RENT) or owners do not report their home value (VALUEH).

[29]The boundaries and PUMA codes are the same for the 2000 Decennial Census and the 2005–2011 ACS samples. PUMA boundaries based on the 2010 Decennial Census are used for 2012-onward ACS samples. Therefore, for the 2008–2012 ACS sample, survey respondents from 2008 to 2011 are assigned to the Census 2000 based PUMAs, while respondents from 2012 are assigned to the Census 2010 based PUMAs; see https://usa.ipums.org/usa-action/variables/PUMA#comparability_section. Since these county-to-CZ crosswalk files are available only until 2000, we use the 2000 crosswalk also for the 1% ACS 2012 sample in the 5-year ACS 2008–2012, because there were very few changes in county borders between 2000 and 2010. As a robustness check, we instead used the 2007–2011 ACS sample, which only contains PUMA boundaries based on the 2000 Decennial Census, and we obtained similar results.

[30]See https://usa.ipums.org/usa-action/variables/VALUEH#comparability_section.

To make the universe of workers consistent across decennial censuses, we use the highest earner among Census "head/householder" or "spouse" (RELATE) to define the representative household member (whether working full-time or not).[31] Finally, we restrict income observations to households where the highest earner is between 25 and 60 years old (except when measuring Social Security income (INCSS), which we restrict to individuals older than 65 and younger than 90 years old).

A.1.3 Market Rents and Home Prices

Home values are self-reported and available for owners only. Values are reported continuously in 1940 and 2010 and as the midpoint of an interval in all other years. Except in 1940, estimates include the full value of both the house and the land, even if the respondent resides on only part of the property. In 1940, the value of owner-occupied units that were part of a building containing other households or businesses included only the part of the house in which the owner's household lived. For example, if the owning household of a two-family house rented half of the house to another household, only half of the house's value would have been reported.

Figure A.3 addresses the concern that there might be systematic biases over time in survey responses about home values. Panels (a) and (b) show that Census home values align closely with both the county-level transaction-based house prices from Zillow and with the county-level transaction-based home price index from the Federal Housing Finance Agency (FHFA) during the shorter period when transaction prices are available. Importantly, panels (c) and (d) show that log *changes* of median Census home values are highly correlated with log changes of transaction-level prices or price indices. In fact, panels (e) and (f) show that Census home values align as closely if not more closely with the transaction-based measures than those two transaction-based measures align with each other. That self-assessed home values align closely with market prices is in keeping with the results of Davis and Quintin (2017).

We use contract rent, which may or may not include utilities, instead of gross rent (RENTGRS), which includes utilities, because the latter is not available in the 1940 Census, but Table A.9 shows that the relationship between wages and rents are similar when using gross rents from 1960–2010. This is because utilities are only weakly correlated with local income changes (see Figure A.14).

[31] An alternative would be to use the values of the "head of household" as the representative household member, but the definition of the head has changed considerably over this period (see https://usa.ipums.org/usa-action/variables/RELATE).

A.1.4 Market Wages

To measure market wages, we use the wage and salary income (INCWAGE) of the household member with the highest earnings and who worked full time, which the BLS defines as working at least 35 hours in the "usual week" or previous week (depending on the survey questionnaire; UHRSWORK and HRSWORK2).[32] We also require that full-time workers were employed at least 50 weeks in the past year (WKSWORK2, EMPSTAT).

A.2 Isolating Location-Specific Risk

We decompose the log price (wage, rent, home value) of the individual worker or home i in year t and in i's location $c(i)$ as:

$$\ln x_{it} \equiv \underbrace{(\ln x_{c(i),t} - \ln x_t)}_{\ln x_{ct}^{loc}} + \underbrace{\ln x_t}_{\ln x_t^{agg}} + \underbrace{(\ln x_{it} - \ln x_{c(i),t})}_{\ln x_{it}^{idio}}. \tag{A.1}$$

The first term, $\ln x_{ct}^{loc}$, is the location-specific component, capturing variation across locations at a given time. The second term, $\ln x_t^{agg}$, is the aggregate- or country-level component, capturing national-level variation over time. The third term, $\ln x_{it}^{idio}$, is the idiosyncratic component, capturing within location-year variation. We use the IPUMS USA data to measure the location-specific price component. We discuss measurement of the aggregate and idiosyncratic price component in Sections A.3 and A.4.

To isolate prices per "efficiency unit", we remove the effect of observable differences in quality and quantity on reported prices X^{ipums} in the IPUMS-USA microdata.Specifically, we estimate the following hedonic regression year-by-year with weighted least squares using household sampling weights:[33]

$$\ln X_{it}^{ipums} = \underbrace{\alpha_t + \lambda_{c(i),t} + \varepsilon_{it}}_{\ln x_{it}\text{: quality-adjusted log price}} + \underbrace{q_{it}'\,\delta_t}_{\text{quality}}. \tag{A.2}$$

The quality- and quantity-adjusted log price contains the three price components in

[32] See https://www.bls.gov/news.release/pdf/wkyeng.pdf.

[33] Before estimating each regression, we winsorize individual prices at the 1st and 99th percentiles to reduce the effect of extreme observations and changes in top-codes. For each regression, we adjust the household weights so that the regression sample is representative of the location-year (i.e., household weights sum to location Census populations in each year). This is necessary because each regression uses a different set of households, since households are either renters or owners and not all representative household members are full-time workers and hence included in our final sample. We also adjust the weights so that each location has the same weight in each year by first computing population weights in each year for each location, and then averaging those weights within location across years.

equation (A.1): the aggregate component (year fixed effect α_t), the location-specific component (location-year fixed effects λ_{ct}), and the idiosyncratic component (residual ε_{it}). The vector of quality and quantity characteristics q_{it} includes fixed effects for the age and educational attainment (variables AGE and EDUC) of the representative household member to obtain quality-adjusted full-time wages, and fixed effects for the age of the home (BUILTYR, BUILTYR2), the number of rooms (ROOMS), the number of housing units in the structure (UNITSSTR), and whether the home had complete plumbing (PLUMBING) or shared use of the kitchen (KITCHEN).[34]

Our main specification uses the estimated location-year fixed effects to measure quality-adjusted location-specific average log prices:

$$\ln x_{ct}^{loc} = \lambda_{ct}. \tag{A.3}$$

Including year fixed effects in equation (A.2) implies that the estimated location-year fixed effects control for aggregate (country-level) and idiosyncratic price changes over time. Including year effects also means that the system is over-identified which allows us to identify location-year fixed effects by normalizing them such that they have mean zero across locations in each year:[35]

$$E(\ln x_{ct}^{loc} \mid t) = 0. \tag{A.3$'$}$$

Location-year fixed effects therefore measure the location-specific price component in equation (A.1) while the year fixed effects absorb all aggregate (country-level) risk. This permits us to measure the aggregate component $\ln x_t^{agg}$ separately, including using other data that is more appropriate for this purpose as we discuss in Section A.3 below. Similarly, since we fully saturate the regression (A.2) with fixed effects for each quality q_{it}, the least-square residual $\widehat{\varepsilon}_{it}$ is uncorrelated with the location-year fixed effects:

$$Cov(\ln x_{ct}^{loc}, \widehat{\varepsilon}_{it} \mid t) = 0. \tag{A.3$''$}$$

As with the aggregate component, this property permits us to treat the idiosyncratic risk component $\ln x_{it}^{idio}$ in equation (A.1) separately, including using other sources that are more appropriate for this purpose than the repeated cross-sections of the IPUMS-USA microdata.

[34] Except for the 1970 1% "metro form 2" and the 1940 samples, number of bedrooms is also reported (BEDROOMS), but this variable is coarser than the number of rooms (ROOMS) and the two are almost co-linear.

[35] $E(\cdot \mid t)$ denotes the population average holding fixed the time dimension (i.e., summing over all individuals or locations in period t), and $E(\cdot \mid t, c)$ denotes the population average holding fixed the time dimension and location c (i.e., summing over all individuals in location c and period t).

Quality-Adjusted Prices in 1940 .— For the 1940 Census, which lacks dwelling characteristics q_{it}, for each location we use the average hedonic price component in later years to adjust for average location-specific quality of housing. Specifically, we first estimate unconditional location fixed effects, γ_c separately year-by-year for $t \in \{1960, 1970, ..., 2010\}$:

$$\ln X_{it}^{ipums} = a_t + \gamma_{c(i),t} + e_{it}.$$

The unconditional location fixed effects capture location-specific deviations of log prices from the country-level average:

$$\gamma_{ct} = E(\ln X_{i,t}^{ipums} \mid c(i), t) - E(\ln X_{j,t}^{ipums} \mid t),$$

while the conditional location fixed effects, which are not identified in 1940, represent quality-adjusted deviations:

$$\lambda_{ct} = E(\ln X_{it}^{ipums} \mid c(i), t) - E(\ln X_{jt}^{ipums} \mid t) - \Big[E(q_{it} \mid c(i), t) - E(q_{jt} \mid t)\Big]' \delta_t.$$

We then impute quality-adjusted location fixed effects in 1940 as:

$$\widehat{\lambda}_{c,1940} = \widehat{\gamma}_{c,1940} - \frac{1}{6} \sum_{s=0}^{5} (\widehat{\gamma}_{c,1960+s\cdot 10} - \widehat{\lambda}_{c,1960+s\cdot 10}),$$

where $\gamma_{c,t} - \lambda_{c,t} = \big[E(q_{it} \mid c(i), t) - E(q_{jt} \mid t)\big]' \delta_t$.[36]

Detroit and San Francisco case study.— To simplify the description of the case studies of Detroit and San Francisco and of Figure 1 without having to provide all the details outlined in this appendix, we instead extract the location-year effects directly by estimating the hedonic regression (A.2) in levels rather than logs,

$$X_{it}^{ipums} = \lambda_{c(i),t}^{\$} + \alpha_t^{\$} + q_{it}' \delta_t^{\$} + \varepsilon_{it}. \tag{A.4}$$

Figure 1 plots the cross-section of 50-year differences in location-year effects, $\lambda_{c,2010}^{\$} - \lambda_{c,1960}^{\$}$. We obtain similar results if we use hedonic regression (A.2) in logs and then follow the steps described in Section B to transform the 50-year differences of location-year effects to 50-year changes in real dollars.

Descriptive statistics.— The main text discusses the results obtained from the analysis of the location-specific price components in detail. Tables 1 and 2 provide descriptive

[36] We demean $\widehat{\lambda}_{c,1940}$ because the regression uses sample weights for later years so that $E(\widehat{\lambda}_{c,1940} \mid t = 1940)$ is not zero, while $E(\widehat{\lambda}_{c,t} \mid t) = 0$ for all $t > 1940$.

statistics of these price components. Figure 4 graphs the relationships between location-specific hedonically-adjusted wage income and housing cost changes obtained from this procedure, and Figures 3, 5, and 6 illustrate the risk of the present value of the location-specific component of wage income and housing costs.

A.3 Measuring Aggregate Risk

This section investigates the correlation of wages and housing costs at the country level and explains how we measure the aggregate price component.

A.3.1 Data for the Non-Parametric Risk Model

To model aggregate risk non-parametrically, we need data over a span of 70 years, ideally matching the period 1940–2010 of the IPUMS-USA microdata samples. Using the Decennial Census is the most natural approach, but this leaves us with only eight observations or seven decadal changes. The resulting relationship between housing prices and wages is thus noisy. Using alternative long-run macro data series does not yield a clearer pattern as these series are only imperfectly correlated over such long periods and are subject to various biases pointed out by previous research. More commonly used annual macroeconomic series on the other hand do unfortunately not go back to 1940 but typically start only in the 1960s or later, which makes them unsuitable for the non-parametric risk model, but suitable for the parametric risk model. To have an alternative source of data to model aggregate risk non-parametrically, we also employ the long-run cross-country panel from the macro-history database compiled by Jordà et al. (2019). Despite some limitations—most importantly the lack of full-time wages—it yields a robust pattern for the correlation between national rents, home prices, and incomes.

Decennial Census data.— *Home prices.—* To measure 10-year changes in aggregate home prices, we use medians reported in published Census Tables 1940–2010. These medians are based on the full population of owner-occupied housing units rather than the IPUMS samples, are not subject to top-coding, and are also available in 1950 where housing costs are missing in the IPUMS microdata.

Rents.— Census Tables only report median gross rents, not contract rents. We therefore use median contract rents from the IPUMS microdata. We make two adjustment to this series. First, we impute median rent in 1950, when housing prices are not available in the IPUMS sample, using median gross rents from the Census Tables in 1950 and 1960. Specifically, we multiply median gross rent from the 1950 Census Table with the ratio of

median contract rent from the 1960 IPUMS sample to median gross rent form the 1960 Census Table. This ratio of median gross rent in the Census Tables to media contract rent in the IPUMS samples is relatively stable in the first half of the period (about 1.25). Second, because the rent question was different in 1940 than in later years as discussed above (e.g., it included dollar equivalents of non-cash payments), reported rent values are too high in 1940 relative to later years, both for contract rents and gross rents.[37] Consequently, the ratio of median home price to median annual contract rent of 12.2 in 1940 is unreasonably low compared to an average ratio of 17.1 from 1950–1970. Using the IPUMS microdata, we compute the median ratio of contract rent to full-time wage and of home price to full-time wage to confirm that the low price-to-rent ratio in 1940 is indeed due to high rents rather than low home values. Specifically, while the median price-to-wage ratio is fairly similar over the period from 1940 to 1970 (about 2.0), the median rent-to-wage ratio is much higher in 1940 (0.24) than in 1960 and 1970 (0.19). We therefore adjust median contract rent down in 1940 using the ratio of gross rents from the Census Table to full-time wages from IPUMS observed in both 1940 and 1950 and the ratio of gross rent to contract rent when first observed in both Census Tables and IPUMS microdata in 1960.

Wages.— Full-time wages are not reported in Census Tables, only median family income and consistently only from 1953 on. Fortunately, we can use the IPUMS samples for all ears from 1940 on. While housing costs are not available in the IPUMS microdata in 1950, wages and hours worked are. We can therefore use the IPUMS samples to create a continuous series of country-level median full-time wages for all decades from 1940 to 2010.

Findings.— Figure A.4 graphs our national-level price series and compares them to other frequently used US price series. All nominal prices are deflated using the CPI for all urban consumers and normalized to 100 in 1980 to make them comparable. In general, our series (in black) align reasonably well with these other series, especially in the later part of the sample. For instance, our series of median full-time wages from the IPUMS sample is very similar to the series of median usual weekly full-time earnings from the Current Population Survey (CPS) Tables published by the BLS (U.S. Bureau of Labor Statistics 1979–2021c). Uncertainty about the quality of the data for earlier decades is however substantial. For instance, Crone et al. (2010) show that both the annual CPI and the annual PCE rent series are too high in the first half of the 20th century and propose an adjustment for non-response bias from 1940 to 1985 for the CPI rent series. Their adjusted CPI rent series is also shown in panel (b) of Figure A.4, together with the PCE rent series (U.S. Bureau of Economic Analysis 1929–2021b), the median

[37] See https://usa.ipums.org/usa-action/variables/RENT#comparability_section.

gross rent series from the Census Tables (raw and adjusted for the different survey question in 1940; U.S. Census Bureau 1940–2000), and our series of median contract rents from the IPUMS samples (Ruggles et al. 2018). The rent series obtained from Decennial Censuses suggests that this correction might be too modest when looking at long time periods. This is consistent with recent work by Agorastos et al. (2024) based on newly collected historical data on rent and home sale listings in newspapers, which show that market rents increased rapidly in the 1940s after World War II rent controls were lifted. Similarly, panel (c) compare our measure with the repeat-sale index from Shiller (1890–2021) and with transaction-based indices from U.S. Department of Housing and Urban Development (1963–2021) and U.S. Federal Housing Finance Agency (1975–2021). It shows that over such long time spans, measures of aggregate home prices can diverge substantially.

Keeping in mind these caveats, Figure A.5 illustrates that the positive relationship between wage and housing cost changes, notably rents, persists at the national level. These three series are our baseline measure of aggregate risk for the non-parametric risk model, and we supplement them with aggregate risk measure from cross-country data that paint a similar but more robust picture of the correlation between aggregate prices.

Cross-country data.— An alternative approach to measuring aggregate risk is to view national-level shocks as risk draws from a statistical population of countries. We can measure such aggregate shocks using annual cross-country data of real housing costs and national income (real GDP per capita at purchasing power parity) obtained from Jordà et al. (2019). Because for each country, data on income and housing costs becomes available in different years, for our purpose this is an unbalanced panel which includes the following 16 countries with the corresponding first year observed in the data in brackets: Australia (1901), Belgium (1890), Denmark (1875), Finland (1920), France (1870), Germany (1870), Italy (1927), Japan (1931), Netherlands (1870), Norway (1871), Portugal (1948), Spain (1900), Sweden (1883), Switzerland (1901), the UK (1895), and the US (1890).

The data contains valuation ratios such as the housing capital gain rate and housing rental yield from which we can recover one-year growth rates of real rents and real home prices. We then cumulate these growth rates from one up to 70 years, convert them to log changes, and winsorize their unweighted distribution at the 1^{st} and 99^{th} percentile.

A few of the one-year growth rates are missing in later years for some countries, less than 5% of observations in total. However, since we construct long-run growth rates

from these one-year rates, these missing data reduce the number of observations of the cumulated longer-horizon growth rates substantially, especially at very long horizons. In our analysis, we therefore proceed in two steps. First, for the period 1940–2010 which coincides with the IPUMS data sample, we do not make any adjustments and just report population-weighted statistics. As a result, we cannot estimate volatilities and correlations for very long horizons in that sub.sample, as those would be based on very few observations, typically much fewer than 100 (e.g., we only have 9 countries over the period 1940–2010 with full 70-year growth rates even though we observe income and housing data for each of the 16 countries already before 1940). Second, we impute these missing one-year growth rates so that, while we still have an unbalanced panel, we have no missing data between the initial year a country is observed and the final year 2015, and we therefore have many more observations also for long-run growth rates. Specifically, we linearly interpolate nine observations of one-year growth rates, four truly missing rates for Finland (1927–1928 and 1945–1946) and five observations where the one-year growth rates are exactly zero (for France and Germany), which is likely due to measurement error. The remaining cases of missing data concern longer periods, ranging from 7 years for the UK (1940–1946) to 24 years for Germany (1939–1962). For those cases, we use the same number of one-year growth rates that immediately follow the missing data (e.g. 1947–1953 in the case of the UK). Relative to random sampling, this has the advantage that it preserves the auto-correlation structure non-parametrically, at least for the length of the missing period. However, we obtained similar results using alternative imputation methods.

We therefore draw aggregate rent, home price, and income growth rates from this data spanning 1870–2015 as an alternative source for the non-parametric risk model of aggregate risk discussed in Section B.1 below. Compared to the baseline data from the Decennial Census, the cross-country panel does not rely on the single time series drawn for the US from the population of potential alternative histories (i.e., only 8 data points), and it is available at annual rather than just decadal frequency. Its main disadvantage is that it does not directly measure full-time wage income but national income in the form of GDP per capita.

Findings.— Is the positive relationship between income and housing costs specific to housing and labor markets in the U.S.? Table A.3 and Figure A.6 show for horizons from one up to 70 years the standard deviation of annualized log changes ($\Delta_T \ln x^{agg}_{t+T}/\sqrt{T}$), the correlation of log changes ($Corr(\Delta_T \ln x^{agg}_{t+T}, \Delta_T \ln y^{agg}_{t+T})$), and the corresponding cross-price elasticities $\beta_{wls,x|y}$ obtained from regressing the log price changes on each other ($\Delta_T \ln x^{agg}_{t+T} = a + \beta_{wls,x|y}(T) \cdot \Delta_T \ln y^{agg}_{t+T} + u_{t+T}$). Each statistic is computed on two samples: on the sample period 1940–2010 that matches the rest of the analysis and on the

entire sample from 1870–2015 with missing data imputed as described above. Statistics for the 1940–2010 sample (without imputations) use weights corresponding to country population size relative to total population in each year that add up to the number of countries in each year of the unbalanced panel so that each year has the same weight in the calculations. Weights for statistics on the entire 1870–2015 sample are population shares that are constant across all years and are computed over the period where all 16 countries are present (mostly years 1949–2015). This is to avoid over-weighting country-years early in the sample where only few countries are present in the panel. All statistics shown use at least 100 observations and are omitted otherwise.

Table A.3 and Figure A.6 show a similarly strong positive association between growth in aggregate housing costs and growth in income across countries over long periods as the correlations obtained for location-specific changes within the US from 1940–2010. The positive correlations across countries obtains for the same 71-year period from 1940–2010 spanned by the IPUMS microdata, and over the entire 146-year period from 1870–2015. This suggests that our findings of strong correlations between wages and housing prices extend to other countries and time periods, and also extends from location-specific risk to aggregate risk.

Comparing location-specific with national-level risk (Figure A.2 with Figure A.8 and Tables 1 and 2 with Tables A.1, A.2, and A.3) we see that despite using completely different data sources and levels of aggregation, they lead to similar findings. First, the standard deviation of annualized log changes is remarkable stable over time, suggesting that location-specific and national-level shocks are very persistent.

Second, housing price growth rates are much more volatile than income growth rates, although not in real dollars because income is of course much larger than housing costs. Surprising to us, we find that national-level risk is substantially larger than location-specific risk. Depending on the horizon, the volatility of annualized log changes of housing cost is about 9% at the national-level, while location-specific volatilities are only about 5% for home prices and 4% for rents. It is possible that part of the difference, especially for shorter horizons, is due to more measurement error in the cross-country data, because it is notoriously difficult to measure national income and housing costs consistently across countries and over time, especially over such a long period. By measuring location-specific risk against the national average in each year, many of these issues can be avoided or controlled for.

