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On non-robustness of income polarization measures to housing cycles

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Abstract

A cross-country comparison of the middle class as measured by income polarization indices is commonplace in welfare economics. Using the 2001–2007 housing cycle and data for Australia, the United States, Germany, and Switzerland (and an array of methods, including triple-difference design), I show that polarization indices based on disposable income are unreliable. The cycle changes the relative importance of non-monetary income from housing (imputed rent), particularly for middle-income households. Therefore, to ensure that convenient income measures do not misrepresent the size of the middle class, researchers should verify the absence of swings in housing prices during their study period.

Keywords: homeownership, imputed rent, income distribution, polarization, middle class
JEL codes: D31, I3, C81.

** The work benefited from the comments made during seminars held at TTPI, ANU and ARC CEPAR, UNSW. Furthermore, I extend my heartfelt gratitude to Matthew Smith, Nathan Deutscher, Robert Breunig, and the numerous anonymous (though mostly critical) reviewers whose comments have greatly contributed to the improvement of this work. First draft: 1st August 2021. This draft: 9th August 2023. Seek the latest version on ResearchGate. Author contact:* sergey.alexeev@sydney.edu.au

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Working paper

On non-robustness of income polarisation measures to housing cycles

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Abstract: A cross-country comparison of the middle class as measured by income polarisation indexes is commonplace in welfare economics. Using the 2001–2007 housing cycle and data for Australia, the United States, Germany, and Switzerland (and an array of methods, including triple-difference design), I show that the Esteban-Ray index based on disposable income is unreliable. The cycle changes the relative importance of non-monetary income from housing (imputed rent; IR), particularly for middle-income families. Therefore, to ensure that convenient income measures do not misrepresent the size of the middle class, researchers should verify the absence of swings in housing prices during their study period.

KEYWORDS. Homeownership, Imputed rent, Income distribution, polarisation, Middle class.

JEL CLASSIFICATION. D31, I3, C81.

1. INTRODUCTION

The empirical literature concerning the distribution of economic outcomes from a theoretical perspective should utilise economic income (Smeeding and Weinberg 2001). In practice, disposable income is almost universally used instead. A long-acknowledged problem of this practice is that it may misinterpret the standards of living because economic well-being is also determined by sources of income that are in-kind, such as public health or education, and non-monetary, such as the consumption of one's own produce or imputed rent (IR) for occupied accommodation (Atkinson and Bourguignon 2000, 2015).

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For better intuition, note that in a village where farmers grow their own food and exchange it, reserving money for occasional city visits, quantifying economic well-being and ignoring non-monetary sources would be impossible. Therefore, income measures that include non-monetary components are conceptually better than those that do not (Brandolini and Smeeding 2016; Canberra Group 2011; Kuznets 1963; Smeeding and Weinberg 2001).

This long-running debate on the concept of income has been renewed with the 'Stiglitz-Sen-Fitoussi' report (Stiglitz 2017; Stiglitz, Sen and Fitoussi 2009). The report stresses that significant changes in the function of households and society have gone undetected because of selectively observed income components. In particular, shifts from non-market to market provision of goods or services may give false impressions of a change in living standards or inequality without an actual increase in economic development. This effect becomes particularly perplexing when examining the extent of the gender gap, as historical gender roles have led females to supply non-market services more frequently than men.