Third, the correlations are consistently positive, large, and similar for location-specific and national-level risk. Depending on the horizon, the correlation is about 60% for rent and income (50% for national-level risk), 40% for home price and income, and typically

above 70% for home prices and rents, although correlations in national-level data appear higher pre-WWII than in the post-war period.

Fourth, for the cross-country data which is available at annual frequency, correlations increase steeply over the first 10 to 20 years. This could reflect sluggish adjustment of market price, but it is also consistent with measurement error or transitory shocks which might attenuate correlations at higher frequencies while at lower frequencies the persistent long-run shocks dominate the measurement.

Fifth, putting these facts together, the regression-based cross-price elasticities are remarkably similar for location-specific and national-level prices, over time, and across pairs of prices. Housing cost elasticities with income are about one for horizons longer than 10 years (both home price and rent elasticties), while they are increase monotonically and quickly at shorter horizons. Remarkably, the income elasticity of both housing costs tend to be even larger than the elasticity of home prices to a change in rents, especially for national-level changes.

A.3.2 Annual Data for the Parametric Risk Model

To quantify the aggregate price component for the parametric data generating process of the numerical life cycle model, we use aggregate data series available at annual frequency. In addition to the annual cross-country data described above, we also collect aggregate US data series that are most commonly used in macroeconomics. While most of these series do not go back in time as far as 1940 as is the case for the Decennial Census price series, they are available at annual frequency.

Wages.— Our main measure of aggregate market wages uses the IPUMS ASEC microdata (i.e., the Annual Social and Economic (ASEC) supplement to the Current Population Survey (CPS); Flood et al. 2021) to define full-time wages in the same way as described in Section 3 and Appendix A.1.4 for the IPUMS-USA Decennial Census microdata. The IPUMS-CPS ASEC sample is 1962–2021, with work and income questions referencing the previous calendar year, hence CPS full-time wage ends in 2020. As a robustness check, we also use the shorter series "median usual weekly real earnings of wage and salary workers 16 years and over employed full time" provided by the BLS (U.S. Bureau of Labor Statistics 1979–2021c). This series is based on the CPS "merged outgoing rotation group" (or "earner study" or "annual earnings files") of the basic monthly samples, which began in 1979. Work and income questions reference current pay, hence this series ends in 2021.

Rents.— Our main measure of aggregate rents uses the Consumer Price Index (CPI)

"rent of primary residence in U.S. city average" series (U.S. Bureau of Labor Statistics 1914–2021b) provided by the BLS. Note that the upward bias in this series, pointed out by Crone et al. (2010), pertains to the period before 1980 and does not affect our analysis of annual aggregate rents and their co-movement with aggregate wages. As a robustness check, we also use the Personal Consumption Expenditures (PCE) series "rental chain-type price index of tenant-occupied nonfarm housing" provided by the BEA (U.S. Bureau of Economic Analysis 1929–2021c).

Home prices.— Our main measure of aggregate home prices using the "all-transactions house price index for the United States" series provided by the FHFA (U.S. Federal Housing Finance Agency 1975–2021). As a robustness check, we also measure aggregate home prices using real U.S. home price data for Figure 3.1 in Robert J. Shiller's Irrational Exuberance (Shiller 1890–2021).[38]

Price index.— All nominal prices are deflated using the CPI for All Urban Consumers (U.S. Bureau of Labor Statistics 1913–2021a). As a robustness check, we also use the PCE Price Index (U.S. Bureau of Economic Analysis 1929–2021b) and the Gross Domestic Product (GDP) Implicit Price Deflator (U.S. Bureau of Economic Analysis 1947–2021a) and find similar results.

Findings.— Figure A.7 graphs the relationship between annual macro housing prices and wage income using the different measures described above. It shows strong positive correlations between country-level growth in wages and housing costs, comparable with the evidence for location-specific growth rates.

Tables A.1, A.2, and A.3 and Figure A.8 provide additional information.

A.4 Measuring Idiosyncratic Risk

B Risk Models

Most of our analysis uses a non-parametric risk model with minimal assumptions about the data generating process. The main exception is the numerical life-cycle model, which we solve with dynamic programming. The dynamic program uses a recursive structure, for which we need to specify a parametric model of the risk process. We describe both non-parametric and parametric risk models in detail in this section.

[38] Figure A.4.c shows that the alternative series of home prices, "Median Sales Price for New Houses Sold in the United States" provided by HUD (U.S. Department of Housing and Urban Development 1963–2021), does not correlate well with the other aggregate home price series.

B.1 Non-Parametric Risk Model

We use a non-parametric model of the evolution of prices across locations, individuals, and time that requires minimal assumptions about the data generating process. Equation (A.1) decomposes the log of individual prices and hence estimates the hedonic price regressions (A.2) in logs. To compute present values of wage income and housing costs, we need to transform these quality-adjusted log prices back to prices in real dollars.[39]

We want the non-parametric risk model of these prices to be flexible so that we can explore different scenarios (e.g., with our without aggregate and idiosyncratic risk, with or without moves, etc.). We also want the resulting price paths to have some desirable properties. First, as will become clear when we discuss moves across locations, we need to distinguish between the price path faced by an individual household and the price path of a location, even in the case where there is no idiosyncratic risk. This is because in our baseline specification, all individuals start from the same initial condition, which can be thought of as the price vector of the 'typical' location. The actual locations on the other hand start from their respective initial price vector, which is different for each location (e.g. there are expensive and cheap locations). When individuals later move, they move to one of these 'actual' locations and this new location's price vector reflects the initial cross-sectional dispersion of prices which is thus an additional source of risk if moves occur at least partially for exogenous reasons. Second, we also want to start our risk analysis from initial conditions that reflect today's prices rather than the prices in 1940. Third, we want to incorporate all three types of risks—location-specific, aggregate, and idiosyncratic—and we want to be able to easily turn each source of risk on or off.

Local price paths.— We start with as simple decomposition of the price path in location c, $X^*_{c,t}$:

$$\begin{aligned} X^*_{c,t+s} &= \frac{\frac{X^*_{c,t+s}}{X^*_{t+s}}}{\frac{X^*_{c,t}}{X^*_t}} \times \frac{X^*_{t+s}}{X^*_t} \times \frac{X^*_{c,t}}{X^*_t} \times X^*_t \\ &= e^{\Delta_s(\ln X^*_{c,t+s} - \ln X^*_{t+s})} \times e^{\Delta_s \ln X^*_{t+s}} \times e^{\ln X^*_{c,t} - \ln X^*_t} \times X^*_t. \end{aligned} \tag{A.5*}$$

We use $*$ to distinguish between the price path X^* that we could potentially observe in an ideal data set (i.e., in the statistical population) from the price path X that we

[39] We prefer decomposing prices in logs rather than in levels so that the coefficients from the hedonic regression reflect relative price changes, because when computing present values over long periods, we can compare relative prices more easily without the risk of over-weighting later census years as those later years had higher prices in real dollars and hence would contribute much more to the price variation.

can actually study in our non-parametric analysis of housing and income risk with the available data. We use $s = 0, 1, 2, \ldots, T$ to index years within the housing spell that lasts T periods and begins in year t (i.e., in period $s = 0$).

X_t^* measures the country-level price level. To start our risk analysis from today's prices rather than the prices in 1940, we replace the initial aggregate price X_t^* with the price faced by the median full-time worker in 2010, which we denote by X^o (i.e., \$58k wage income, \$13k rent, and \$186k home price): $X_t^* = X^o$. For example, for a homeowner whose spell starts in $t = 1953$, the initial home price is the median home price of a full-time worker in 2010, $P^o_{1953} =$ \$186k, which then grows at the growth rate observed from 1953 on. For a housing spell that starts in $t = 1979$, the initial price level is also $P^o_{1979} =$ \$186k but then grows at the growth rate of home prices observed from 1979 on.[40]

$\Delta_s \ln X^*_{t+s}$ is aggregate risk and corresponds conceptually to $\Delta_s \ln x^{agg}_{t+s}$ in the s-period differenced version of equation (A.1):

$$\Delta_s \ln x_{i,t+s} = \Delta_s \ln x^{loc}_{c,t+s} + \Delta_s \ln x^{agg}_{t+s} + \Delta_s \ln x^{idio}_{i,t+s}. \tag{A.1′}$$

We can measure it either with observed changes of a measure of country-level prices in real dollars (e.g., the single realization of observed median prices from the US Decennial Census data, $\Delta_s \ln x^{agg}_{t+s} = \Delta_s \ln X^{Census}_{t+s}$) or with changes of a measure of a country-level price index (e.g., non-parametric draws from the cross-country panel data set of price index values obtained from Jordà et al. 2019, $\Delta_s \ln x^{agg}_{t+s} = \Delta_s \ln X^{Country(i)}_{t+s}$).

$\ln X^*_{c,t+s} - \ln X^*_{t+s}$ is the log of the location's price relative to the country-level price level. It represents heterogeneity in initial prices across locations, which might not reflect risk because it is know at the beginning of the housing spell in year t (in period $s = 0$). Its s-period difference, $\Delta_s(\ln X^*_{c,t+s} - \ln X^*_{t+s})$, on the other hand represents location-specific risk. $\ln X^*_{c,t+s} - \ln X^*_{t+s}$ corresponds conceptually to $\ln x^{loc}_{c,t+s}$ in equation (A.1), which we measure using the hedonic regression (A.2) applied to the IPUMS USA microdata. Clearly, we would like this relative price to be one on average, i.e., $E(\frac{X^*_{c,t+s}}{X^*_{t+s}} \mid t+s) = E(e^{\ln X^*_{c,t+s} - \ln X^*_{t+s}} \mid t+s) = 1$, such that $E(X^*_{c,t+s} \mid t+s) = X^*_{t+s}$. Since the location-year fixed effect $\ln x^{loc}_{c,t+s}$ from the hedonic regression has mean zero by construction (see equation A.3′), we need to adjust for the effect of the convex transformation on the expected price level due to Jensen's inequality. To see this, note that if we substituted $\ln x^{loc}_{c,t+s}$ for $\ln X^*_{c,t+s} - \ln X^*_{t+s}$ without adjustment, we would obtain $E(e^{\ln x^{loc}_{c,t+s}} \mid t+s) \approx e^{\mu_{t+s} + \frac{1}{2}\sigma^2_{t+s}} = e^{\frac{1}{2}\sigma^2_{t+s}} > 1$, where $\mu_{t+s} = E(\ln x^{loc}_{c,t+s} \mid t+s) = 0$ by equation (A.3′) and

[40] For the case study of price changes in Detroit and San Francisco from 1960–2010, we instead use the average of the median prices in 1960 and 2010 as we discuss in more detail below.

$\sigma^2_{t+s} = Var(\ln x^{loc}_{c,t+s} \mid t+s) > 0$. The approximation would be exact if $\ln x^{loc}_{c,t+s}$ was normally distributed. Dividing the realized location-specific growth shock, $e^{\ln x^{loc}_{c,t+s}}$, by its cross-sectional average, $E(e^{\ln x^{loc}_{c,t+s}} \mid t+s)$, makes this adjustment non-parametrically (without assuming that $x^{loc}_{c,t+s}$ is log-normal).

Hence, we define the path of local prices $X_{c,t+s}$ used for our risk analysis based on the price decomposition (A.5*), but using data from different sources to model location-specific and aggregate risk non-parametrically, as follows:

$$\begin{aligned} X_{c,t+s} &= k^{loc}_{t,t+s}\, e^{\Delta_s \ln x^{loc}_{c,t+s}} \times e^{\Delta_s \ln x^{agg}_{t+s}} \times \kappa^{loc}_t e^{\ln x^{loc}_{c,t}} \times X^o \\ &= \underbrace{\kappa^{loc}_{t+s}\, e^{\ln x^{loc}_{c,t+s}}}_{X^{loc}_{c,t+s}} \times \overbrace{X^o}^{X^{agg}_{t+s}} \times\ e^{\Delta_s \ln x^{agg}_{t+s}}, \end{aligned} \tag{A.5}$$

where k^{loc} and κ^{loc} are adjustment factors for the effect of Jensen's inequality on the expected price level as described above.[41] The second line defines the aggregate price path X^{agg}_{t+s}, with the desirable property that $E(X_{c,t+s} \mid t+s) = X^{agg}_{t+s}$, and the location-specific price paths $X^{loc}_{c,t+s}$. These location-specific prices are the local prices that would apply in location c if there was no aggregate risk ($\Delta_s \ln x^{agg}_{t+s} = 0$).

Individual price paths.— Next, we turn to the price that an individual household i faces who is planning to own or rent housing in initial location $c(i)$ in the initial year t for a spell of T periods. In addition to the location-specific and aggregate risk already included in the definition of the local price path (A.5), the household is also exposed to idiosyncratic risk as shown in equation (A.1′). Building on equation (A.5*), we can decompose individual prices as:

$$\begin{aligned} X^*_{i,t+s} &= \frac{\frac{X^*_{i,t+s}}{X^*_{c,t+s}}}{\frac{X^*_{i,t}}{X^*_{c,t}}} \times \frac{X^*_{i,t}}{X^*_{c,t}} \times X^*_{c,t+s} \\ &= e^{\Delta_s(\ln X^*_{i,t+s} - \ln X^*_{c,t+s})} \times e^{\ln X^*_{i,t} - \ln X^*_{c,t}} \times X^*_{c,t+s}. \end{aligned} \tag{A.6*}$$

The first two new terms reflect the addition of the idiosyncratic price component. The first captures idiosyncratic risk and the second additional heterogeneity in the initial price, which again might not reflect risk because it is know at the beginning of the housing spell in year t (in period $s = 0$). The last term $X^*_{c,t+s}$ is the local price path if we had access to the ideal data, with its decomposition given by equation (A.5*). To bring

[41] $\kappa^{loc}_{t+s} = 1/E(e^{\ln x^{loc}_{c,t+s}} \mid t+s)$, $\kappa^{loc}_t = 1/E(e^{\ln x^{loc}_{c,t}} \mid t)$, and $k^{loc}_{t,t+s} = 1/E(e^{\Delta_s \ln x^{loc}_{c,t+s}} \mid t, t+s)$.

this price decomposition to the data that we do have available, we replace $\ln X^*_{i,t} - \ln X^*_{c,t}$ with its conceptual counterpart $\ln x^{idio}_{i,t}$ in equation (A.1) and $\Delta_s(\ln X^*_{i,t} - \ln X^*_{c,t})$ with $\Delta_s \ln x^{idio}_{i,t}$ in equation (A.1′), and we adjust for the effect of Jensen's inequality from transforming hedonically-adjusted log prices to prices in real dollars. Finally, as before, we replace unobserved $X^*_{c,t+s}$ with observed $X_{c,t+s}$ from equation (A.5). If we had access to panel data for individual workers over a 70 year period, we could also model idiosyncratic risk non-parametrically in the same way we model it for the local price paths using our 70-year panel data of local prices. That is, we could replace $e^{\ln X^*_{i,t} - \ln X^*_{c,t}}$ with $e^{\ln x^{idio}_{i,t}} / E(e^{\ln x^{idio}_{i,t}} \mid t)$, where $\ln x^{idio}_{i,t}$ would be the hedonically-adjusted idiosyncratic price component. Since we do not have access to such long-run panel data for individuals, which would also have to include information about their location, we add idiosyncratic risk as i.i.d. draws from a log-normal distribution, where we calibrate its variance using studies that measure such idiosyncratic risk for the three prices (full-time wage income, rents, and home prices): $\ln x^{idio}_{i,t} \sim i.i.d.N(0, \sigma^2_{idio,x})$.

Hence, we define the path of individual prices used for our risk analysis based on the price decomposition (A.6*), but using data from different sources to model idiosyncratic, location-specific, and aggregate risk, as follows:

$$X_{i,t+s} = k^{idio}_{t,t+s}\, e^{\Delta_s \ln x^{idio}_{i,t+s}} \times \kappa^{idio}_t\, e^{\ln x^{idio}_{i,t}} \times X_{c,t+s} \tag{A.6}$$

$$= \underbrace{k^{loc}_{t,t+s}\, e^{\Delta_s \ln x^{loc}_{c,t+s}} \times \overbrace{X^o}^{X^{agg}_{t+s}} \times e^{\Delta_s \ln x^{agg}_{c,t+s}}}_{X^{loc}_{i,t+s}} \times k^{idio}_{t,t+s}\, e^{\Delta_s \ln x^{idio}_{i,t+s}} \times \kappa^{loc}_t e^{\ln x^{loc}_{c,t}} \times \kappa^{idio}_t\, e^{\ln x^{idio}_{i,t}},$$

where $k^{idio}_{t,t+s}$ and κ^{idio}_t are again adjustment factors for the effect of Jensen's inequality on the expected price level.[42] The second line defines the location-specific price paths, $X^{loc}_{i,t+s}$, that household i faces. These location-specific individual prices are the prices that apply if there was no aggregate and idiosyncratic risk ($\Delta_s \ln x^{agg}_{t+s} = \Delta_s \ln x^{idio}_{i,t+s} = 0 \;\forall\; s$) and no heterogeneity in initial prices ($\ln x^{loc}_{c,t} = \ln x^{idio}_{i,t} = 0$) as is the case in our baseline specification described below. Hence, these are the price paths shown in Figure 2.

Since idiosyncratic risk is independent of location-specific and aggregate risk, its average within a location is also zero, $E(\ln x_{i,t+s} \mid c(i), t+s) = 0$, yielding the desirable property that without any restrictions (such as those of the baseline specification below), individual prices within a location average to the local price level: $E(X_{i,t+s} \mid c(i), t+s) = X_{c(i),t+s}$.

[42] $k^{idio}_{t,t+s} = 1/E(e^{\Delta_s \ln x^{idio}_{i,t+s}} \mid t, t+s)$ and $\kappa^{idio}_t = 1/E(e^{\ln x^{idio}_{i,t}} \mid t)$.

Baseline specification.— Our baseline specification of the non-parametric risk model is as follows:

- *No heterogeneity in initial individual prices* $X_{i,t}$*:* $\ln x_{c,t}^{loc} = \ln x_{i,t}^{idio} = 0$ in (A.6) but not in (A.5), implying $\kappa_t^{loc} = \kappa_t^{idio} = 1$.
- *No aggregate risk:* $\Delta_s \ln x_{t+s}^{agg} = 0 \; \forall \; s$ in both (A.6) and (A.5).
- *No idiosyncratic risk:* $\Delta_s \ln x_{t+s}^{idio} = 0 \; \forall \; s$ in both (A.6) and (A.5), implying $k_{t,t+s}^{idio} = 1$.

Hence, in the baseline, the individual price paths are equal to $X_{i,t+s}^{loc}$ in equation (A.6):

$$X_{i,t+s} = X_{i,t+s}^{loc}, \tag{A.6^{baseline}}$$

and the local price paths are equal to $X_{c,t+s}^{loc}$ in equation (A.5):

$$X_{c,t+s} = X_{c,t+s}^{loc} = X_{i,t+s}^{loc} \times \kappa_t^{loc} e^{\ln x_{c,t}^{loc}}. \tag{A.5^{baseline}}$$

Therefore, in the baseline, the difference between these two local-specific price paths is the initial heterogeneity in prices across locations, $\kappa_t^{loc} e^{\ln x_{c,t}^{loc}}$, which is included in the local price paths but excluded from the individual price paths. As a consequence, the within-location average of individual prices does not equal the local price in the baseline specification: $E(X_{i,t+s} \mid c(i), t+s) \neq X_{c(i),t+s}$.

Moves across locations.— In the descriptive analysis in Section 4, households do not move. Therefore, their price risk is fully characterized by the individual price paths in equation (A.6). We shut down differences in initial prices in the baseline specification because they do not necessarily reflect risk. We do this to be conservative to avoid overstating the amount of housing risk and the value of renting as a hedge against uninsured income risk. The value of the rental hedge would increase if we treat this initial heterogeneity as risk because the cross-sectional correlation between wage income and housing prices is positive and very strong as shown in Figure A.1.

When we extend the analysis to include moves to new locations (e.g., in Section 5 when quantifying risk costs of owning versus renting over a full 70-year life cycle), we keep track of both the initial location (in period $s = 0$) and the subsequent locations to which the household might move in future periods ($s > 0$). For the initial location, we model risk exactly as in equation (A.6): The initial price vector X^o is the same vector of initial wage income, rent, and home price in all states of the world, and the subsequent evolution of prices in that initial location is drawn from the empirical distribution of

price growth across locations. So the risk in prices in the initial location comes entirely from risk in the evolution of prices during the household's stay (starting from a common initial price), not from any pre-existing differences across locations in the initial prices. For subsequent locations to which the household might move, the price paths are the observed price paths in each location, i.e., prices evolve according to equation (A.5) not from a single initial price vector that is the same in all locations, but from the observed heterogeneous prices across locations at time zero (Figure A.1 plots the distribution of this initial price heterogeneity). We adopt this risk model both (i) to exclude any differences in prices in the initial period from our measures of risk (since our baseline perspective considers risk from that point forward, though as an alternative perspective we also consider risk from behind the veil, e.g., from being born into an initial location) and (ii) to account for the considerable differences across locations in price levels (which can be an important determinant of the riskiness of owning versus renting for movers).