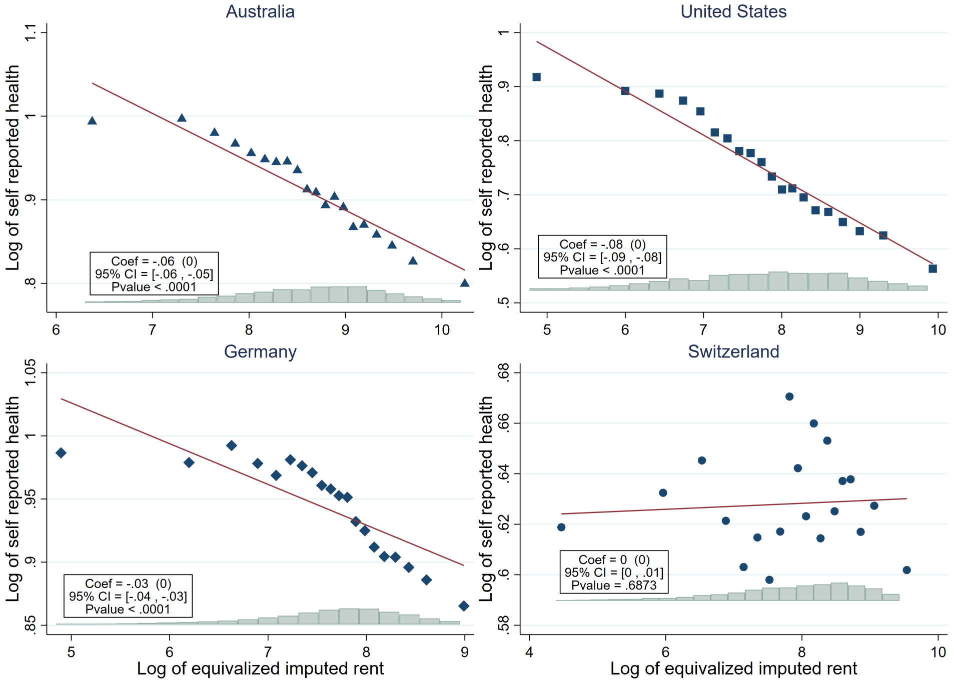
Homeownership or subsidised tenancy is one component of non-monetary income for developed countries with particular quantitative significance. One's home is a financial asset and a source of numerous psychological and consumption benefits. The importance of these benefits became particularly clear during the COVID-19 lockdowns (Horne et al. 2020; Ong 2020).

Since many who do not own their accommodation have to pay rent to access these benefits and the fraction of homeowners differ across countries and within countries across time, ignoring these differences can produce time-series and cross-sectional comparability problems when quantifying economic well-being. A monetary value should be assigned to these benefits and added to the income of rent-free tenants to amend this. This hypothetical stream of benefits is known as IR. 'Imputed' because it is not directly measured but is inferred based on its assumed relationship to variables that can be directly measured.

Consider, for example, two individuals with similar monetary incomes. One lives in a recently well-built million-dollar house, while the other resides in a granny flat that requires daily maintenance. Conventional income measures might suggest that both individuals are equally well-off. In contrast, considering IR as a component of income would appropriately acknowledge the distinction in economic well-being between them.

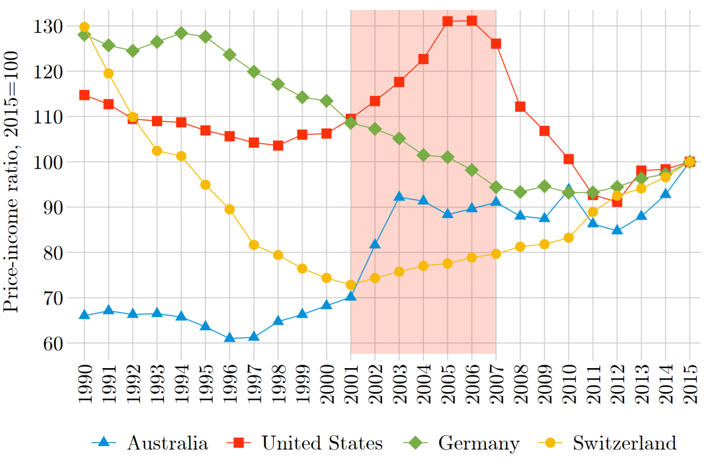
Utilising the data fully presented in Section 2 to stress the potential importance of IR, it can be demonstrated that in some countries, IR could significantly contribute to gen-

FIGURE 1. IR contributes to well-being



Notes: The figure shows a binned scatterplot with a linear fixed effect (household and year) regression line of self-reported health on IR (both variables are logged) controlling for age fixed effects. The sample is limited to household heads and their partners. Errors are clustered on the household level. Self-reported health is rated on a scale of 1 to 5, with 1 being 'excellent' and 5 'poor.' For the data on IR, refer to Section 2. To facilitate the interpretation of the horizontal distance between points, the IR histogram is also displayed in the figure (Pinna 2022).
Source: Waves 2001-2007 of HILDA, PSID, SOEP and SHP.

FIGURE 2. House prices to income ratios, 1990–2015



Notes: Nominal house prices divided by nominal disposable income. The shaded area indicates the study period.
Source: OECD (2021).

eral health. As depicted in Figure 1, a 1% increase in IR corresponds to a 0.6% increase in health status in Australia, a 0.8% rise in the United States, and a 0.3% increase in Germany. However, this contribution could also be absent, as in the case of Switzerland, where IR is small, showing that the IR effects are differential and, thus, could pose comparability problems. This shows that IR, or rental payments to oneself, have a significant psychological and physiologic meterage and contribute immediately, but differentially, to well-being.

To address these potential measurement issues for country comparison, guidelines outlined by United Nations (1977) stipulate that IR should be included in the calculation of GDP, and the official agencies routinely calculate IR for this purpose (e.g., The UK Statistics Authority 2016). In numerous countries, the IR represents the most substantial element of personal consumption expenditure, underscoring the magnitude of quantities that could be easily overlooked (Australian Bureau of Statistics 2023). This highlights further that housing is fundamental to economies, but – as noted by Duca, Muellbauer and Murphy (2021) – remains outside the mainstream of academic research.

In line with its significance, IR has been employed for a wide range of purposes, including inflation measurement (Hill, Steurer and Waltl 2023), affordability measurement (Naidin, Waltl, Ziegelmeyer et al. 2022), and tax-related purposes (Fagereng, Holm and Torstensen 2020). However, when quantifying the distribution of economic outcomes, IR has only been considered in relation to poverty and income inequality.[1] IR tends to reduce measured income inequality and poverty. The effect, however, is likely not significant enough to change the ranking of countries (Ceriani, Olivieri and Ranzani 2019). Thus, it is often perceived that the influence of IR, for practical purposes, can be disregarded.

Alexeev (2020) argues that the marginal effect of IR on measured inequality provides only a limited case to ignore IR in empirical studies. He shows that housing also influences measures of intergenerational income mobility. Whereas IR reduces income inequality indexes, it often has the opposite effect on intergenerational income mobility measures. Mobility estimates were particularly affected by IR in Australia, relative to the impact in Germany and the United States.

[1] Buckley and Gurenko (1997), De Vreyer and Lambert (2020), Eurostat (2010), Ferreira and Litchfield (1999), Frick, Goebel and Grabka (2014), Frick, Grabka et al. (2010), Garner and Short (2009), Gasparini and Sosa Escudero (2004), Lerman and Lerman (1986), Naidin, Waltl, Ziegelmeyer et al. (2022), Onrubia, Rodado and Ayala (2009), Saunders and Siminski (2005), Smeeding, Saunders et al. (1993), Torrey, Smeeding and Bailey (1999) and Yates (1994).

This is an unexpected result. Given that IR reduces measured inequality in Australia (Saunders and Siminski 2005), one might infer, following the implications of the Great Gatsby Curve (which suggests that inequality correlates with immobility), that the IR would enhance mobility. However, the opposite is shown in the data. This indicates that within a more intricate analysis of economic outcomes, where factors extend beyond mere inequality, the role of the IR is impossible to predict ex-ante.

It is limiting to only focus on inequality and poverty indexes when quantifying economic outcomes as they are usually defined by the households at the income distribution tails (Cobham, Schlögl and Sumner 2016). However, since homeownership is generally associated with the middle class, it is reasonable to assume that IR would affect the position of households in the middle of the income distribution rather than the tails. Using data from Australia, Germany, the United States, and Switzerland, the current study is the first to demonstrate this. The focus is on cross-country comparisons during the 2001-2007 housing cycle, as depicted in Figure 2.

The paper unfolds as follows. Section 2 introduces the data. Imputing rents is a methodologically ambiguous exercise with new methods still emerging (Gallin et al. 2021), so instead, I rely on a Cross-National Equivalent File (CNEF). In CNEF, the methods of imputation are chosen by the data custodians. This approach has its advantages and disadvantages. The key advantage is that the chosen method is optimised for each institutional setting, making IRs most economically meaningful for each country. A disadvantage is that differences in imputation methods may compromise cross-country comparisons and their associated conclusions. The latter, however, depends on the methods used.