Example.— Consider an owner who moves to a new location every 20 years, and for simplicity assume that there is no aggregate and idiosyncratic price risk as in our baseline specification of the risk model. The initial known purchase price at time zero is $P^{buy}_{i,t} = P^o = \$186\text{k}$. Subsequent growth in this initial price is drawn from the empirical distribution of price growth. Suppose the draw for the first 20 years is New York City from 1940 to 1960. Then this price grows according to the observed price growth in New York City from 1940 to 1960, $\Delta_{20} \ln x^{loc}_{c,t+20}$, where c is New York City. Therefore, the sale price at time $t+20$, which is unknown at t, is $P^{sell}_{i,t+20} = k^{loc}_{t,t+20}\ e^{\Delta_{20} \ln x^{loc}_{c,t+20}}\ P^o = \frac{e^{\Delta_{20} \ln x^{loc}_{NYC,1960}}}{E(e^{\Delta_{20} \ln x^{loc}_{c,1960}})}\ P^o$ according to equation (A.6^{baseline}). The unknown purchase price in the new location c' in period $t+20$ (e.g., Austin in 1960), which provides housing service starting in period $t+21$ (1961), is $P^{buy}_{i,t+20} = \kappa_{t+20}\ e^{\ln x^{loc}_{c',t+20}}\ P^o = \frac{e^{\ln x^{loc}_{Austin,1960}}}{E(e^{\ln x^{loc}_{c,1960}})}\ P^o$ according to the second line of equation (A.5), with $\Delta_{20} \ln x^{agg}_{t+20} = 0$ in the baseline specification. This price then evolves at the growth rate of home prices in location c' for 20 years, $\Delta_{20} \ln x^{loc}_{c',t+40}$ (Austin from 1960 to 1980), such that the unknown sale price is $P^{sell}_{i,t+40} = \kappa_{t+40}\ e^{\ln x^{loc}_{c',t+40}}\ P^o = \frac{e^{\ln x^{loc}_{Austin,1980}}}{E(e^{\ln x^{loc}_{c,1980}})}\ P^o$ and the unknown purchase price in the next location c'' (e.g., Chicago in 1980) is $P^{buy}_{i,t+40} = \kappa_{t+40}\ e^{\ln x^{loc}_{c'',t+40}}\ P^o = \frac{e^{\ln x^{loc}_{Chicago,1980}}}{E(e^{\ln x^{loc}_{c,1980}})}\ P^o$, and so on.

B.2 Parametric Risk Model

For the numerical life cycle model, we use a parametric model of the evolution of prices across locations and time, including idiosyncratic risk within location-time.

Location-specific risk.— We model the vector of location-specific log prices, $\ln x_{ct}^{loc} = (\ln y_{ct}^{loc}, \ln R_{ct}^{loc}, \ln P_{ct}^{loc})'$, which we observe every decade, as a first-order panel vector auto-regressive process (pVAR):

$$\ln x_{ct}^{loc} = \rho_{10}^{loc} \ln x_{c,t-10}^{loc} + u_{ct}^{loc} \quad \text{for } t \in \{10, 20, \ldots, 70\}. \tag{A.7}$$

The decadal growth rate shock u_{ct}^{loc} is modeled as independently and joint normally distributed:

$$u_{ct}^{loc} \sim i.i.d.\ N(0, \Sigma_{10}^{loc}). \tag{A.8}$$

The joint distribution of initial log prices is also modeled as being normally distributed:

$$\ln x_{c,0}^{loc} = u_{c,0}^{loc} \sim N(0, \Sigma_{0}^{loc}). \tag{A.9}$$

We estimate the matrix of auto-regression coefficients, $\widehat{\rho}_{10}^{loc}$, and the covariance matrix of the decadal growth rate shock, $\widehat{\Sigma}_{10}^{loc}$, by weighted least squares using data from 1940–2010 and location population weights.

We use the sample variance of log prices pooled across all decades to estimate the covariance matrix of the joint distribution of initial log prices:

$$\widehat{\Sigma}_{0}^{loc} = Var(\ln x_{ct}^{loc}). \tag{A.10}$$

Alternatively, we could also use the implied long-run variance of log prices of the stationary process (A.7), $\ln x_{c,0}^{loc}$ (see Figure A.1):

$$vec(\widehat{\Sigma}_{0}^{loc}) = (I_9 - \widehat{\rho}_{10}^{loc} \otimes \widehat{\rho}_{10}^{loc})^{-1} vec(\widehat{\Sigma}_{10}^{loc}),$$

where $\otimes$ denotes the matrix direct (Kronecker) product and $vec(\Sigma)$ is the vectorization operator applied to Σ. We also explored using the cross-section of log prices in 1940 as starting values. Using either estimate of Σ_0^{loc} or the 1940 distribution yields similar results. In our baseline, we use the covariance matrix (A.10) which leads to a price dispersion that lies between that obtained when using the implied long-run variance and that obtained when using the 1940 cross-section.

Annualizing the decadal panel VAR.— We annualize process (A.7) as follows. Let the unobserved vector of annual location-specific log prices also follow a first-order pVAR:

$$\begin{aligned} \ln x_{ct}^{loc} &= \rho_1^{loc} \ln x_{c,t-1}^{loc} + v_{ct}^{loc} \quad \text{for } t \in \{1, 2, \ldots, 70\} \\ &= (\rho_1^{loc})^{10} \ln x_{c,t-10}^{loc} + \sum_{i=0}^{9} \rho_1^{loc} v_{c,t-i}. \end{aligned} \tag{A.11}$$

The annual growth rate shock v_{ct}^{loc} is independently and joint normally distributed:

$$v_{ct}^{loc} \sim i.i.d.\ N(0, \Sigma_1^{loc}).$$

Using the eigenvalue decomposition of the matrix of auto-regressive coefficients, ρ_{10}^{loc}, we can recover the implied auto-regressive coefficients of the annual process:

$$\rho_1^{loc} = (\rho_{10}^{loc})^{\frac{1}{10}} = V\ D^{\frac{1}{10}}\ V^{-1},$$

where D is the diagonal matrix containing the eigenvalues of ρ_{10}^{loc} and V is the matrix of its eigenvectors, such that $\rho_{10}^{loc} = V\ D\ V^{-1}$. Using this matrix ρ_1^{loc} and the two properties of the *vec* operator and the Kronecker product, $vec(A + B) = vec(A) + vec(B)$ and $vec(ABC) = (C' \otimes A)\ vec(B)$, we can recover the implied covariance matrix of the unobserved annual growth rate shocks:

$$vec(\Sigma_1^{loc}) = \left(\sum_{i=0}^{9} (\rho_1^{loc})^i \otimes (\rho_1^{loc})^i \right)^{-1} vec(\Sigma_{10}^{loc}).$$

Estimation results.— Table A.4 summarizes the results of the estimation of this parametric DGP of location-specific risk. Panel A estimates (A.7) as multivariate system of equations. As a robustness check, Panel B estimates the corresponding univariate first-order auto-regressive regressions with correlated error terms across equations. The standard errors of the estimated auto-correlation coefficients (AR) are clustered by state in both panels. This simple first-order Markov process fits the evolution of location-specific log prices well. The R^2 for all three log prices is above 80%. The process is highly persistent but stationary, with all eigenvalues of the AR coefficient matrix ρ_{10} within the unit circle. Using steps described before to recover the implied panel VAR at annual frequency, we find that the AR(1) coefficient matrix of the annualized panel VAR is nearly an identity matrix. Its three eigenvalues are 0.99, 0.98, and 0.96. The persistence is also evident from Panel B, which shows that the annualized coefficients (clustered standard errors) of the univariate auto-regressive coefficients are 0.982 (0.0017) for log wages, 0.987 (0.0029) for

log rents, and 0.988 (0.0029) for log prices. Hence, while the process is highly persistent, it is stationary. All annualized AR coefficients in Panel B are statistically significantly smaller than one.

The error terms u_{ct}^{loc} have substantial volatility and are strongly positively correlated as expected. The standard deviation of annualized shocks to location-specific home prices is 6.1%, which is in line with previous estimates in the literature. For instance, Table 1 of the handbook chapter by Piazzesi and Schneider (2016) reports location-specific house price volatilities of 5% at the state and 7% at the city level, keeping in mind that commuting zones are larger than cities and smaller than states. The volatility of annual location-specific shocks is 3.7% for log rents and 1.7% for log wages. We are not aware of similar estimates provided in the literature. Because of the high persistence of the process, these numbers correspond closely to the annualized standard deviations obtained by dividing the standard deviation of the s-period log price change by the square-root of s (see Figure A.2 for s ranging from 10 to 70 years).

Finally, Panel C of Table A.4 shows that the estimated standard deviation of initial log prices ranges from 10–15% for wages, 27–33% for rents, and 40–42% for home prices depending on whether one uses the sample variance of the pooled log prices or the long-run variance implied by the panel VAR in equation (A.7).

Model fit.— To assess the fit of the parametric risk model, we simulate data for 720 locations over 70 decades from the DGP (equation A.7) and compare it with the empirical data using different moments. Table A.5 shows that we obtain very similar regression slopes from regressing changes in log housing costs on log wages as those reported in Figure 4 in the main text.

Table A.6 compares the first four moments of the marginal distributions of the simulated error terms with the corresponding moments of the residuals of the panel VAR (Panel A), and of the simulated and empirical log prices (Panel B). Panel A shows that the simulated growth rate shock distribution matches that of the regression residuals well, with the exception that home price residuals are more positively skewed and leptokurtic. Panel B shows that the marginal distributions of simulated and empirical data are also closely aligned, with the exception that the empirical distribution has slightly higher cross-sectional standard deviations.

Aggregate risk.—

Idiosyncratic risk.— Idiosyncratic risk is by definition independent of all other risk, not only of location-specific and aggregate risk, but also of idiosyncratic risk of the other

prices. E.g., idiosyncratic rent risk is independent of idiosyncratic home price risk and of idiosyncratic wage income risk. We therefore add idiosyncratic risk in the same way as we do in the non-parametric risk model.

C Risk Analysis

Present values.— We compute the volatility of the present value of prices under different scenarios. In the baseline, we only use location-specific price changes non-parametrically. In extensions and robustness checks, we also add aggregate and idiosyncratic price changes, and we also use a parametric model to simulate price risk.

Timing assumption.— Stocks are measured at the end of period during which flows occur. That is, owners need to own the home in period t to consume the housing service flow in period $t+1$ (i.e., home prices are "ex-dividend" prices). Renters instead buy the housing service flow in the same period they consume it (i.e., rents and wages are spot prices unknown at time t).

Wage and housing risk by tenure spell.— The present value of real wages and rents X of individual i over a T-period housing spell, starting in period t and location $c(i)$, is:

$$PV(X)_{it}^{(T)} \equiv \sum_{s=1}^{T} \frac{X_{i,t+s}}{(1+r_t)^s}. \tag{A.12}$$

Individual prices are computed using equation (A.6). r_t is the long-run risk-free interest rate in period t. In the baseline, we keep the interest rate constant at $r = 2.5\%$, the average real yield of 10-year U.S. Government bonds in the post-war period 1950–2010. As a robustness exercise, we also use the contemporaneous real risk-free yields on 10-year Government bonds in each year from 1940–2010, taken from Jordà et al. (2019).

For owners, the housing cost during this spell is the present value of the net capital loss:

$$P_{i,t} - \frac{P_{i,t+T}}{(1+r_t)^T}, \tag{A.13}$$

where prices are also computed using equation (A.6), i.e., $P_{i,t} = P^o = \$186k$ is the known price at the beginning of the planning period in year t, and $P_{i,t+T}$ is the unknown T-year ahead sale price in location $c(i)$. We then compute the standard deviation of these present values for a fixed spell-length of T years across all calendar years 1940–2010 and across all individuals in these locations, where each present value is weighted by is location's average population share from 1940–2010.

Figure 3 plots these standard deviations for all spells from 1 to 70 years for the case with no idiosyncratic and no aggregate risk, using for each location one representative individual who does not move and always owns or rents during the spell. For example, for a spell of $T = 6$ years, the standard deviation of rents is computed across all locations from year 1940 (i.e., the 720 present values $PV(R)^{(6)}_{i,1940} = \sum_{s=1941}^{1946} \frac{R_{i,s}}{(1+r_{1940})^{s-1940}}$) up to year 2004 (the 720 present values $PV(R)^{(6)}_{i,2004} = \sum_{s=2005}^{2010} \frac{R_{i,s}}{(1+r_{2004})^{s-2004}}$), a total of 46,080 present values (720 locations c times $71-6=65$ calendar years t). Similarly, the standard deviation of an owner's housing costs with a 6-year spell is computed using the 6-year discounted capital losses across locations from year 1940 (i.e., the 720 values $P^o - \frac{P_{i,1946}}{(1+r_{1940})^6}$) up to year 2004 (the 720 values $P^o - \frac{P_{i,2010}}{(1+r_{2004})^6}$).

D Descriptive Analysis: Robustness and Extensions

This section tests the robustness of our main findings of the descriptive analysis of housing risk described in Section 4.

Placebo-type Tests using Social Security Income and Utility Costs.— A reasonable concern is that the results might be driven by some spurious relation or omitted variable. To show that this is not the case, we redo the analysis using values that are largely determined at the national level, such as Social Security income, a federal program, or electricity demand, which is serviced competitively on a national grid—or at least supplied competitively across many CZs—as well as all other components of utility costs. Figure A.14 shows that location-level wage changes indeed have a negligible association with changes in these values.

Migration.— Another concern is that these correlations might be the outcome of economically selected migration and hence a reflection of increased cross-location economic segregation. This might be the case for instance if factor price changes were largely driven by migrants and the correlations would therefore not reflect the net income risk of non-moving residents. Table A.9 addresses this concern showing that the correlations are largely unchanged when we control for population growth or when we drop cross-county movers from the sample.[43]

Using Means and Medians of Residuals Instead of Location-Year Fixed Effects.— A natural concern is that fixed effects, which are means, could be sensitive to extreme observations (although the Census applies top coding which mitigates this concern) and,

[43]Movers are identified as households that report having moved to the county within 5 years prior to the survey (1 year in the 2012 5-year ACS).

more importantly, to changes in top codes over time. The latter could create spurious positive correlations in prices due to positively correlated changes in top codes across censuses. In addition, fixed effects can be difficult to interpret when including other controls. To address these concerns, we use quality-adjusted *median* prices as an alternative measure of the location-specific component. For that purpose, we estimate the same hedonic regression as in equation (A.2) but without location fixed effects:

$$X_{it}^{ipums} = \alpha_t + q_{it}'\delta_t + u_{it}. \tag{A.14}$$

We then calculate location-specific medians of the regression residuals, median$\{\widehat{u}_{it} \mid c(i), t\}$. Because these location-specific medians of the regression residuals do not have a zero mean, we also demean them for each year to control for aggregate changes: $x_{ct}^{loc} \equiv$ median$\{\widehat{u}_{it} \mid c(i), t\} - E\Big(\text{median}\{\widehat{u}_{it} \mid c(i), t\}\Big|t\Big)$. Panel B of Figure A.15 shows that these approaches yield results that are almost identical to using the location-year fixed effects of the main specification (A.2); see panel A of Figure 4.

Housing Market Segmentation and Differences Across the Distribution.— Related to the previous concern, one might worry that the relationship between wage growth and housing price growth only applies to the typical home, but not to the entire housing market. However, Figure A.15 shows that these correlations apply across the entire cross-sectional distribution of wages and housing costs. It shows that the 25th, 50th, and 75th percentiles of the wage and housing cost distribution—obtained from estimating equation (A.14)—move similar to each other and to the conditional mean captured by location-year fixed effects shown in Figure 4.

Housing market for different types of homes might be segmented within a location. For instance, it is conceivable that prices of different types of homes—say two and four room housing units—are set independently from each other, and the hedonic regression used in our main analysis might not sufficiently account for such market segmentation. To assess whether our findings are sensitive to such frictions, Table A.9 limits the sample of housing units to single-family homes and obtains similar results. We then further restrict it to single-family homes of modal size of five rooms as well as to the modal four-room home for renters, respectively the modal six-room home for owners. The table shows that the correlations between wages and housing costs are very similar as when we use the full sample of housing units.

We also extend the hedonic regressions to control for differential changes in observed quantity and quality of the housing stock and the workforce within a location over time by projecting off changes in average home characteristics and average worker age

and education within a location over time across two consecutive Decennial Censuses, $\Delta q_{ct} = q_{ct} - q_{c,t-10}$. The table shows that the resulting relationship between wages and housing costs is similar to the baseline specification shown in Figure 4.

Changes in Unobserved Quality and Quantity.— A related reasonable concern are unobserved changes in quality and quantity. We assess this concern using two extreme approaches. First, Table A.9 shows that using the unadjusted local median price, thereby not adjusting for the characteristics of homes and workers, has little effect on the estimated correlations. While the hedonic regressions reduce the cross-sectional dispersion of full-time wages, home prices, and rents—especially within locations—substantially, they have little effect on the *growth rates* of these prices.

Second, as noted above, the relationship between market wages and housing prices remains strong even after controlling for changes in a location's average home and worker characteristics between two consecutive Decennial Censuses. Nevertheless, one might still be concerned that some unobserved quality changes might not be fully absorbed by year fixed effects, but might reflect differential changes in housing quality across space. For example, the adoption of air conditioning might have been unequal across locations in the early 20th century. To address this, we also include state fixed effects. While these fixed effects also absorb location-specific risk that we are interested in (e.g., the decline of the automotive industry was geographically concentrated), they flexibly control for such remaining unobserved quality changes. Table A.9 shows that the strong positive relationship between wages and housing costs also holds using variation across cities within the same state, using either CZ-level or county-level data.

Mortgages.—

Property Taxes and Other Carrying Costs.— Figure A.16 shows that that property tax rate changes—an important component of carrying costs—are uncorrelated with local income changes. We use this evidence and the fact that other costs such as utilities also do not correlate with local income (Figure A.14) to justify excluding carrying costs from our main analysis.

Dropping the 2000s Housing Boom and the 1940 Census.— Table A.9 drops the 2000s to test whether the wage and housing costs correlations are driven by the recent housing boom and bust cycle and the Great Recession. Excluding this decade does not affect the correlations much.

The 1950 Decennial Census lacks information on housing costs. In the main analysis, we

interpolate values between 1940 and 1960, effectively giving this period the weight of two decades. Table A.9 shows similar results after dropping the 1940 Census and limiting the sample to 1960-2010, thereby dropping much of the post-war economic boom.

Decade-by-decade.— Table A.10 shows the relationship between wage growth and housing price growth separately for each decade, both at the CZ-level and at the county level. We observe a robust positive relationship between wage risk and housing risk in all decades except the 1970s. The disconnect in the 1970s is consistent with recent work by Agorastos et al. (2024) who compare Census data with newly collected historical listing of rents and home prices in newspaper of major cities. One possibility for the weaker relationship in the 1970s might be related to the high inflation during that period, which might have dampened the reaction of relative prices to location-specific shocks or made top-coding issues more binding in some cities, even when using median values.

Alternative Definition of Local Labor and Housing Markets.— Table A.9 also shows that the correlations are robust to using alternative geographic definitions of labor and housing markets. The strength of the correlation between wages and housing costs is similar at the county and the CZ level, both for the full sample period from 1940–2010 as well as for the more recent sample periods from 1960–2010 that does not require any imputation of quality-adjusted prices.

Alternative Income Concept: Family Income.— Table A.9 extends the analysis to total family income, which reflects both changes in the market price of labor as well as changes in labor supply (hours and labor market participation), financial income, and government transfers. Total family income is similarly correlated with housing costs as are full-time wages. Hence, housing tenure choice provides households with an important opportunity to manage income risk that is not already covered by other financial assets in the household's portfolio or by the government (e.g., by social insurance such as unemployment insurance, old-age pensions, or disability insurance).

Alternative Rent Concept: Gross Rent.— Table A.9 shows that the correlation of changes in wages and gross rents—which include utility costs—is attenuated relative to the correlation when using contract rents. The reason is that these additional housing costs are largely uncorrelated with location-specific wage changes as shown in Figure A.14.

Large and Small Commuting Zones.— Finally, Table A.9 shows that the relationship between wages and housing costs tends to be stronger in larger CZs, consistent with lower housing supply elasticities in larger CZs.