Section 3 employs a transition matrix to demonstrate that IR impacts the position of households in the middle of the income distribution. Transition matrices are commonly used in mobility studies (Jäntti and Jenkins 2015); my study is the first to note that these matrices are remarkably useful in investigating the distributional effects of IR. The transition matrices are particularly powerful, as they unambiguously and without any assumption on data show that the IR affects the relative position of households and, for all countries, the effects are concentrated in the middle-income household. The evidence presented in this section takes full advantage of country-specific methods of imputing rent (the economic relevance) while not suffering the potential disadvantage of the imputation methods not being uniform across countries.

Given the relatively small sample sizes of the national household survey I am working with, the transition matrices that I specify can only have three income groups. This is necessary to maintain a reasonable number of households per cell. However, to dig deeper, I define the changes in position across distributions on the household level and perform a regression analysis. The results confirm the distributional effects of IRs and allow us to conclude that IR increases the position of older households at the expense of younger ones. This confirms a theoretical connection between homeownership and the life cycle and allows us to speculate on the consequences of potential IR tax.

After establishing that IR affects middle-income households, in Section 4, I demonstrate the potential consequences of these effects, using the Esteban-Ray (ER) polarisation index (Esteban and Ray 1994). The ER index is the oldest and most cited index in the literature (Duclos and Taptué 2015), with the statistical packages for its calculation available in `Stata` (Gradin 2014) and `R` (Sohn 2019). I show that ranking based on the ER indexes is unstable throughout the housing cycle. The least polarised countries might become the most polarised from one year to another. However, equalising the tenants' rental status by adding IR to the income measure stabilises the ranking. The intuition for the results is the following. Polarisation characterises, loosely speaking, the flatness at the centre of the income distribution – the middle class. Then, without IR, the middle class (the flatness) cannot be detected (polarisation is high).

This discovery carries significance because, in a typical research scenario, scholars utilise cross-sectional data to measure polarisation and subsequently draw conclusions based on those measurements.[2] The results presented in the current study suggest that the ranking might be affected if selected countries go through a housing cycle. This a likely scenario, particularly for a country that goes through development as business and housing cycles are tightly linked (Leamer 2007, 2015). These conclusions contrast to those of Hussain (2009). His work demonstrates that neither household equivalence scales nor the inclusion of taxes in income calculations affects the rankings of income polarisation.

My subsequent investigation, presented in Section 5, is based on regression analysis of ER indexes. My study is the first to offer a regression analysis of this kind. Typically, welfare economics papers calculate indexes that incorporate or exclude IR, followed by visual comparisons across different periods and locations, similar to the analysis presen-

[2] Clementi, Dabalen et al. (2017, 2020), Clementi, Molini and Schettino (2018), Clementi and Schettino (2015), Esteban, Gradín and Ray (2007), Gochoco-Bautista et al. (2013), Gornick and Jäntti (2014), Gradín (2000), Rodas, Molini and Oseni (2019), Schettino, Gabriele and Khan (2021), Schettino and Khan (2020), Wan and Clementi (2021), Wang, Caminada, Goudswaard et al. (2017) and Wang, Caminada and Wang (2017).

ted in Section 4. I make a natural methodology step forward and instead use a regression model for the comparison.

Cross-country regressions can be unreliable if the imputation methods are not uniform across countries. To obviate this issue, I employ a panel data technique, which allows for more robust findings. This technique utilises the within-country variation, making it a suitable choice for my context as it addresses the ambiguity of IR calculations when comparing countries. In other words, as IR methods are time-invariant (fixed for a country across years), fixed effects work well. As a result, the model uses the IR values with the highest economic relevance (since they are computed by the custodians and optimal for each setting), and, at the same time, the model accounts for differences between countries fully non-parametrically with country and year fixed effects.

Regression analysis has the key advantage of producing standard errors, which allows us to evaluate statistical significance. By utilising cluster robust inference and feasible generalised squares with variance-covariance matrix with different specifications, the analysis confirms that IR has a significant and economically meaningful impact on ER indexes. However, its effect on Gini coefficients is not statistically significant. This is consistent with the evidence presented with the transition matrix, where IR changes the position of households in the middle rather than at the tails of the distribution.

After confirming that the effects of IR are concentrated on the ER index, but the Gini coefficient is not affected in a statistically or economically significant manner, I build on this discovery and formulate a triple difference (TD) framework that causally relates the housing cycle to the influence of IR on polarisation. The framework exploits variation in outcomes over time (less intense early vs. more intense late stage of the cycle), across indexes (unaffected inequality vs. affected polarisation), and between definitions of income (with vs. without IR). Consequently, the influence of IR on the measurement of economic outcomes is mediated by swings in the housing market. This, in turn, suggests that the IR component of income can generally be disregarded in empirical applications if researchers have verified that their data is not affected by the large changes in housing prices.