E Risk Cost Calculations

F Life Cycle Model

Table A.1: Volatility of National-Level Housing Cost and Income Growth

	Std of annualized log changes over:					
	1 year	2 years	3 years	4 years	5 years	Period
Panel A: Rent						
BLS: City average rent of primary residence	1.4%	1.7%	1.8%	1.9%	2.0%	1979–2021
	2.4%	3.0%	3.2%	3.4%	3.5%	1940–2021
BEA: Rental index of tenant-occupied housing	1.6%	1.7%	1.8%	1.9%	2.0%	1979–2021
	2.7%	3.3%	3.5%	3.5%	3.7%	1940–2021
Panel B: Home price						
FHFA: All-transactions house price index	4.0%	5.2%	5.8%	6.1%	6.3%	1979–2021
	3.9%	5.1%	5.8%	6.1%	6.2%	1975–2021
Shiller: Repeat sales house price index	5.2%	6.8%	7.7%	8.2%	8.3%	1979–2021
	5.1%	6.3%	6.8%	7.1%	7.0%	1940–2021
Panel C: Income						
CPS: Full-time wage and salaray income	1.8%	2.2%	2.4%	2.4%	2.4%	1979–2020
	3.2%	3.0%	3.2%	3.3%	3.4%	1961–2020
BLS: Usual weekly real full-time earnings	1.7%	1.8%	1.8%	1.7%	1.7%	1979–2021

Notes. The table shows the standard deviation of annualized log changes (or growth rates) of national wage income, rents, and home prices over different horizons. Annualized log changes are defined as $\Delta_h \ln x^{agg}_{t+h} / \sqrt{h}$, with $\Delta_h \ln x^{agg}_{t+h} = \ln x^{agg}_{t+h} - \ln x^{agg}_{t}$. The first row of each series uses the common sample period 1979–2021; the second row uses data back to 1940 or to earliest year available for the series. All nominal prices are deflated using the CPI for All Urban Consumers (U.S. Bureau of Labor Statistics 1913–2021a). Using the PCE Price Index (U.S. Bureau of Economic Analysis 1929–2021b) or the Gross Domestic Product (GDP) Implicit Price Deflator (U.S. Bureau of Economic Analysis 1947–2021a) produces similar results. Figure A.4 and Table A.2 provide additional information.
Source: *Panel A, Rent.* Rows one and two (BLS) measure rents using the Consumer Price Index (CPI) "rent of primary residence in U.S. city average" series (U.S. Bureau of Labor Statistics 1914–2021b) provided by the BLS. Rows three and four (BEA) instead measure rents using the Personal Consumption Expenditures (PCE) series "rental chain-type price index of tenant-occupied nonfarm housing" provided by the BEA (U.S. Bureau of Economic Analysis 1929–2021c). *Panel C: Home price.* Rows five and six (FHFA) measure home prices using the "all-transactions house price index for the United States" series provided by the FHFA (U.S. Federal Housing Finance Agency 1975–2021). Rows seven and eight (Shiller) instead measure home prices using real U.S. home price data for Figure 3.1 in Robert J. Shiller's Irrational Exuberance (Shiller 1890–2021). *Panel A: Income.* Rows nine and ten (CPS) measure wage income using the IPUMS ASEC microdata (i.e., the Annual Social and Economic (ASEC) supplement to the Current Population Survey (CPS); Flood et al. 2021) to define full-time wage income the same way as described in Section 3 and A for the IPUMS-USA Decennial Census microdata (Ruggles et al. 2018). The IPUMS-CPS ASEC microdata is available from 1962–2021, with work and income questions referencing the previous calendar year, hence CPS full-time wage income ends in 2020. Row eleven (BLS) instead measures wage income using the series "median usual weekly real earnings of wage and salary workers 16 years and over employed full time" provided by the BLS (U.S. Bureau of Labor Statistics 1979–2021c). This series is based on the CPS "merged outgoing rotation group" (or "earner study" or "annual earnings files") of the basic monthly samples, which began in 1979. Work and income questions reference current pay, hence this series ends in 2021.

Table A.2: Correlation of National-Level Growth in Housing Costs and Income

		Income:		Rent:		Home price:	
		CPS	BLS	BLS	BEA	FHFA	Shiller
Panel A: 1-year log changes							
Income	CPS: Full-time wage and salaray income	100%					
	BLS: Usual weekly real earnings	53%	100%				
Rent	BLS: City average rent of primary residence	80%	72%	100%			
	BEA: Rental index of tenant-occupied housing	61%	73%	92%	100%		
HP	FHFA: All-transactions house price index	60%	41%	58%	46%	100%	
	Shiller: Repeat sales price index	52%	27%	46%	33%	92%	100%
Panel B: 3-year log changes							
Income	CPS: Full-time wage and salaray income	100%					
	BLS: Usual weekly real earnings	75%	100%				
Rent	BLS: City average rent of primary residence	87%	76%	100%			
	BEA: Rental index of tenant-occupied housing	68%	70%	92%	100%		
HP	FHFA: All-transactions house price index	62%	41%	59%	45%	100%	
	Shiller: Repeat sales price index	58%	30%	54%	40%	95%	100%
Panel C: 5-year log changes							
Income	CPS: Full-time wage and salaray income	100%					
	BLS: Usual weekly real earnings	82%	100%				
Rent	BLS: City average rent of primary residence	88%	79%	100%			
	BEA: Rental index of tenant-occupied housing	72%	69%	93%	100%		
HP	FHFA: All-transactions house price index	59%	47%	56%	47%	100%	
	Shiller: Repeat sales price index	61%	41%	59%	49%	97%	100%

Notes. The table shows the correlation of log changes (or growth rates) of national wage income, rents, and home prices, $\Delta_h \ln x^{agg}_{t+h} = \ln x^{agg}_{t+h} - \ln x^{agg}_{t}$, over different horizons $h \in \{1, 3, 5\}$ from 1979–2021. All nominal prices are deflated using the CPI for All Urban Consumers (U.S. Bureau of Labor Statistics 1913–2021a). Using the PCE Price Index (U.S. Bureau of Economic Analysis 1929–2021b) or the Gross Domestic Product (GDP) Implicit Price Deflator (U.S. Bureau of Economic Analysis 1947–2021a) produces similar results. Figure A.4 and Table A.1 provide additional information.

Source: *Income.* "Income CPS" measures wage income using the IPUMS ASEC microdata (i.e., the Annual Social and Economic (ASEC) supplement to the Current Population Survey (CPS); Flood et al. 2021) to define full-time wage income the same way as described in Section 3 and A for the IPUMS-USA Decennial Census microdata (Ruggles et al. 2018). The IPUMS-CPS ASEC microdata is available from 1962–2021, with work and income questions referencing the previous calendar year, hence CPS full-time wage income ends in 2020. "Income BLS" measures wage income using the series "median usual weekly real earnings of wage and salary workers 16 years and over employed full time" provided by the BLS (U.S. Bureau of Labor Statistics 1979–2021c). This series is based on the CPS "merged outgoing rotation group" (or "earner study" or "annual earnings files") of the basic monthly samples, which began in 1979. Work and income questions reference current pay, hence this series ends in 2021. *Rent.* "Rent BLS" measures rents using the Consumer Price Index (CPI) "rent of primary residence in U.S. city average" series (U.S. Bureau of Labor Statistics 1914–2021b) provided by the BLS. "Rent BEA" measures rents using the Personal Consumption Expenditures (PCE) series "rental chain-type price index of tenant-occupied nonfarm housing" provided by the BEA (U.S. Bureau of Economic Analysis 1929–2021c). *Home price (HP).* "HP FHFA" measures home prices using the "all-transactions house price index for the United States" series provided by the FHFA (U.S. Federal Housing Finance Agency 1975–2021). "HP Shiller" measures. home prices using real U.S. home price data for Figure 3.1 in Robert J. Shiller's Irrational Exuberance (Shiller 1890–2021).

Table A.3: Volatility and Correlation of Cross-Country Growth in Housing Costs and Income

	Spell length, T (years): 1	3	5	10	30	70	Period
Panel A: Standard deviation of annualized log changes (%)							
Rent, $(\ln R_{\omega,t+T} - \ln R_t)/\sqrt{T}$	6.4%	8.3%	9.3%	9.9%	7.7%	–	1940–2010
	6.3%	8.5%	9.5%	10.5%	10.8%	10.0%	1870–2015
Home price, $(\ln P_{\omega,t+T} - \ln P_t)/\sqrt{T}$	7.2%	9.1%	9.7%	9.9%	8.8%	–	1940–2010
	8.0%	9.2%	9.6%	9.7%	10.4%	8.2%	1870–2015
Income, $(\ln y_{\omega,t+T} - \ln y_t)/\sqrt{T}$	4.2%	5.2%	4.9%	4.7%	5.0%	–	1940–2010
	4.2%	5.0%	5.3%	5.5%	6.7%	5.3%	1870–2015
Panel B: Correlation of housing costs with income (log changes, in %)							
Rent, $Corr(\ln R_{\omega,t+T} - \ln R_t, \ln y_{\omega,t+T} - \ln y_t)$	15%	26%	36%	43%	57%	–	1940–2010
	12%	22%	33%	45%	64%	55%	1870–2015
Home price, $Corr(\ln P_{\omega,t+T} - \ln P_t, \ln y_{\omega,t+T} - \ln y_t)$	8%	9%	17%	32%	49%	–	1940–2010
	11%	22%	31%	45%	66%	63%	1870–2015
Panel C: Coefficients of regressions, $\Delta_T \ln x_{\omega,t+T} = \alpha + \beta_{x\|y}\Delta_T \ln y_{\omega,t+T} + \epsilon_{\omega,t+T}$							
Rent, $\beta_{R\|y}$	0.14	0.30	0.46	0.70	0.92	–	1940–2010
	(0.11)	(0.189)	(0.21)	(0.29)	(0.26)	–	
	0.18	0.38	0.59	0.87	1.04	1.04	1870–2015
	(0.17)	(0.25)	(0.26)	(0.29)	(0.19)	(0.32)	
Home price, $\beta_{P\|y}$	0.19	0.29	0.39	0.61	0.96	–	1940–2010
	(0.19)	(0.22)	(0.23)	(0.26)	(0.28)	–	
	0.21	0.41	0.56	0.80	1.02	0.96	1870–2015
	(0.21)	(0.26)	(0.26)	(0.30)	(0.28)	(0.21)	

Notes. Aggregate risk of real (purchasing power parity adjusted) incomes (GDP per capita), rents, and home prices by country. Log changes are winsorized at the 1st and 99th percentile of the unweighted distribution. The data set includes the following 16 countries: Australia, Belgium, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, the UK, and the US. Each statistic is computed on two samples: on the sample period 1940–2010 that matches the rest of the analysis and on the entire sample from 1870–2015. The latter imputes missing 1-year growth rates in interim periods. This is the data used for the non-parametric model of aggregate risk described in Section B.1. Statistics for the 1940–2010 sample use weights corresponding to country population size relative to total population in each year that add up to the number of countries in each year of the unbalanced panel so that each year has the same weight in the calculations. Weights for statistics on the entire sample 1870–2015 are population shares that are constant across all years and are computed over the period where all 16 countries are present (mostly years 1949–2015). All statistics shown use at least 100 observations. Since 70-year log changes over the 1940–2010 period are available for only 9 countries, these statistics are not show in column 6. *Panel A.* Standard deviation of annualized log changes of aggregate income and housing prices, $\Delta_T \ln x^{agg}_{t+t}/\sqrt{T}$, with T-period log change $\Delta_T \ln x^{agg}_{t+T} = \ln x^{agg}_{t+T} - \ln x^{agg}_t$. *Panel B.* Correlation coefficients $Corr(\Delta_T \ln x^{agg}_{t+T}, \Delta_T \ln y^{agg}_{t+T})$. *Panel C.* Elasticities $\beta_{wls,x|y}(T)$ are obtained from regressing log changes in housing costs on log changes in GDP per capita , weighted by population: $\Delta_T \ln x^{agg}_{t+T} = a + \beta_{wls,x|y}(T) \cdot \Delta_T \ln y^{agg}_{t+T} + u_{t+T}$. Robust standard errors in brackets are two-way clustered by country and year. Unweighted regressions yield similar results. Source: Jordà et al. (2019).

Table A.4: DGP of Location-Specific Risk for Life Cycle Model

Panel A: Multivariate regression, $\ln x_{ct}^{loc} = \rho_{10}^{loc} \ln x_{c,t-10}^{loc} + u_{ct}^{loc}$, $u_{ct}^{loc} \sim i.i.d.\ N(0, \Sigma_{10}^{loc})$

	AR coefficient matrix, $(\rho_{10}^{loc})'$:			Covariance matrix, Σ_{10}^{loc}:		
	$\ln y_{ct}^{loc}$	$\ln R_{ct}^{loc}$	$\ln P_{ct}^{loc}$	$u_{ct}^{y,loc}$	$u_{ct}^{R,loc}$	$u_{ct}^{P,loc}$
Income, $\ln y_{c,t-10}^{loc}$	0.706	-0.196	-0.360	0.0023		
	(0.0243)	(0.0573)	(0.124)	[0.0477]		
Rent, $\ln R_{c,t-10}^{loc}$	0.0606	0.887	0.333	0.0028	0.0125	
	(0.0202)	(0.0399)	(0.0874)		[0.112]	
Home price, $\ln P_{c,t-10}^{loc}$	0.0108	0.0572	0.749	0.0049	0.0153	0.0323
	(0.0142)	(0.0379)	(0.0615)			[0.180]
R^2 (univariate)	0.875	0.864	0.801	Annualized Std., $\sqrt{\sigma_{ii}^2/10}$:		
Eigenvalues of ρ_{10}^{loc}	0.880	0.800	0.661	0.0151	0.0353	0.0568
	Annualized AR matrix, $(\rho_1^{loc})'$:			Annualized cov, Σ_1^{loc}:		
	$\ln y_{ct}^{loc}$	$\ln R_{ct}^{loc}$	$\ln P_{ct}^{loc}$	$v_{ct}^{y,loc}$	$v_{ct}^{R,loc}$	$v_{ct}^{P,loc}$
Income, $\ln y_{c,t-1}^{loc}$	0.967	-0.023	-0.043	0.00028		
				[0.0166]		
Rent, $\ln R_{c,t-1}^{loc}$	0.007	0.988	0.042	0.00030	0.00135	
					[0.0368]	
Home price, $\ln P_{c,t-1}^{loc}$	0.001	0.007	0.970	0.00056	0.00158	0.00375
						[0.0613]

Panel B: Univariate regressions, $\ln x_{ct}^{i,loc} = \rho_{i,10}^{loc} \ln x_{c,t-10}^{i,loc} + u_{ct}^{i,loc}$, $(u_{ct}^{w,loc}, u_{ct}^{R,loc}, u_{ct}^{P,loc})' \sim i.i.d.\ N(0, \Sigma_{10}^{loc})$

	AR coefficients:			Covariance matrix, Σ_{10}^{loc}:		
	$\ln y_{ct}^{loc}$	$\ln R_{ct}^{loc}$	$\ln P_{ct}^{loc}$	$u_{ct}^{y,loc}$	$u_{ct}^{R,loc}$	$u_{ct}^{P,loc}$
Income, $\ln y_{c,t-10}^{loc}$	0.836			0.0025		
	(0.0145)			[0.0494]		
Rent, $\ln R_{c,t-10}^{loc}$		0.875		0.0030	0.0128	
		(0.0258)			[0.113]	
Home price, $\ln P_{c,t-10}^{loc}$			0.887	0.0053	0.0155	0.0339
			(0.0354)			[0.184]
R^2	0.866	0.860	0.791			
	Annualized AR coefficients:			Annualized Std., $\sqrt{\sigma_{ii}^2/10}$:		
	0.982	0.987	0.988	0.0156	0.0358	0.0582

Panel C: Variance Σ_0 *of initial log prices,* $\ln x_{c,0}^{loc} = u_{c,0}^{loc} \sim N(0, \Sigma_0^{loc})$

	Sample variance, $Var(\ln x_{ct}^{loc})$:			pVAR's long-run variance:		
	$u_{c,0}^{y,loc}$	$u_{c,0}^{R,loc}$	$u_{c,0}^{P,loc}$	$u_{c,0}^{y,loc}$	$u_{c,0}^{R,loc}$	$u_{c,0}^{P,loc}$
Income, $u_{c,0}^{y,loc}$	0.022			0.010		
	[0.147]			[0.099]		
Rent, $u_{c,0}^{R,loc}$	0.040	0.106		0.021	0.073	
		[0.325]			[0.270]	
Home price, $u_{c,0}^{P,loc}$	0.047	0.124	0.177	0.030	0.099	0.159
			[0.420]			[0.398]

Notes. The sample consists of 5,040 CZ-year observations from 1940–2010. Robust standard errors in brackets are clustered by state. Numbers in square brackets are the standard deviations σ_{ii} of the 10-year error term u_{ct}^{loc}. See Appendix B.2 for details. Source: Decennial Census and ACS.

Table A.5: Fit of Location-Specific Risk Model: Regression Slopes

	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Log price changes, $\Delta_s \ln x_{ct}^{loc} = \beta_s \Delta_s \ln \chi_{ct}^{loc} + u_{ct}$						
	$\Delta \ln R$ on $\Delta \ln w$:		$\Delta \ln P$ on $\Delta \ln w$:		$\Delta \ln P$ on $\Delta \ln R$:	
	Model	Data	Model	Data	Model	Data
β_{10}	1.250***	1.245***	1.881***	1.960***	1.159***	1.143***
	(0.024)	(0.053)	(0.040)	(0.269)	(0.015)	(0.071)
R^2	0.351	0.327	0.322	0.326	0.544	0.525
N	5,040	5,040	5,040	5,040	5,040	5,040
β_{20}	1.243***	1.312***	1.693***	1.390***	1.162***	1.040***
	(0.028)	(0.135)	(0.049)	(0.246)	(0.017)	(0.079)
R^2	0.363	0.305	0.280	0.192	0.561	0.608
N	4,320	4,320	4,320	4,320	4,320	4,320
β_{30}	1.287***	1.279***	1.576***	1.464***	1.144***	1.173***
	(0.034)	(0.141)	(0.060)	(0.303)	(0.021)	(0.082)
R^2	0.390	0.333	0.255	0.210	0.570	0.663
N	3,600	3,600	3,600	3,600	3,600	3,600
β_{40}	1.310***	1.184***	1.509***	1.282***	1.139***	1.239***
	(0.038)	(0.146)	(0.069)	(0.283)	(0.024)	(0.104)
R^2	0.424	0.318	0.252	0.171	0.581	0.703
N	2,880	2,880	2,880	2,880	2,880	2,880
β_{50}	1.328***	1.183***	1.429***	1.044***	1.122***	1.146***
	(0.042)	(0.145)	(0.074)	(0.259)	(0.027)	(0.124)
R^2	0.459	0.343	0.249	0.129	0.591	0.635
N	2,160	2,160	2,160	2,160	2,160	2,160
β_{60}	1.337***	1.299***	1.316***	1.029***	1.098***	1.050***
	(0.046)	(0.129)	(0.082)	(0.241)	(0.030)	(0.129)
R^2	0.486	0.458	0.233	0.150	0.596	0.577
N	1,440	1,440	1,440	1,440	1,440	1,440
β_{70}	1.358***	1.439***	1.255***	1.185***	1.094***	1.044***
	(0.051)	(0.128)	(0.093)	(0.270)	(0.034)	(0.137)
R^2	0.494	0.541	0.217	0.183	0.614	0.545
N	720	720	720	720	720	720
Panel B: Log prices (initial condition), $\ln x_{ct}^{loc} = \beta_0 \ln \chi_{ct}^{loc} + u_{ct}$						
	$\ln R$ on $\ln w$:		$\ln P$ on $\ln w$:		$\ln P$ on $\ln R$:	
	Model	Data	Model	Data	Model	Data
β_0	1.696***	1.853***	2.229***	2.166***	1.358***	1.173***
	(0.047)	(0.112)	(0.077)	(0.226)	(0.016)	(0.073)
R^2	0.452	0.700	0.340	0.572	0.802	0.824
N	5,760	5,760	5,760	5,760	5,760	5,760

Notes. The empirical data (data columns) consists of 5,760 CZ-year observations from 1940–2010. The model data (model columns) simulates 8 decadal observations for each of 720 locations from the estimated data generating process (A.11). Robust standard errors in brackets are clustered by the 48 continental states for the empirical data, and by each of the 720 locations for the simulated data. Regressions with the empirical data use population weights. See Appendix B.2 and the notes of Table A.4 for details. Source: Decennial Census and ACS.

Table A.6: Marginal Distribution of Simulated and Empirical Location-Specific Data

	N	Mean	StDev	Skewness	Kurtosis
Panel A: Model error $u_{ct}^{loc} \sim$ i.i.d. $N(0, \Sigma_{10}^{loc})$ vs. regression residual $\widehat{u}_{ct}^{loc}$					
		Simulated data:			
Income, $u_{ct}^{y,loc}$	5,040	0.000	0.048	0.041	2.877
Rent, $u_{ct}^{R,loc}$	5,040	0.002	0.113	0.008	3.046
Home price, $u_{ct}^{P,loc}$	5,040	0.002	0.181	0.038	2.965
		Empirical data:			
Income, $\widehat{u}_{ct}^{y,loc}$	5,040	0.000	0.046	0.108	3.408
Rent, $\widehat{u}_{ct}^{R,loc}$	5,040	0.000	0.108	-0.098	2.964
Home price, $\widehat{u}_{ct}^{P,loc}$	5,040	0.001	0.177	0.729	4.250
Panel B: Distribution of initial log prices (initial condition), $\ln x_{c,0}^{loc}$					
		Simulated data:			
Income, $\ln y_{c,0}^{loc}$	5,760	-0.029	0.109	0.202	3.581
Rent, $\ln R_{c,0}^{loc}$	5,760	-0.012	0.274	-0.013	3.070
Home price, $\ln P_{c,0}^{loc}$	5,760	0.029	0.416	-0.079	3.093
		Empirical data:			
Income, $\ln y_{c,0}^{loc}$	5,760	0.001	0.143	-0.607	2.990
Rent, $\ln R_{c,0}^{loc}$	5,760	0.000	0.320	-0.308	2.706
Home price, $\ln P_{c,0}^{loc}$	5,760	0.001	0.413	0.305	2.686

Notes. Panel A. The empirical data is the distribution of regression residuals obtained from the panel VAR, equation (A.7): $\ln x_{ct}^{loc} = \rho_{10}^{loc} \ln x_{c,t-10}^{loc} + u_{ct}^{loc}$ for $t \in \{10, 20, \ldots, 70\}$. The simulated data consists of decadal growth rate shocks drawn from the multivariate normal distribution (A.8): $u_{ct}^{loc} \sim$ *i.i.d.* $N(0, \widehat{\Sigma}_{10}^{loc})$, where $\widehat{\Sigma}_{10}^{loc}$ is the estimated covariance matrix of the panel VAR. *Panel B.* The empirical data is the distribution of location-year fixed effects $\widehat{\lambda}_{ct}$ of the hedonic regression (A.3). The simulated data consists of initial log prices drawn from the multivariate normal distribution (A.9): $u_{c,0}^{loc} \sim N(0, \widehat{\Sigma}_{0}^{loc})$, where $\widehat{\Sigma}_{0}^{loc}$ is the covariance matrix (A.10) of the estimated location-year fixed effects. The empirical distributions in both panels use population weights and have been winsorized at the 1% and 99% percentile to reduce the influence of extreme values. See Appendix B.2 for details. Source: Decennial Census and ACS.