Overall, each method used in the study is based on different assumptions. However, they all reach a common conclusion that IR affects polarisation (as measured by the ER index) but has no effect on inequality (as measured by the Gini index), and this impact is mediated by housing cycles.

The following section discusses the data.

2. Data

2.1 *Measuring rents*

In his recent review, Balcázar et al. (2017) concludes that there is no universally applicable method of constructing IR values. An optimal approach is institutional-specific. Each country should consider the unique features of the market to equate people by rental status correctly. For example, in Switzerland, where IRs are calculated for tax purposes, each territory (cantons) uses its own methods to estimate rent. This problem mirrors the complexity of computing the comparable measure of disposable income, which, depending on the country, may, for example, include government transfers or be subject to a different tax regime.

A natural starting point involves directly querying homeowners about their estimation of what the market would pay if they were to rent out their home (e.g. Fessler, Rehm and Tockner 2016). An acknowledged concern with this approach is the potential influence of personal attachment to the property, often referred to as the 'owner pride factor.' This factor could distort estimates, particularly for homeowners who have held their properties for extended periods (Agarwal 2007; Gonzalez-Navarro and Quintana-Domeque 2009). Consequently, this method is generally considered unreliable, leading to the adoption of other techniques. I discuss two ways that are used in this study. These are hedonic regression (suitable for situations with lower homeownership rates) and user-cost (more effective than hedonic regression when dealing with high homeownership rates).

The first approach employs data from recently rented properties to estimate rents for owner-occupied or subsidised properties. This technique is known as 'comparisons' or 'rental equivalence.' In practice, a hedonic regression, $\hat{Y}_i = f(X_i, \epsilon_i)$, is frequently applied (e.g. Hill 2013). This method relies on the strict exogeneity assumption, $\mathbb{E}(\epsilon|X_i) = 0$, which implies that properties are rented out in a relatively random manner across the market and in sufficient quantity.

This assumption is more likely to hold true when the rental market is substantial, and homeownership rates are low, as seen in countries like Germany or Switzerland. However, if this method is applied to markets where only specific property types are rented out while the majority of the market is owner-occupied, the hedonic model might suffer from incidental truncation and struggle to estimate rents accurately. Additionally, there is no standardised regression specification; the preferences captured by $f(\cdot)$ are unlikely to be uniform across all markets.

For thin and imbalanced rental markets characterised by high homeownership rates (as observed in the United States or Australia), user-cost methods are often employed as alternatives to hedonic regression. Here, the IR is treated as a cost that homeowners cannot recover. Typically, the IR is calculated using the formula IR $:=$ (interest rate $+$ user costs) $\times$ P_H, where P_H represents housing prices, and a fraction of this value (the terms in the parentheses) is considered as the IR. The interest rate approximates the gains from homeownership, while other user costs encompass location-specific expenses linked to property ownership, such as property tax rates, maintenance costs, or depreciation.

An advantage of user-cost methods is that housing prices, P_H, can be drawn from a wide range of sources, including broad area property indexes (e.g. Garner and Verbrugge 2007). However, a limitation arises in that a single rate is applied to all households, disregarding variations in housing characteristics (IR varies solely by H, while in hedonic regression, it varies by i).

To complete the discussions on imputation methods, we note that these methods are rich and still emerging in literature. There are dedicated papers on imputation methods, and I suggest interested readers refer to them (Ceriani, Olivieri and Ranzani 2023). One promising recent innovation is the application of machine learning techniques (Gallin et al. 2021). Another interesting method is suggested by (Bracke 2015), which uses buy-to-let properties to infer conversion factors for calculating imputed rents. The diversity of practices stresses the complexity and potential ambiguity of imputing rents. In this study, as I now explain, I circumvent the complexity and ambiguity of estimating IR by utilising CNEF.