Table A.7: Table 1 with Simulated and Empirical Data

	Spell length, T (years):						
	10	20	30	40	50	60	70
Panel A: Standard deviation of price changes ($1,000)							
	Simulated data:						
Rent, $R_{\omega,t+T} - R_t$	1.5	2.1	2.6	3.0	3.3	3.7	4.0
Home price, $P_{\omega,t+T} - P_t$	35.8	49.2	59.0	66.4	72.7	78.2	83.6
Income, $y_{\omega,t+T} - y_t$	3.3	4.7	5.7	6.7	7.7	8.5	9.1
	Empirical data:						
Rent, $R_{\omega,t+T} - R_t$	1.6	2.4	2.8	3.0	3.3	3.7	4.2
Home price, $P_{\omega,t+T} - P_t$	37.8	51.9	64.5	70.2	72.0	77.5	90.0
Income, $y_{\omega,t+T} - y_t$	3.2	4.6	5.7	6.6	7.4	8.5	9.7
Panel B: Standard deviation of annualized log price changes (%)							
	Simulated data:						
Rent, $(\ln R_{\omega,t+T} - \ln R_t)/\sqrt{T}$	3.8%	3.7%	3.7%	3.7%	3.7%	3.7%	3.6%
Home price, $(\ln P_{\omega,t+T} - \ln P_t)/\sqrt{T}$	6.0%	5.8%	5.6%	5.5%	5.3%	5.2%	5.1%
Income, $(\ln y_{\omega,t+T} - \ln y_t)/\sqrt{T}$	1.8%	1.8%	1.8%	1.8%	1.9%	1.9%	1.9%
	Empirical data:						
Rent, $(\ln R_{\omega,t+T} - \ln R_t)/\sqrt{T}$	3.8%	4.1%	3.8%	3.6%	3.5%	3.5%	3.7%
Home price, $(\ln P_{\omega,t+T} - \ln P_t)/\sqrt{T}$	6.0%	5.4%	5.5%	5.4%	5.0%	4.8%	5.2%
Income, $(\ln y_{\omega,t+T} - \ln y_t)/\sqrt{T}$	1.7%	1.7%	1.7%	1.7%	1.7%	1.8%	1.9%
Panel C: Standard deviation of present values ($1,000)							
Prices:	Simulated data:						
Rent, $R_\omega = \sum_{s=1}^{T} \frac{R_{\omega,t+s}}{(1+r_t)^s}$	6.1	16.4	26.9	36.5	45.2	52.8	62.4
Own, $O_\omega = P_t - \frac{P_{\omega,t+T}}{(1+r_t)^T}$	22.5	27.1	26.2	23.5	20.2	17.1	14.8
Income, $Y_\omega = \sum_{s=1}^{T} \frac{y_{\omega,t+s}}{(1+r_t)^s}$	13.1	36.0	60.4	84.9	108.8	128.3	143.1
	Empirical data:						
Rent, $R_\omega = \sum_{s=1}^{T} \frac{R_{\omega,t+s}}{(1+r_t)^s}$	6.5	18.7	31.3	41.8	51.6	69.1	93.7
Own, $O_\omega = P_t - \frac{P_{\omega,t+T}}{(1+r_t)^T}$	23.3	27.5	27.3	23.6	19.1	16.7	16.0
Income, $Y_\omega = \sum_{s=1}^{T} \frac{y_{\omega,t+s}}{(1+r_t)^s}$	12.8	35.5	58.4	81.8	110.2	143.0	180.5
Net Income:	Simulated data:						
Renter, $Y_\omega - R_\omega$	10.7	29.1	48.4	67.4	85.7	100.8	112.5
Owner, $Y_\omega - O_\omega$	31.5	54.5	76.0	97.8	119.1	136.3	149.6
Full housing insurance, $Y_\omega - E(R_\omega)$	13.1	36.0	60.4	84.9	108.8	128.3	143.1
	Empirical data:						
Renter, $Y_\omega - R_\omega$	10.6	29.3	49.1	70.2	89.7	108.3	131.7
Owner, $Y_\omega - O_\omega$	31.1	50.2	69.9	87.4	113.6	146.2	183.6
Full housing insurance, $Y_\omega - E(R_\omega)$	12.8	35.5	58.4	81.8	110.2	143.0	180.5

Notes. See notes of Table 1 and Appendix B.2 for details. Source: Decennial Census and ACS.

Table A.8: Table 2 with Simulated and Empirical Data

	Spell length, T (years): 10	20	30	40	50	60	70
Panel A: Dollar price changes							
	Simulated data:						
Rent, $Corr(R_{\omega,t+T} - R_t, y_{\omega,t+T} - y_t)$	0.59	0.60	0.62	0.64	0.67	0.69	0.70
Home price, $Corr(P_{\omega,t+T} - P_t, y_{\omega,t+T} - y_t)$	0.57	0.53	0.50	0.50	0.49	0.47	0.45
	Empirical data:						
Rent, $Corr(R_{\omega,t+T} - R_t, y_{\omega,t+T} - y_t)$	0.57	0.55	0.58	0.54	0.57	0.66	0.71
Home price, $Corr(P_{\omega,t+T} - P_t, y_{\omega,t+T} - y_t)$	0.54	0.38	0.39	0.32	0.27	0.28	0.30
Panel B: Log price changes							
	Simulated data:						
Rent, $Corr(\ln R_{\omega,t+T} - \ln R_t, \ln y_{\omega,t+T} - \ln y_t)$	0.59	0.60	0.62	0.65	0.68	0.70	0.70
Home price, $Corr(\ln P_{\omega,t+T} - \ln P_t, \ln y_{\omega,t+T} - \ln y_t)$	0.57	0.53	0.50	0.50	0.50	0.48	0.47
	Empirical data:						
Rent, $Corr(\ln R_{\omega,t+T} - \ln R_t, \ln y_{\omega,t+T} - \ln y_t)$	0.57	0.55	0.58	0.56	0.59	0.68	0.74
Home price, $Corr(\ln P_{\omega,t+T} - \ln P_t, \ln y_{\omega,t+T} - \ln y_t)$	0.57	0.44	0.46	0.41	0.36	0.39	0.43
Panel C: Present values of prices							
	Simulated data:						
Rent, $Corr(R_\omega, Y_\omega)$	0.60	0.61	0.62	0.64	0.66	0.67	0.66
Own, $Corr(O_\omega, Y_\omega)$	-0.53	-0.48	-0.46	-0.46	-0.44	-0.42	-0.40
	Empirical data:						
Rent, $Corr(R_\omega, Y_\omega)$	0.57	0.56	0.54	0.51	0.59	0.68	0.71
Own, $Corr(O_\omega, Y_\omega)$	-0.43	-0.26	-0.23	-0.10	-0.09	-0.14	-0.16
Panel D: Coefficients of reverse regressions of present values, $Y_\omega = \alpha + \beta_{Y\|H} H_\omega + \epsilon_\omega$							
	Simulated data:						
Rent, $\beta_{Y\|R}$	1.29	1.33	1.40	1.50	1.60	1.63	1.50
	(0.02)	(0.03)	(0.04)	(0.04)	(0.05)	(0.06)	(0.06)
Own, $\beta_{Y\|O}$	-0.31	-0.64	-1.06	-1.65	-2.39	-3.16	-3.82
	(0.01)	(0.02)	(0.04)	(0.08)	(0.14)	(0.19)	(0.28)
	Empirical data:						
Rent, $\beta_{Y\|R}$	1.12	1.07	1.01	1.01	1.27	1.41	1.37
	(0.08)	(0.10)	(0.13)	(0.17)	(0.17)	(0.15)	(0.13)
Own, $\beta_{Y\|O}$	-0.24	-0.34	-0.49	-0.36	-0.53	-1.16	-1.77
	(0.04)	(0.14)	(0.24)	(0.32)	(0.50)	(0.73)	(0.94)

Notes. See notes of Table 2 and Appendix B.2 for detail. Source: Decennial Census and ACS.

Table A.9: Robustness and Extensions of Descriptive Analysis (Location-Specific Risk)

	10-year wage elasticity of:					
	Rents			Home prices		
	β	SE(β)	R^2	β	SE(β)	R^2
Panel A: Baseline specification for comparison						
Specification shown in Figure 4, panel A	1.24	(0.05)	0.33	1.96	(0.27)	0.33
Panel B: Robustness checks and extensions						
Controlling for relative population growth	1.22	(0.05)	0.31	1.93	(0.27)	0.31
Drop cross-county movers (median prices, 1960–2010)	0.85	(0.12)	0.22	1.85	(0.26)	0.33
Using median housing cost of modal home instead of hedonic regressions (1960–2010):						
Single family homes	1.05	(0.08)	0.22	1.95	(0.28)	0.34
Modal five-room home	0.92	(0.11)	0.21	1.93	(0.28)	0.31
Modal four-room home for renters, six-room home for owners	0.97	(0.14)	0.25	1.98	(0.28)	0.33
Controlling for changes in the composition of homes and workers (1960–2010)	1.01	(0.08)	0.31	1.80	(0.20)	0.33
Using unadjusted location median prices	1.20	(0.06)	0.36	1.59	(0.24)	0.24
Controlling for state fixed effects:						
Using CZ-level data	1.19	(0.05)	0.30	2.04	(0.29)	0.34
Using county-level data	1.08	(0.04)	0.29	1.61	(0.19)	0.29
Dropping 1940 (sample period 1960–2010)	0.96	(0.10)	0.26	2.04	(0.29)	0.36
Dropping 2000s boom-bust cycle (1940–2000)	1.30	(0.06)	0.34	1.98	(0.31)	0.33
Using counties instead of CZs (1940–2010)	1.11	(0.04)	0.31	1.58	(0.18)	0.29
Using counties instead of CZs (1960–2010)	0.93	(0.07)	0.25	1.66	(0.21)	0.30
Using family income (1960–2010)	0.75	(0.11)	0.27	1.73	(0.22)	0.42
Using gross rent (1960–2010)	0.87	(0.08)	0.31			
Dropping small locations:						
CZ population above 25th percentile (>25k in 2010)	1.26	(0.06)	0.33	1.99	(0.28)	0.33
CZ population above 50th percentile (>50k in 2010)	1.28	(0.06)	0.34	2.10	(0.30)	0.35
CZ population above 75th percentile (>250k in 2010)	1.39	(0.06)	0.36	2.39	(0.34)	0.40

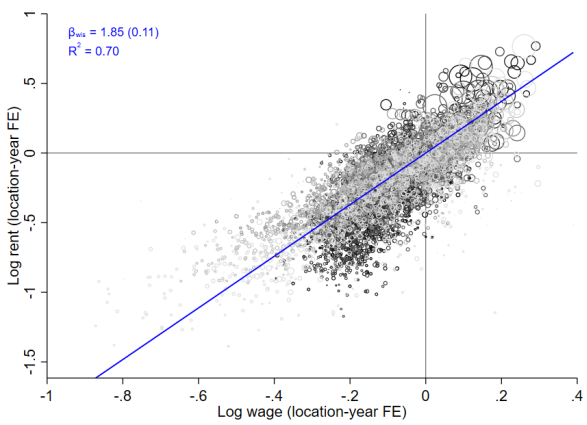
Notes. See Appendix D for detail. Source: Decennial Census and ACS.

Table A.10: Robustness of Descriptive Analysis: Decade-by-decade

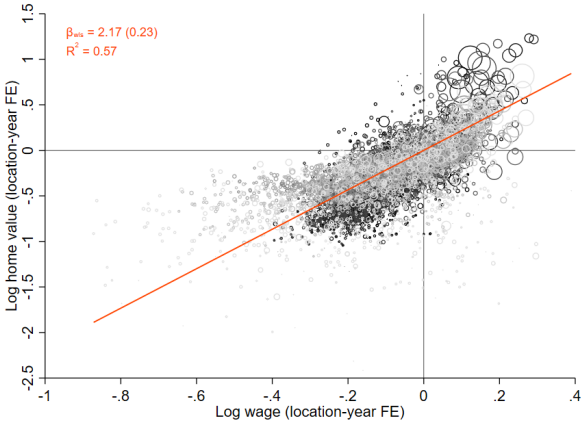
	10-year wage elasticity of:					
	Rents			Home prices		
	β	SE(β)	R^2	β	SE(β)	R^2
Panel A: Using CZ-level data						
1940–1950 and 1950–1960	2.11	(0.27)	0.52	1.73	(0.30)	0.24
1960–1970	0.51	(0.14)	0.10	0.56	(0.19)	0.10
1970–1980	0.03	(0.28)	0.00	0.57	(0.48)	0.05
1980–1990	1.62	(0.17)	0.73	3.54	(0.29)	0.82
1990–2000	1.37	(0.16)	0.36	1.98	(0.44)	0.17
2000–2010	0.80	(0.12)	0.24	1.83	(0.44)	0.30
Panel B: Using county-level data						
1940–1950 and 1950–1960	1.47	(0.16)	0.43	1.42	(0.16)	0.25
1960–1970	0.79	(0.10)	0.24	0.64	(0.13)	0.14
1970–1980	0.24	(0.18)	0.02	0.64	(0.29)	0.08
1980–1990	1.53	(0.15)	0.61	3.22	(0.27)	0.73
1990–2000	1.18	(0.11)	0.31	1.57	(0.23)	0.18
2000–2010	0.81	(0.08)	0.25	1.60	(0.40)	0.25

Notes. See Appendix D for detail. Source: Decennial Census and ACS.

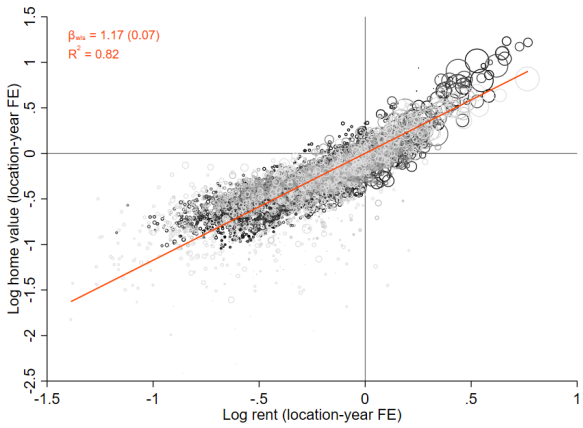
Figure A.1: Relationship Between Initial Wages and Initial Housing Costs



(a) rent vs. wage



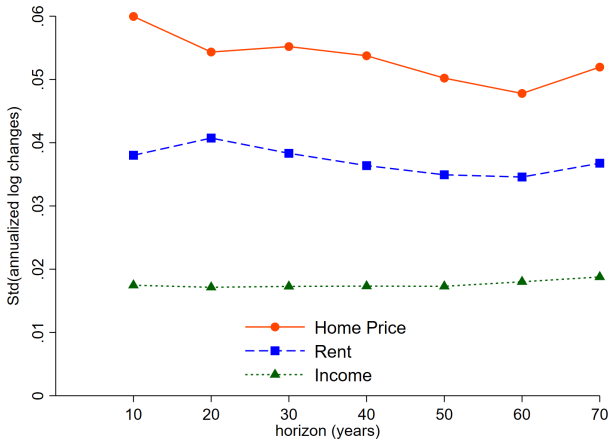
(b) home price vs. wage



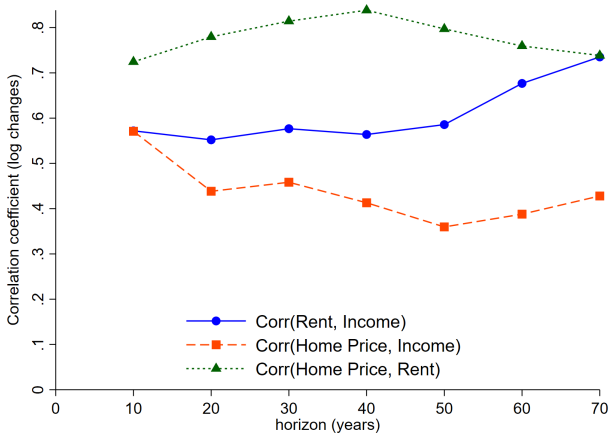
(c) home price vs. rent

Notes. This figure shows the positive association between (log) wages and housing costs in the pooled cross-sections. In the baseline, we control for this heterogeneity by starting each household from the same initial price vector of wages and housing at time t (wage \$58k, rent \$13k, and home price \$186k). When introducing moves in the risk cost analysis, and in extensions where households make choices "behind the veil", this cross-sectional dispersion is also part of the price risk that households face. Each circle is a housing cost wage pair in a census year. Circle areas are proportional to the location's population share. Observations in darker shades are from more recent decades. Linear regressions are weighted using CZ population and standard errors in brackets are clustered by state. Source: Decennial Census and ACS.

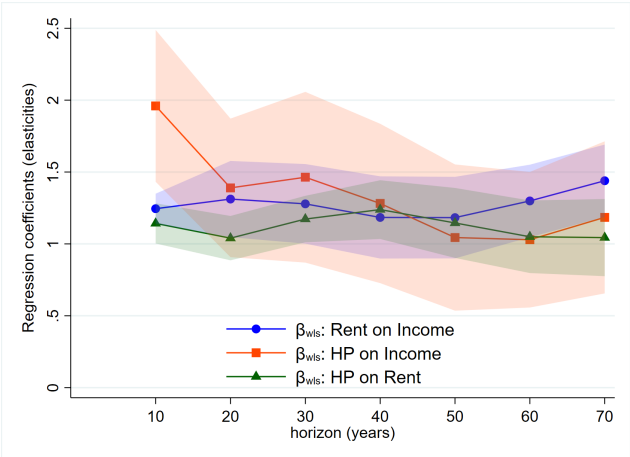
Figure A.2: Volatility and Correlation of Location-Specific Risk by Horizon



(a) standard deviations of annualized log changes



(b) correlation coefficients of log changes

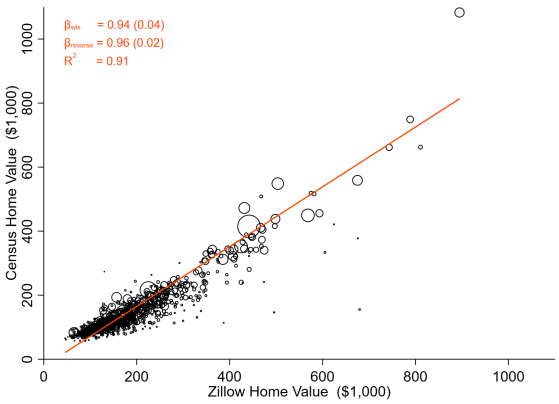


(c) regression coefficients of log changes (elasticities)

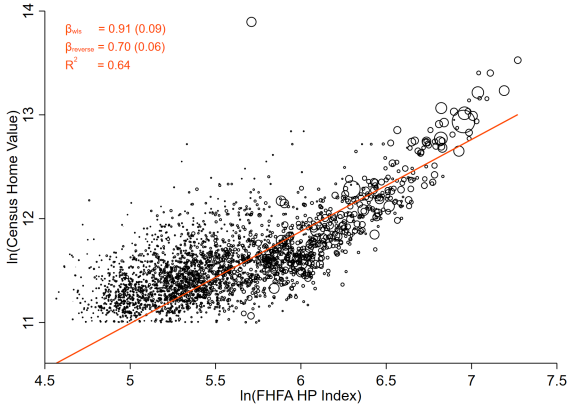
Notes. Location-specific risk of real wages, rents, and home prices. All calculations weigh CZ-year observations using a location's populations share, separately in each decade so that the weights sum to one in each decade (i.e., each decade receives the same total weight in all calculations, independent of the growth of total U.S. population from 1940 to 2010). *Panel A.* Standard deviation of annualized log changes of location-specific wage income and housing prices, $\Delta_h \ln x^{loc}_{c,t+h}/\sqrt{h}$, with quality-adjusted log change $\Delta_h \ln x^{loc}_{c,t+h} = \ln x^{loc}_{c,t+h} - \ln x^{loc}_{c,t}$. *Panel B.* Correlation coefficients $Corr(\Delta_h \ln x^{loc}_{t+h}, \Delta_h \ln y^{loc}_{t+h})$. *Panel C.* Elasticities $\beta_{wls,y|x}(h)$ are obtained from regressing log changes in housing costs on log changes in wage income and log changes in home prices on log changes in rents, weighted by population: $\Delta_h \ln y^{loc}_{t+h} = a + \beta_{wls,y|x}(h) \cdot \Delta_h \ln x^{loc}_{t+h} + u_{t+h}$. Shaded areas are 95% confidence intervals based on robust standard errors clustered by state. Source: Decennial Census and ACS.

Figure A.3: Comparison of Census Home Values with FHFA and Zillow Transaction Prices

A. Census vs. Transaction Prices: (Log) Levels

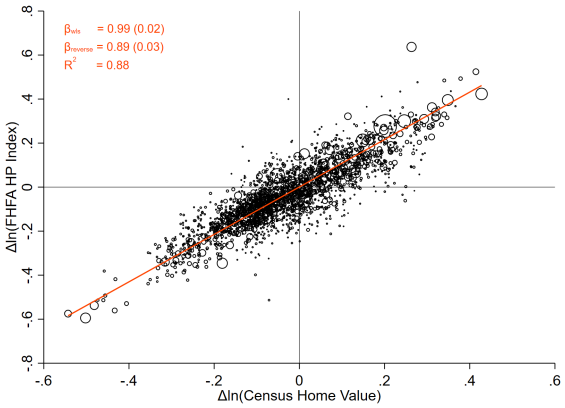


(a) dollars

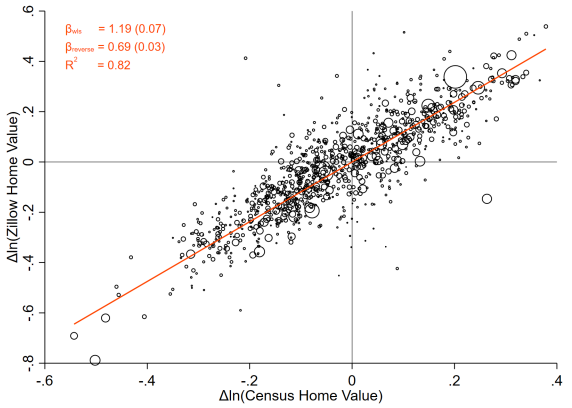


(b) logs

B. Census vs. Transaction Prices: Growth Rates

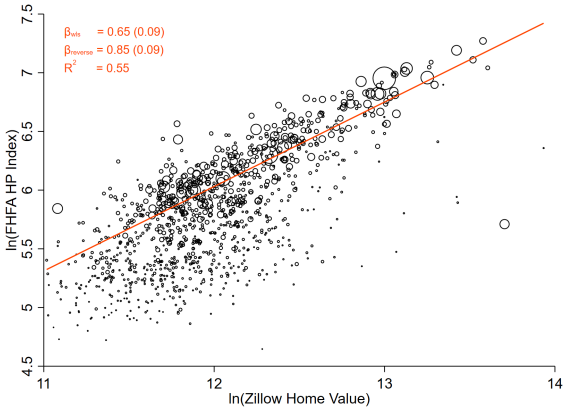


(c) logs

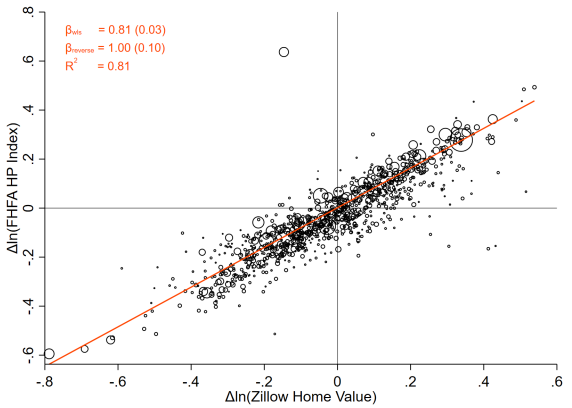


(d) 10-year log changes

C. Comparing the Transaction Prices with Each Other



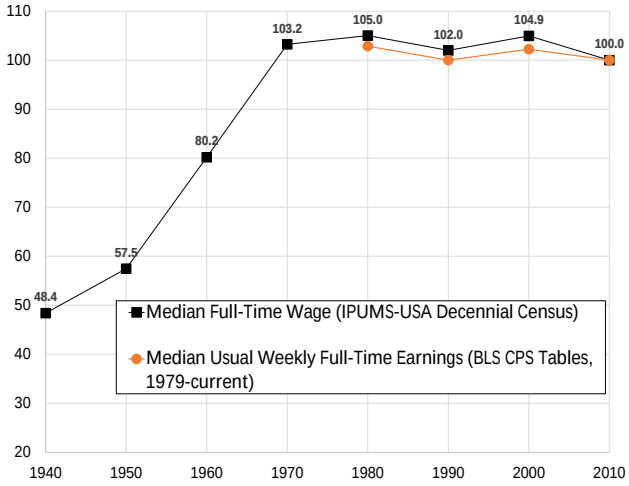
(e) logs



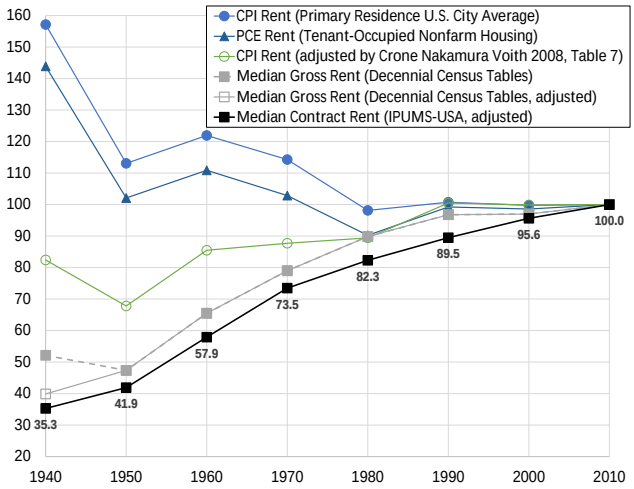
(f) 10-year log changes

Notes. Census county median home value (in 2020 dollars), ZHVI Zillow Home Value Index All Homes Time Series (in 2020 dollars), and FHFA Federal Housing Finance Agency Annual House Price Index (index value). Circle areas represent county population size relative to total population. Source: Decennial Census and ACS (Ruggles et al. 2018), FHFA (U.S. Federal Housing Finance Agency 1975–2016), and Zillow (Zillow 1996–2015).

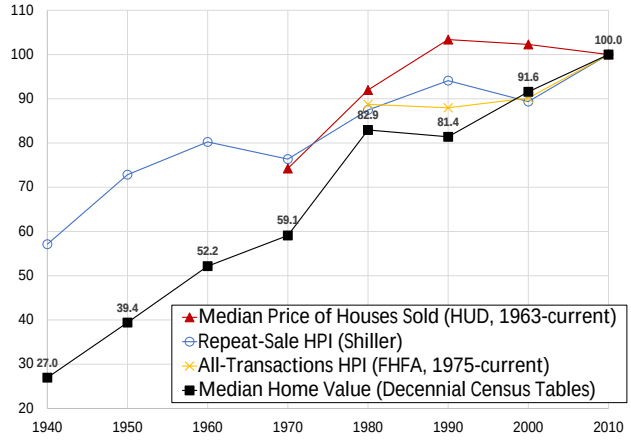
Figure A.4: Decadal Series of National Wage, Rent, and Home Price



(a) wage series



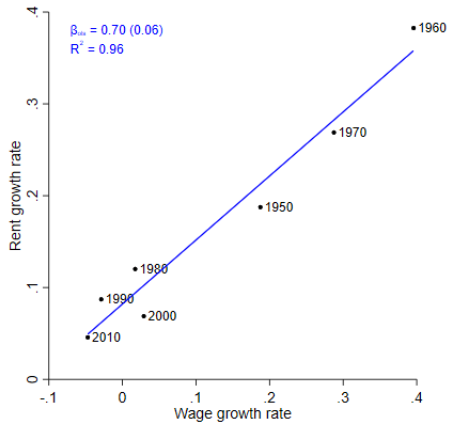
(b) rent series



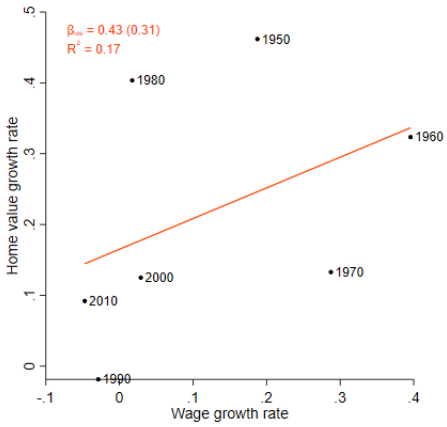
(c) home price series

Notes. All price series are real, deflated using the Consumer Price Index (CPI) for all urban consumer, and normalized to 100 in 2010 for comparability. Source: *Wage series.* "Median Full-Time Wages": IPUMS-USA microdata 1940–2010, Ruggles et al. (2018). "Median Usual Weekly Full-Time Earnings": Current Population Survey (CPS) Tables, U.S. Bureau of Labor Statistics (1979–2021c). *Rent series.* "CPI Rent of Primary Residence in U.S. City Average": Bureau of Labor Statistics (U.S. Bureau of Labor Statistics 1913–2021a), with adjustment by Crone et al. (2010) for non-response bias from 1940 to 1985. "PCE Rent (Tenant-Occupied Nonfarm Housing)": Bureau of Economic Analysis (U.S. Bureau of Economic Analysis 1929–2021b). "Median Gross Rent": Historical Census of Housing Tables 1940–2000 (U.S. Census Bureau 1940–2000) and ACS 2010 Tables (Flanagan and Wilson 2013; American Community Survey 2000–2019), with adjustment for differences in the rent survey question in 1940. "Median Contracts Rent": IPUMS-USA (Ruggles et al. 2018), with adjustment and imputation of 1950 as described in the text. *Home price series.* "Median Sales Price for New Houses Sold in the United States": Department of Housing and Urban Development (U.S. Department of Housing and Urban Development 1963–2021). "Repeat-Sale HPI (Shiller): Shiller (1890–2021). "All-Transaction HPI: U.S. Federal Housing Finance Agency (1975–2021). "Median Home Value": Historical Census of Housing Tables (U.S. Census Bureau 1940–2000) and ACS 2010 Tables (Flanagan and Wilson 2013; American Community Survey 2000–2019).

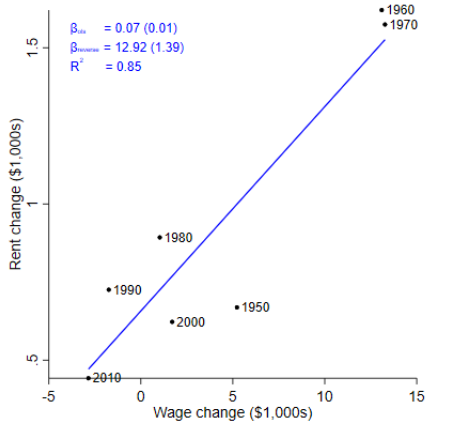
Figure A.5: Correlation of Growth in Wage, Rent, and Home Price: US Decennial Census Data



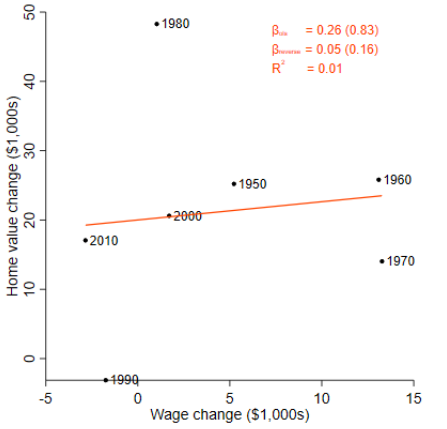
(a) rent vs. wage



(b) home price vs. wage



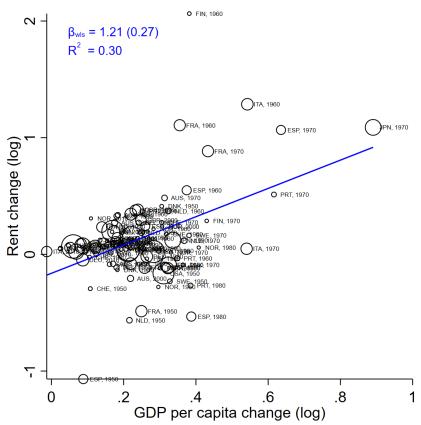
(c) rent vs. wage



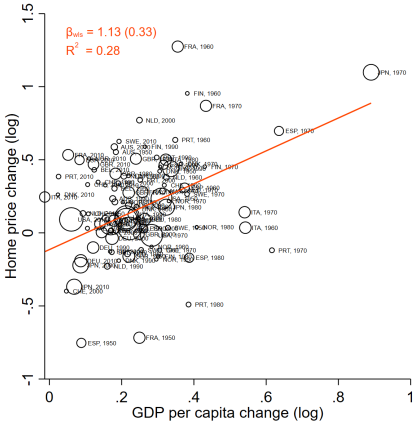
(d) home price vs. wage

Notes. 10-year national log changes in wages, rents, and home prices, $\Delta_10 \ln x^{agg}_{t+10} = \ln x^{agg}_{t+10} - \ln x^{agg}_{t}$, in (a) and (b), respectively real dollar changes, $\Delta_10 x^{agg}_{t+10} = x^{agg}_{t+10} - x^{agg}_{t}$, in (c) and (d). β_{wls} and R^2 are obtained from regressing housing costs on wages (growth rates and real dollar changes). $\beta_{reverse}$ is obtained from the corresponding reverse regression of wages on housing costs. All nominal prices are deflated using the Consumer Price Index (CPI) for All Urban Consumers (U.S. Bureau of Labor Statistics, 1913–2021a). Source: Median full-time wages and contract rents are computed from the IPUMS-USA microdata (Ruggles et al. 2018), with rent adjustment and 1950 imputation as described in the text. Median home values are taken from Historical Census of Housing Tables (U.S. Census Bureau 1940–2000) and ACS 2010 Tables (Flanagan and Wilson 2013; American Community Survey 2000–2019).

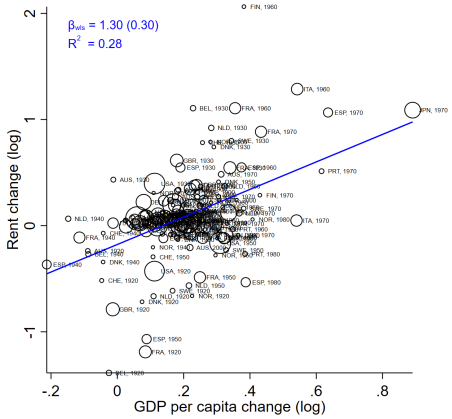
Figure A.6: Correlation of Cross-Country Growth in Income, Rent, and Home Price



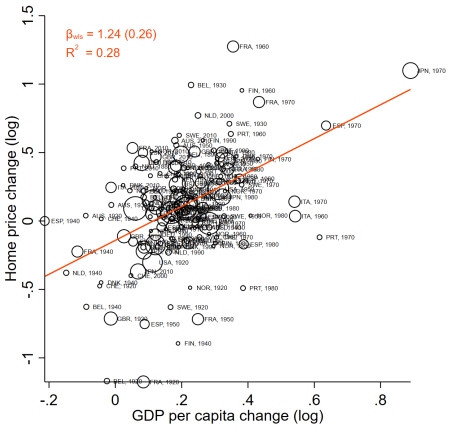
(a) rent vs. income: 1940–2010



(b) price vs. income: 1940–2010



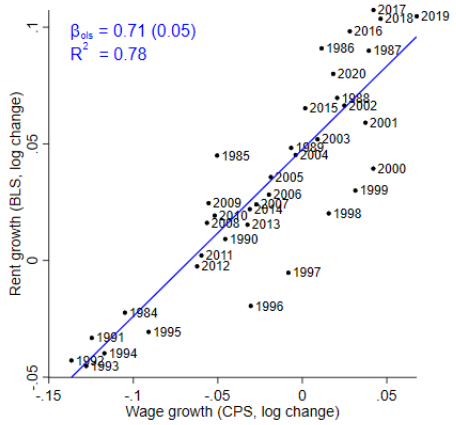
(c) rent vs. income: 1870–2010



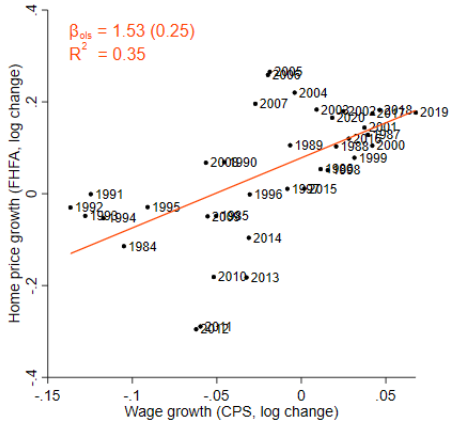
(d) price vs. income: 1870–2010

Notes. 10-year log changes $\Delta_{10} \ln x_{t+10}^{agg}$ of real (purchasing power parity adjusted) income (GDP per capita), rents, and home prices by country. Log changes are winsorized at the 1st and 99th percentile of the unweighted distribution. The data set includes the following 16 countries: Australia, Belgium, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, the UK, and the US. (a) and (b) match the sample period of the rest of the analysis, 1940–2010, and (c) and (d) use the full sample of non-overlapping 10-year growth rates, 1870–2010. Circle areas represent country population size relative to total population in each year. β_{wls} and R^2 are obtained from regressing log changes of housing costs on log changes of GDP per capita, weighted by population. Robust standard errors in brackets are two-way clustered by country and year. Weights add up to the number of countries in each year of the unbalanced panel so that each year has the same weight in the regressions. Unweighted regressions yield similar results. Source: Jordà et al. (2019).

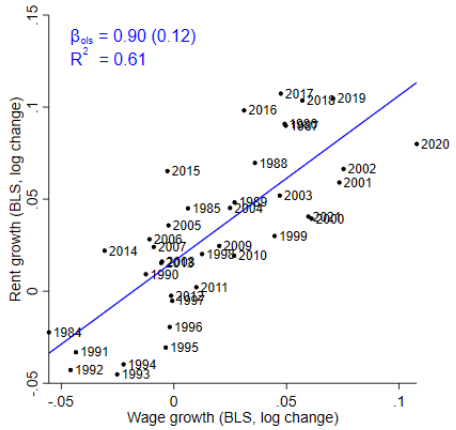
Figure A.7: Correlation of National Growth in Wage, Rent, and Home Price: Annual Data



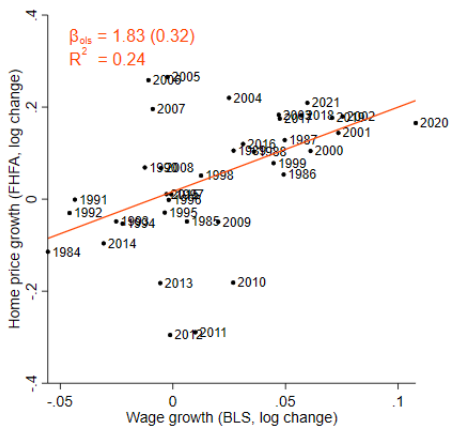
(a) rent (BLS) vs. wage (IPUMS-CPS)



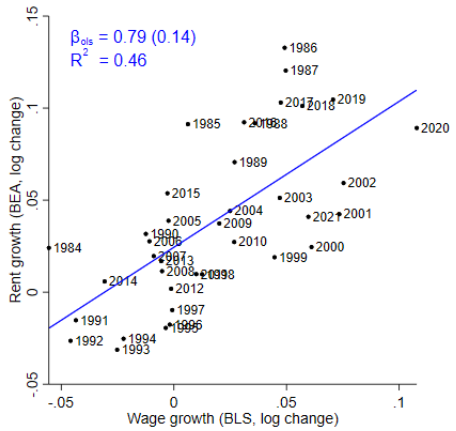
(b) price (FHFA) vs. wage (IPUMS-CPS)



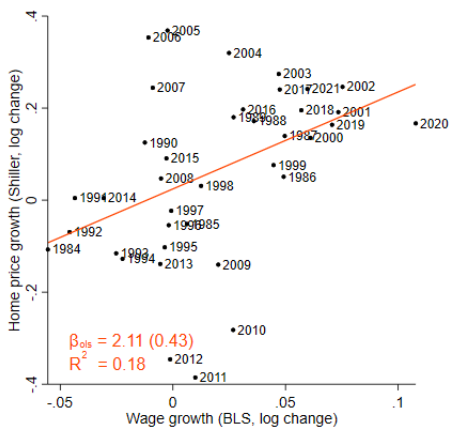
(c) rent (BLS) vs. wage (BLS-CPS)



(d) price (FHFA) vs. wage (BLS-CPS)



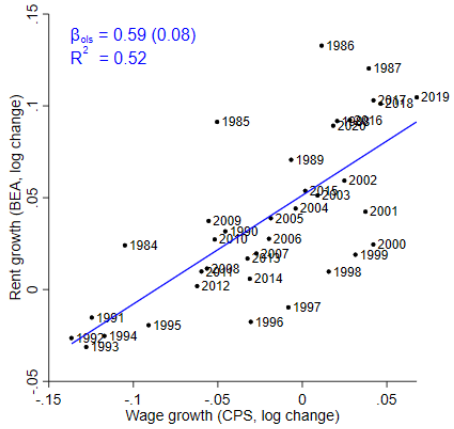
(e) rent (BEA) vs. wage (BLS-CPS)



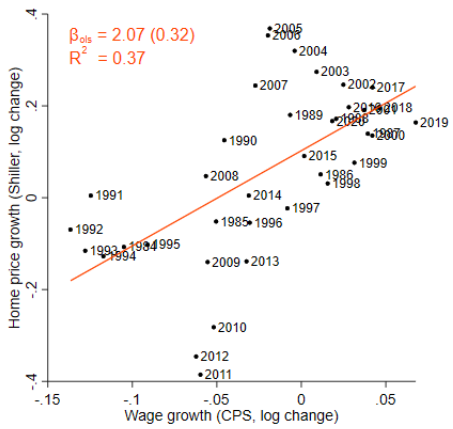
(f) price (Shiller) vs. wage (BLS-CPS)

(continued)

Figure A.7: Correlation of National Growth in Wage, Rent, and Home Price: Annual Data *(cont.)*



(g) rent (BEA) vs. wage (IPUMS-CPS)



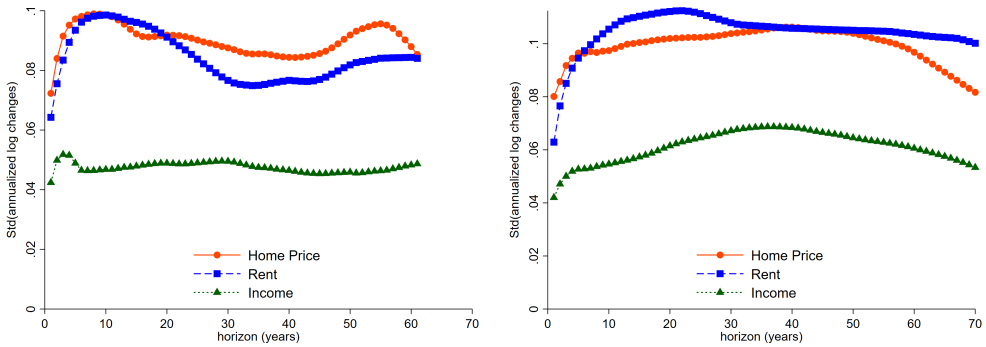
(h) price (Shiller) vs. wage (IPUMS-CPS)

Notes. 5-year national log changes, 1979–2021, with overlapping periods, $\Delta_5 \ln x_{t+5}^{agg} = \ln x_{t+5}^{agg} - \ln x_t^{agg}$. β_{ols} and R^2 are obtained from regressing log changes in housing costs on log changes in wages, with robust standard errors shown in brackets. All nominal prices are deflated using the CPI for All Urban Consumers (U.S. Bureau of Labor Statistics 1913–2021a). Using the PCE Price Index (U.S. Bureau of Economic Analysis 1929–2021b) or the Gross Domestic Product (GDP) Implicit Price Deflator (U.S. Bureau of Economic Analysis 1947–2021a) produces similar results.

Source: *Wages.* (a), (b), (g), and (h) measure wages using the IPUMS ASEC microdata (Annual Social and Economic supplement to the Current Population Survey (CPS); Flood et al. 2021) to define full-time wages in the same way as described in Section 3 and Appendix A for the IPUMS-USA Decennial Census microdata (Ruggles et al. 2018). The IPUMS-CPS ASEC microdata is available from 1962–2021, with work and income questions referencing the previous calendar year, hence CPS full-time wage ends in 2020. (c), (d), (e), and (f) instead measure wages using the series "median usual weekly real earnings of wage and salary workers 16 years and over employed full time" provided by the BLS (U.S. Bureau of Labor Statistics 1979–2021c). This series is based on the CPS "merged outgoing rotation group" (or "earner study" or "annual earnings files") of the basic monthly samples, which began in 1979. Work and income questions reference current pay, hence this series ends in 2021. *Rents.* (a) and (c) measure rents using the Consumer Price Index (CPI) "rent of primary residence in U.S. city average" series (U.S. Bureau of Labor Statistics 1914–2021b) provided by the BLS. (e) and (g) measure rents using the Personal Consumption Expenditures (PCE) series "rental chain-type price index of tenant-occupied nonfarm housing" provided by the BEA (U.S. Bureau of Economic Analysis 1929–2021c). *Home prices.* (b) and (d) measure home prices using the "all-transactions house price index for the United States" series provided by the FHFA (U.S. Federal Housing Finance Agency 1975–2021). (f) and (h) measure home prices using real U.S. home price data for Figure 3.1 in Robert J. Shiller's Irrational Exuberance (Shiller 1890–2021).

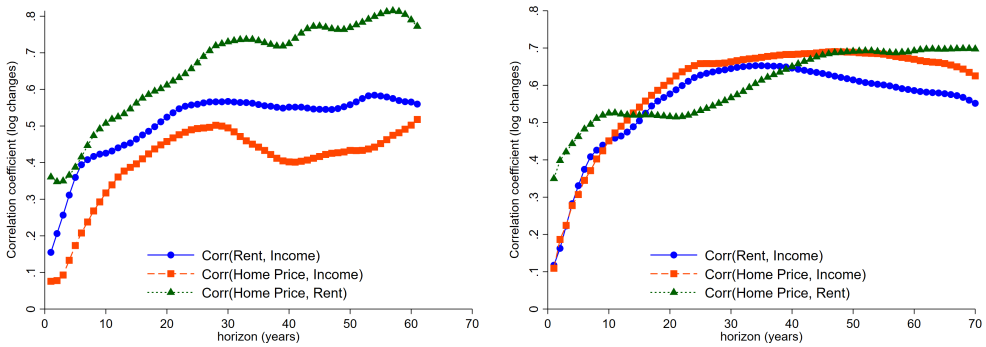
Figure A.8: Volatility and Correlation of Aggregate Risk by Horizon: Cross-country Data

A. Standard deviations of annualized log changes



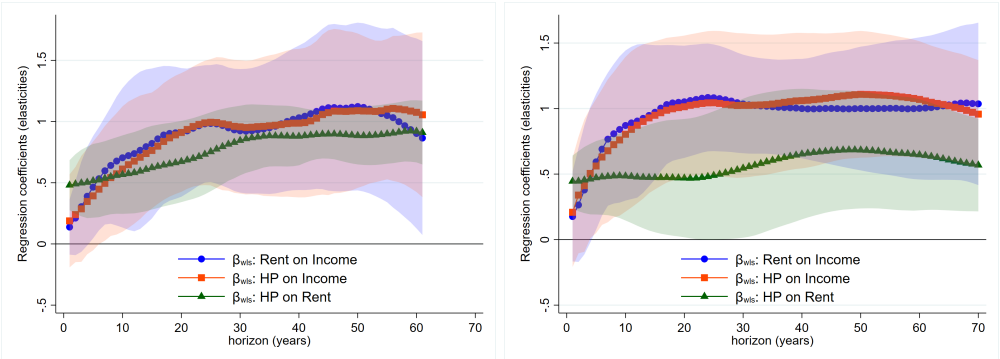
(a) 1940–2010 (b) 1870–2015

B. Correlation coefficients of log changes



(c) 1940–2010 (d) 1870–2015

C. Regression coefficients of log changes (elasticities)

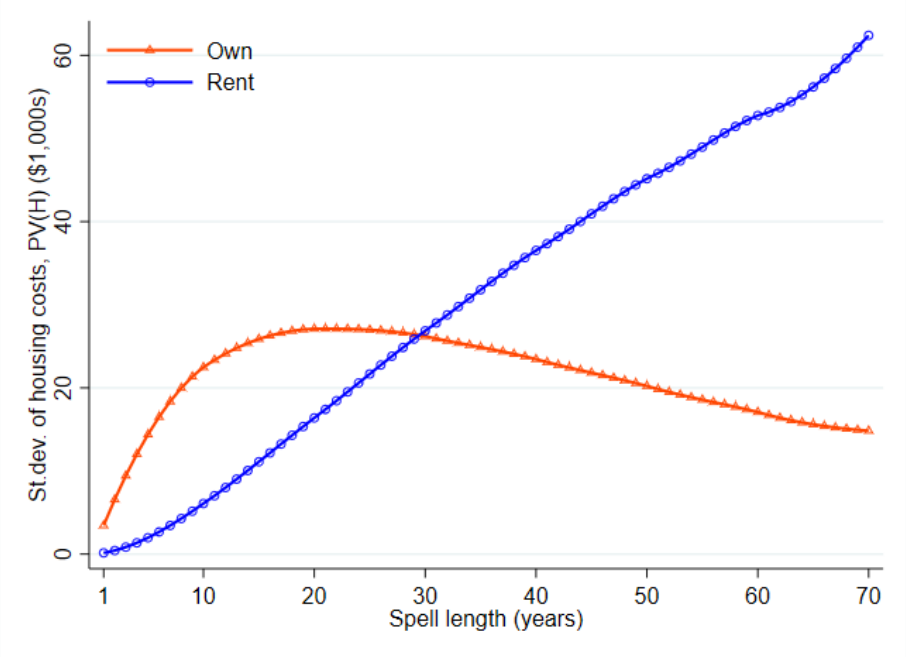


(e) 1940–2010 (f) 1870–2015

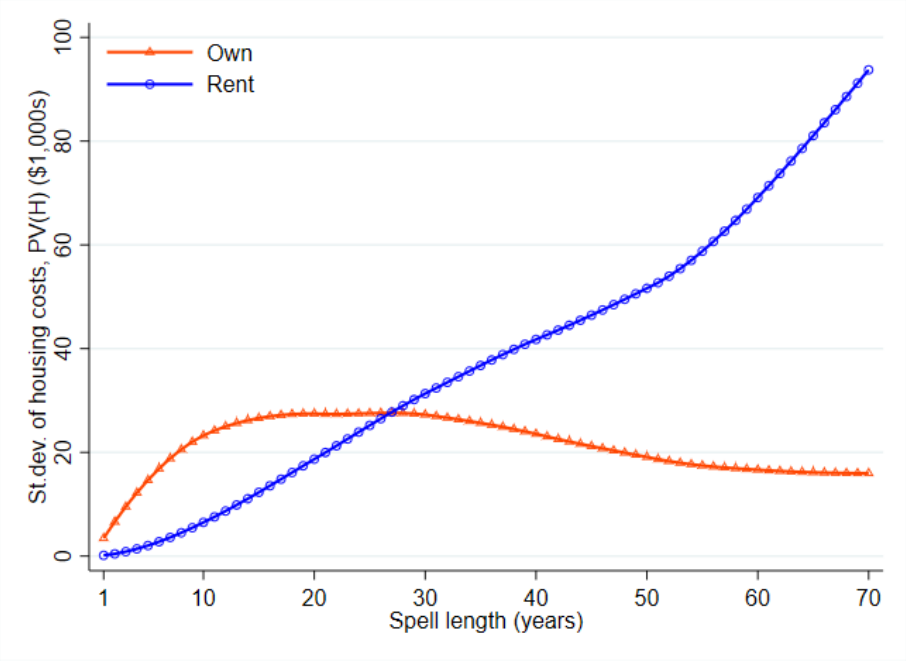
Notes. Aggregate risk of real (purchasing power parity adjusted) incomes (GDP per capita), rents, and home prices by country. Log changes are winsorized at the 1st and 99th percentile of the unweighted distribution. The data set includes the following 16 countries: Australia, Belgium, Denmark, Finland, France, Germany, Italy, Japan, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, the UK, and the US. Figures in the left column match the sample period of the rest of the analysis, 1940–2010; figures in the right column use the entire sample from 1870–2015 and impute missing 1-year growth rates in interim periods. This is the data used for the non-parametric model of aggregate risk described in Section B.1. Statistics in the left column use weights corresponding to country population size relative to total population in each year that add up to the number of countries in each year of the unbalanced panel so that each year has the same weight in the calculations. Weights for statistics shown in the right column are population shares that are constant across all years and are computed over the period where all 16 countries are present (mostly years 1949–2015). All figures show results based on at least 100 observations, corresponding to a horizon up to 61 years for the left column and 70 years for the right column, respectively. *Panel A.* Standard deviation of annualized log changes of aggregate income and housing prices, $\Delta_h \ln x^{agg}_{t+h} / \sqrt{h}$, with h-period log change $\Delta_h \ln x^{agg}_{t+h} = \ln x^{agg}_{t+h} - \ln x^{agg}_{t}$. *Panel B.* Correlation coefficients $Corr(\Delta_h \ln x^{agg}_{t+h}, \Delta_h \ln y^{agg}_{t+h})$. *Panel C.* Elasticities $\beta_{wls,y|x}(h)$ are obtained from regressing log changes in housing costs on log changes in GDP per capita and log changes in home prices on log changes in rents, weighted by population: $\Delta_h \ln y^{agg}_{t+h} = a + \beta_{wls,y|x}(h) \cdot \Delta_h \ln x^{agg}_{t+h} + u_{t+h}$. Shaded areas are 95% confidence intervals based on two-way clustered robust standard errors by country and year. Unweighted regressions yield similar results. Source: Jordà et al. (2019).

Figure A.9: Figure 3 with Simulated and Empirical Data

A. Simulated data



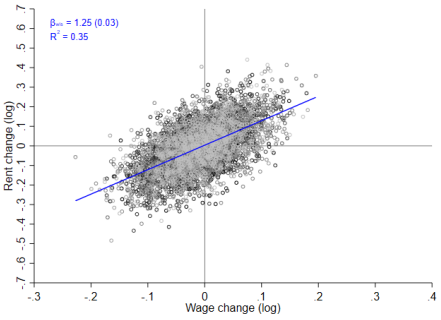
B. Empirical data



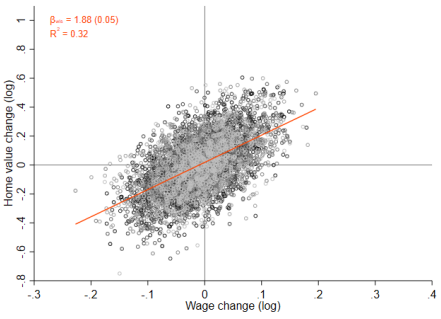
Notes. See notes of Figure 3. Source: Decennial Census and ACS.

Figure A.10: Figure 4 with Simulated and Empirical Data

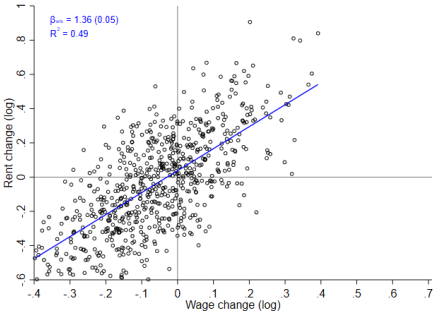
A. Simulated data:



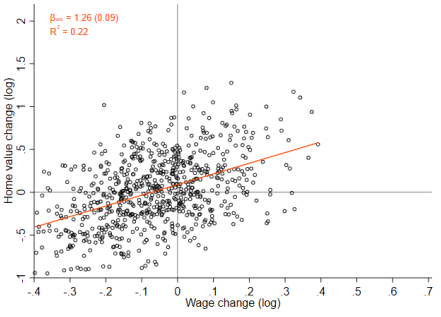
(a) 10-year R vs. y growth



(c) 10-year P vs. y growth

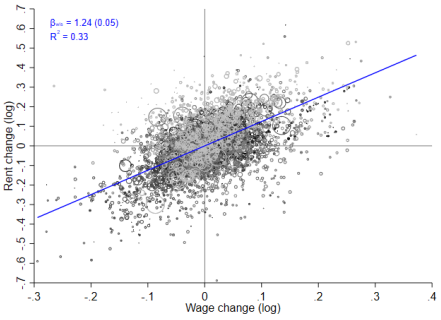


(e) 70-year R vs. y growth

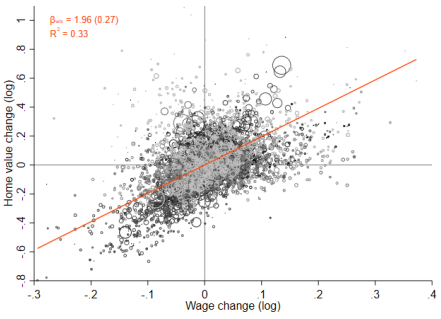


(g) 70-year P vs. y growth

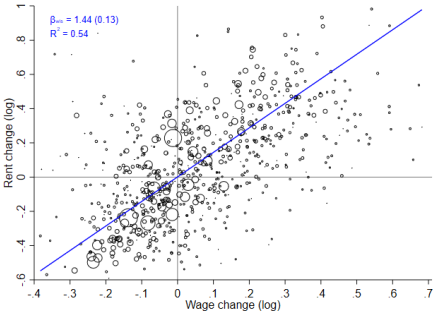
B. Empirical data:



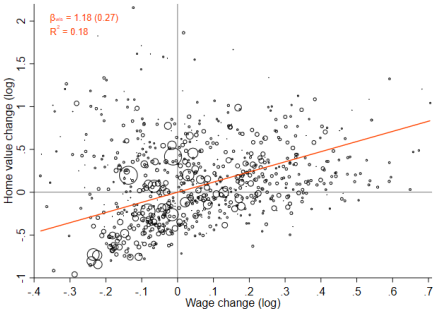
(b) 10-year R vs. y growth



(d) 10-year P vs. y growth



(f) 70-year R vs. y growth

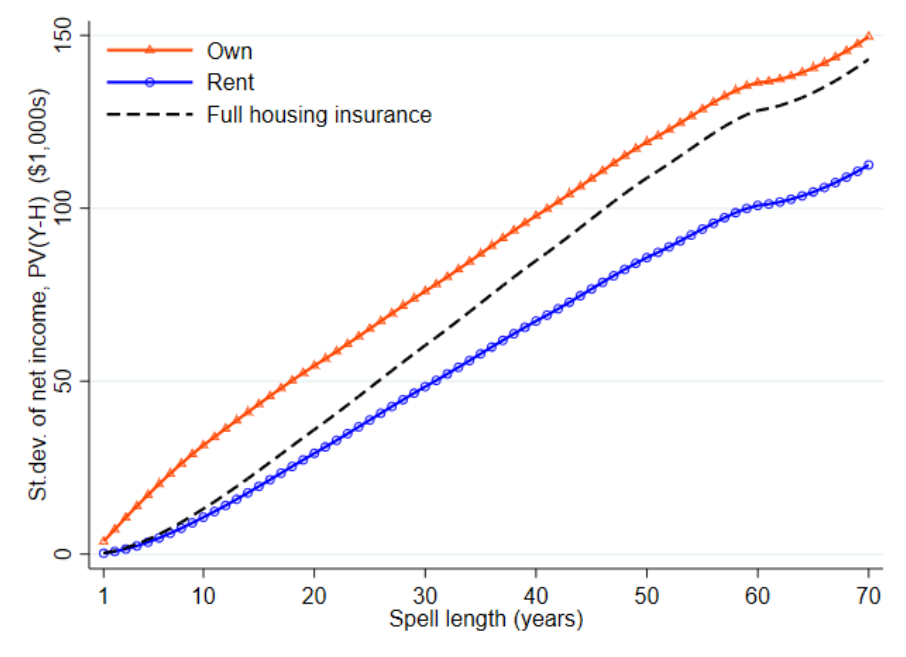


(h) 70-year P vs. y growth

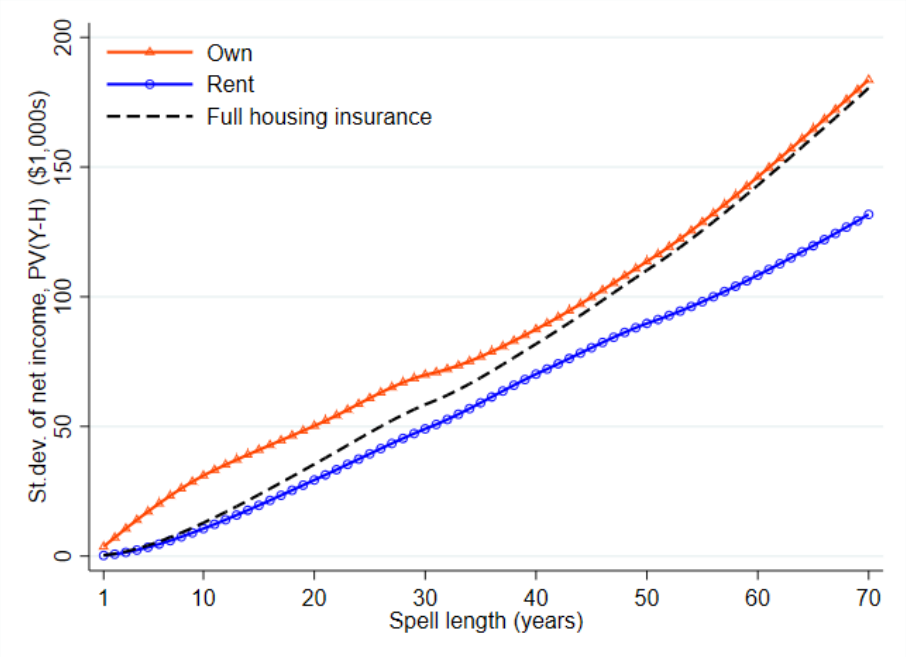
Notes. See notes of Figure 4. Source: Decennial Census and ACS.

Figure A.11: Figure 5 with Simulated and Empirical Data

A. Simulated data



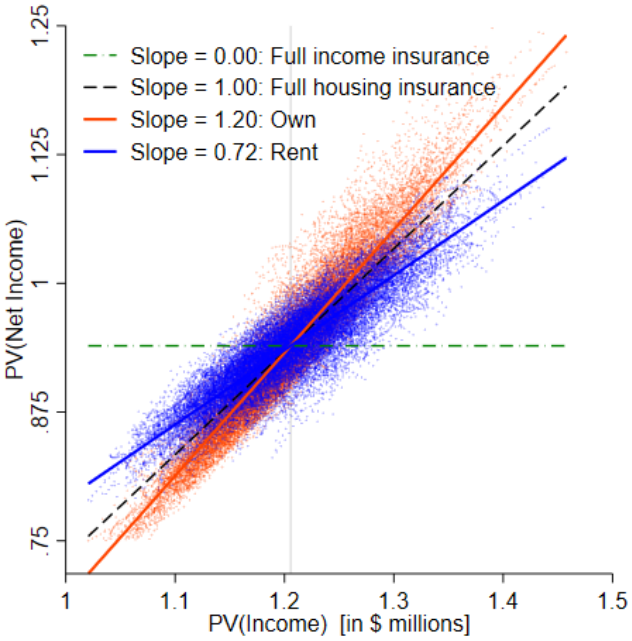
B. Empirical data



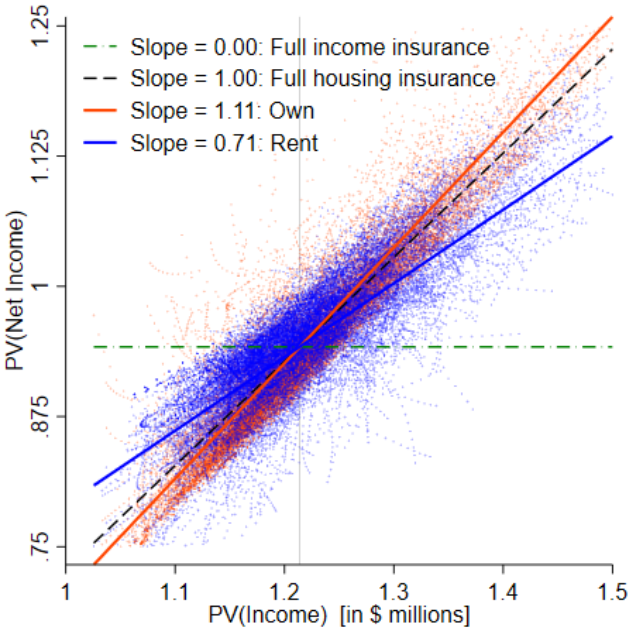
Notes. See notes of Figure 5. Source: Decennial Census and ACS.

Figure A.12: Figure 6 with Simulated and Empirical Data

A. Simulated data



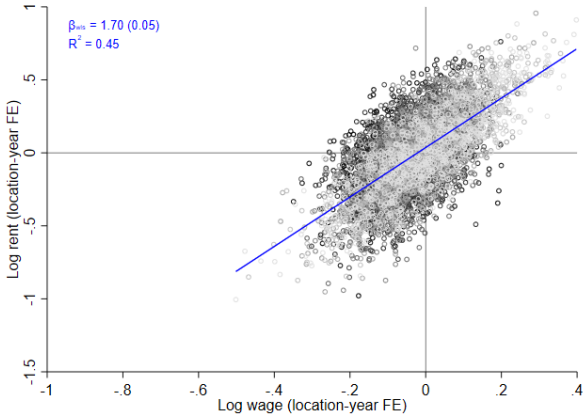
B. Empirical data



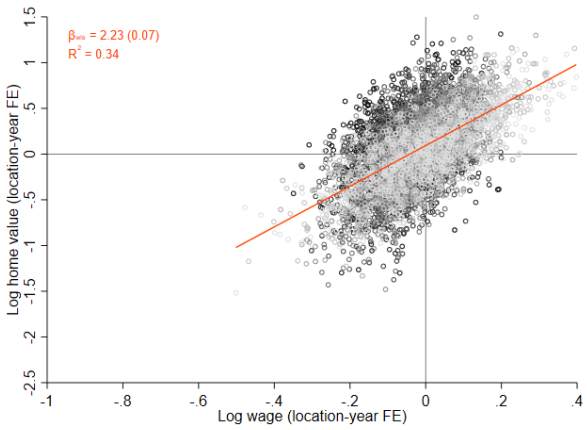
Notes. See notes of Figure 6. Source: Decennial Census and ACS.

Figure A.13: Figure A.1 with Simulated and Empirical Data

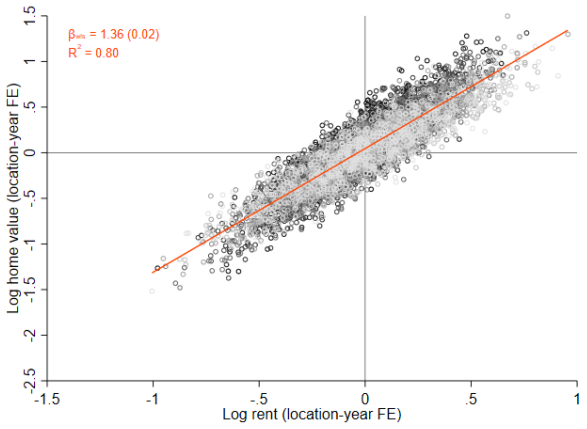
A. Simulated data:



(a) rent vs. wage

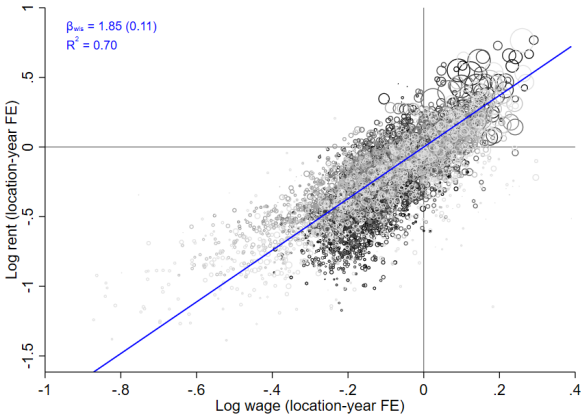


(c) home price vs. wage

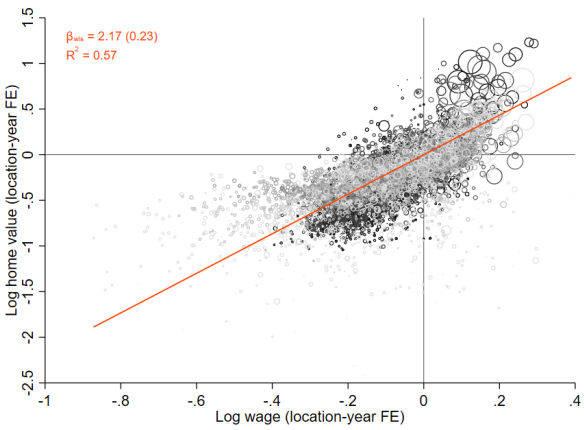


(e) home price vs. rent

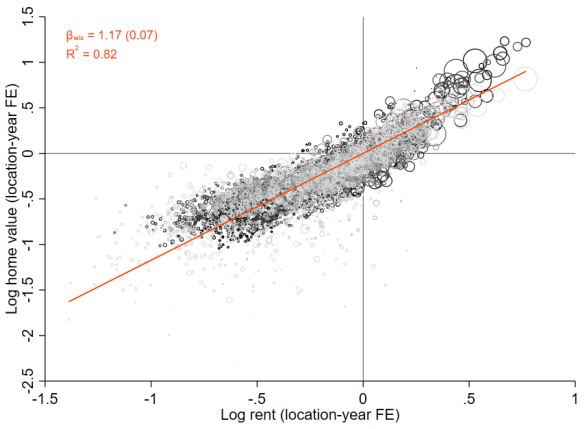
B. Empirical data:



(b) rent vs. wage



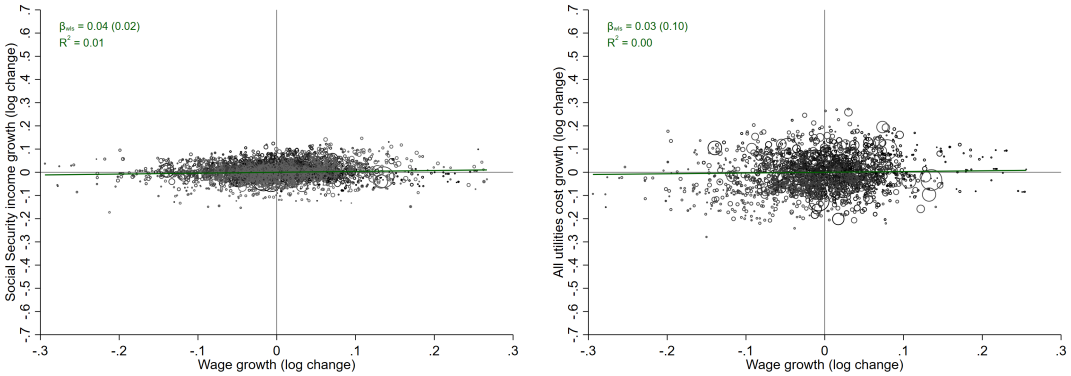
(d) home price vs. wage



(f) home price vs. rent

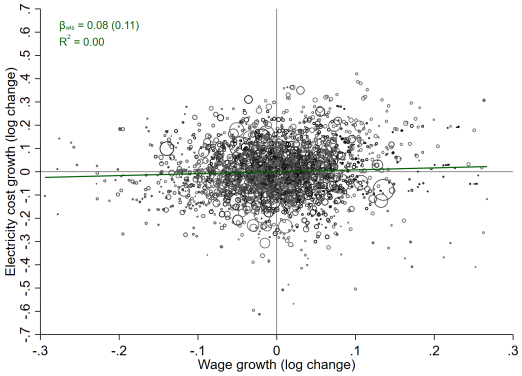
Notes. See notes of Figure A.1. Source: Decennial Census and ACS.

Figure A.14: Correlation of Changes in Social Security Income and Utility Costs with Wages

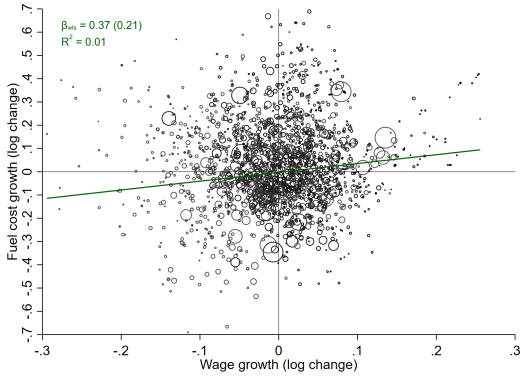


(a) Social Security income

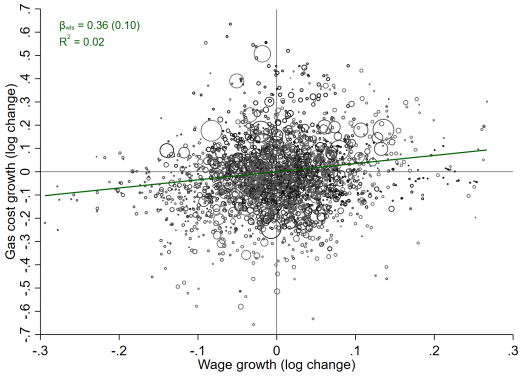
(b) all utilities



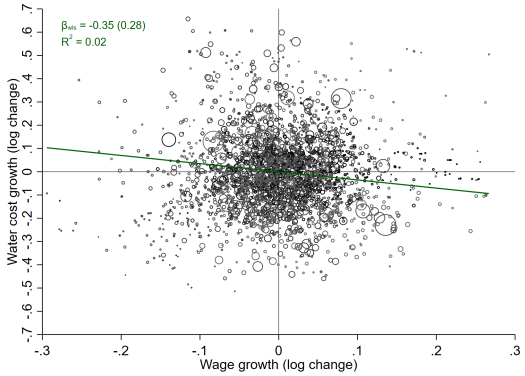
(c) electricity cost



(d) fuel cost



(e) gas cost

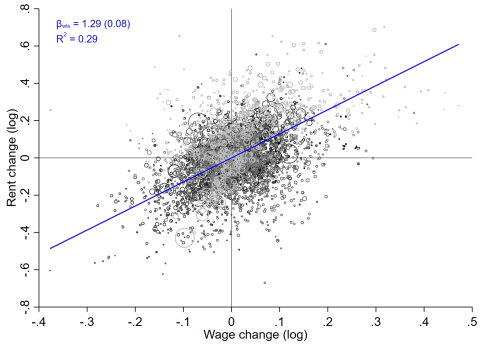


(f) water cost

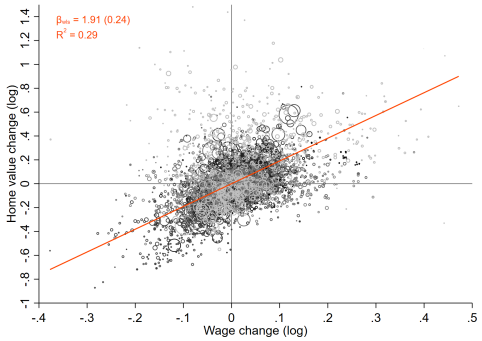
Notes. 10-year CZ-level log changes in wages, rents, and home prices relative to national-level log changes; all quality-adjusted as described in the text. For comparison, the vertical axis has the same range as that for rent changes. Social Security is a federal program without local cost of living adjustments. Hence, it should not be strongly related to local wage changes. Similarly, electricity is supplied competitively on a national grid and prices are mostly determined at the national level. (a) uses Social Security income; (b) to (e) use utility expenses and its components, which are available from 1970 on, except for home heating fuel cost (d), which is consistently available since 1980; see https://usa.ipums.org/usa-action/variables/COSTFUEL#comparability_section). Circle areas are proportional to the location's population share. Observations in darker shades are from more recent decades. β_{wls} and R^2 are obtained from regressing log changes in housing costs on log changes in wages, weighted by population. Robust standard errors in brackets are clustered by state. Prices are adjusted to dollars of 2020 using the Consumer Price Index (CPI) for All Urban Consumers. Source: Decennial Census and ACS.

Figure A.15: Wage, Rent, and Home Price Relationships Across Distribution Within Location

A. 25th percentile

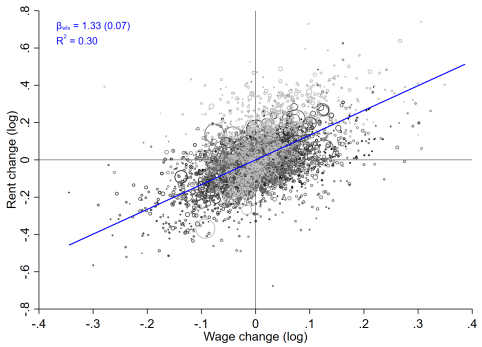


(a) rent vs. wage

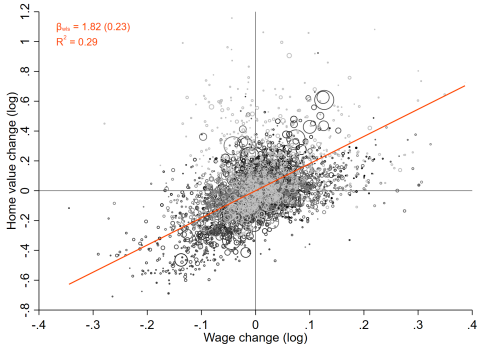


(b) home price vs. wage

B. 50th percentile

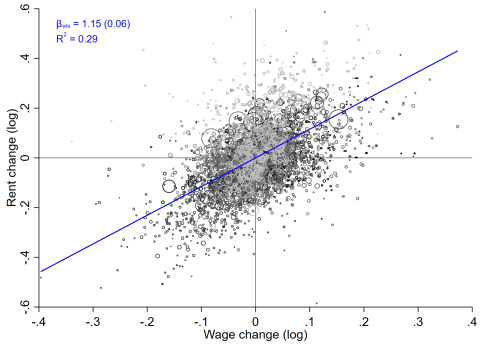


(c) rent vs. wage

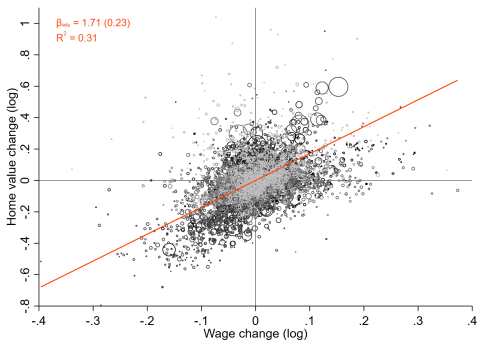


(d) home price vs. wage

B. 75th percentile



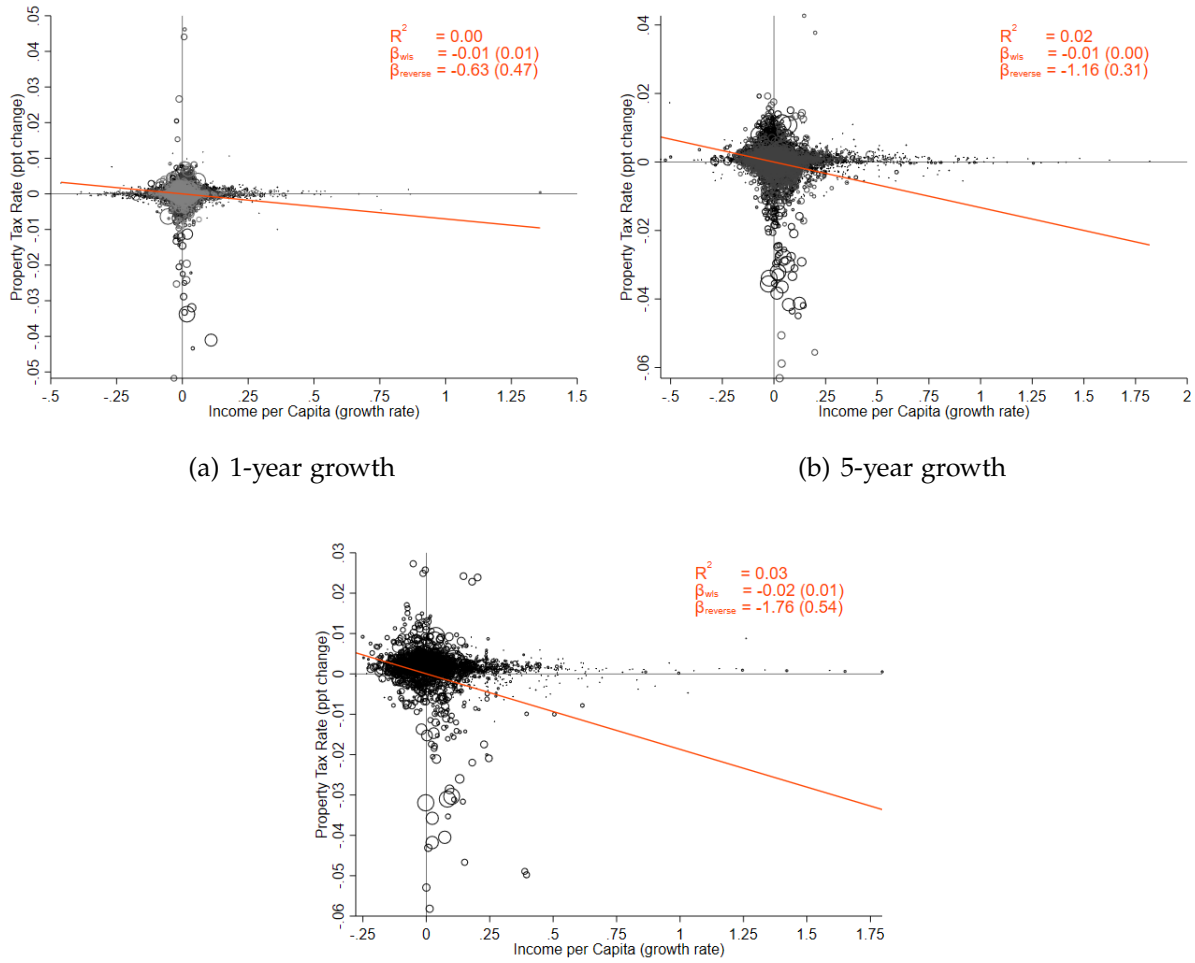
(e) rent vs. wage



(f) home price vs. wage

Notes. 10-year CZ-level changes in wages, rents, and home prices relative to national-level changes; all quality-adjusted as described in the text. Specifically, each figure shows the demeaned year by location-level percentile $Q_p(\hat{u}_{it} \mid c(i), t)$ of the residual $\hat{u}_{it}$ from the hedonic regression of prices x_{it}^{ipums} observed in the IPUMS microdata on measures of quality and quantity without location-year fixed effects, i.e., $x_{it}^{ipums} = \alpha_t + q_{it}'\delta_t + u_{it}$. (a) and (b) use the 25th percentile, (c) and (d) the 50th percentile, and (e) and (f) the 75th percentile of the residual of wages, rents, and home values by CZ c. The sample includes years 1940–2010. Circle areas are proportional to the location's population share. Observations in darker shades are from more recent decades. β_{wls} and R^2 are obtained from regressing log changes in housing costs on log changes in wages, weighted by population. Robust standard errors in brackets are clustered by state. Prices are adjusted to dollars of 2020 using the Consumer Price Index (CPI) for All Urban Consumers. Source: Decennial Census and ACS.

Figure A.16: Correlation of Changes in Property Tax Rates with Income Growth



(a) 1-year growth

(b) 5-year growth

(c) 10-year growth

Notes. County-level growth rates relative to national-level growth rates of personal income per capita and changes in property tax rates by county, 2003–2015. (a) uses 1-year growth rates; (b) uses 5-year growth rates; and (c) uses 10-year growth rates. Circle areas are proportional to the county's population share. Observations in darker shades are from more recent years. Source: Baker et al. (2020) and Bureau of Economic Analysis (BEA).

Figure A.17: Illustration of the Marginal Excess Burden of Owner-Occupation Net Subsidy