2.2 *Data source and descriptive statistics*

The CNEF is used for the analysis. The File adapts the national surveys to be directly comparable (Frick, Jenkings et al. 2007). CNEF datasets are available for ten countries; however, only the following four panels include IR: the Household, Income and Labour Dynamics in Australia (HILDA), the United States's Panel Study of Income Dynamics (PSID), the German Socio-Economic Panel (SOEP) and the Swiss Household Panel (SHP). The data for all panels are collected annually, except for PSID, where data is collected every other year. So 2002, 2004, and 2006 are unavailable for the United States.

The variable $X^{\overline{r}}$ represents household post-government income (referred to as income), which is the sum of all recorded sources of household income from labour, assets, private transfers, private pensions, public transfers, and social security pensions, minus total household taxes. The original PSID does not include income derived from non-refundable tax credits (the Earned Income Tax Credit) or near-cash benefit income in the form of food stamps (now called the Supplemental Nutrition Assistance Program). CNEF employs the TAXSIM model to simulate taxes to derive a harmonised income (Feenberg and Coutts 1993). Similarly, due to incomplete reporting, Schwarze (1995) methods are used to simulate the tax burden of SOEP and SHP.

The income-receiving unit of the analysis is a household. To look at the income on a comparable basis, I follow the OECD recommendations (OECD 2020) and equivalise the household income measures by dividing it by the square root of the household size. As noted in the introduction, the polarisation is known to be largely insensitive to the household equivalization methods (Hussain 2009).

TABLE 1. Descriptive statistics

	Mean	S.D	Min	0.25	Med	0.75	Max	Mean	S.D	Min	0.25	Med	0.75	Max
	Australia							Germany						
IR	6.10	8	0.00	0.63	4.20	8.50	155.00	1.10	1.8	0.00	0.00	0.00	2.00	26.00
Income	35.00	28	-2010.00	19.00	30.00	45.00	773.00	23.00	20	0.00	14.00	19.00	27.00	2003.00
+IR	41.00	30	-2010.00	24.00	35.00	51.00	871.00	24.00	21	0.00	15.00	20.00	28.00	2016.00
IR/income	0.31	4.7	-22.0	0.0	0.1	0.3	626.0	0.06	0.4	0.0	0.0	0.0	0.1	121.0
HH size	2.7	1.4	1	2	2	4	14	2.6	1.2	1	2	2	3	13
N				78,081							133,726			
	United States							Switzerland						
IR	2.90	6	0.00	0.00	0.76	3.40	246.00	0.99	2.7	0.00	0.00	0.00	0.02	121.00
Income	32.00	39	-281.00	15.00	25.00	39.00	3382.00	52.00	37	0.00	34.00	47.00	64.00	3051.00
+IR	35.00	42	-129.00	16.00	27.00	43.00	3392.00	53.00	37	0.00	34.00	48.00	65.00	3051.00
IR/income	0.18	6.3	-173.0	0.0	0.0	0.1	900.0	0.03	0.18	0.0	0.0	0.0	0.0	21.0
HH size	2.9	1.5	1	2	3	4	13	2.8	1.3	1	2	2	4	12
N				46,834							49,488			

Notes: Table reports descriptive statistics for income variables used for estimate index (3) (in thousands). IR stands for imputed rent. HH stands for household. Income stands for household disposable income. +IR stands for household disposable income plus imputed rental value. HH size is reported because nominal equivalised household income is used in polarisation measures.
Source: Waves 2001-2007 of HILDA, PSID, SOEP and SHP.

The variable X^r represents a household post-government income plus household IR value (referred to as the +IR). Similarly to income, the method of IR construction for each country is chosen by the data custodians and is optimal for a given institutional environment. Thus, using CNEF bypasses the complexity and ambiguity of estimating IR myself. This also fosters the comparability of my estimates with potential future studies.

Table 1 reports the descriptive statistics for the household-level variable used to measure polarisation. The country-level variables used in the regression analysis (produced after ER indexes are applied to the panels) are shown in Figure 5. The ratios of IR to income reported in the table are particularly interesting. For Australia, the IR stands at 31% of income, followed by the United States with 18%. These ratios are much smaller for continental European countries. It is also of note that (as shown in column 'Max') there are always households where IR is many times higher than income.

I now report more on the IR method used in CNEF.

2.3 *Imputed rent validation*

In the introduction, Figure 1 provides validation of IR as a contributor to well-being. In this subsection, as a further exploration into the nature of IR as reported in CNEF, I compare IR with the basic homeownership rates.

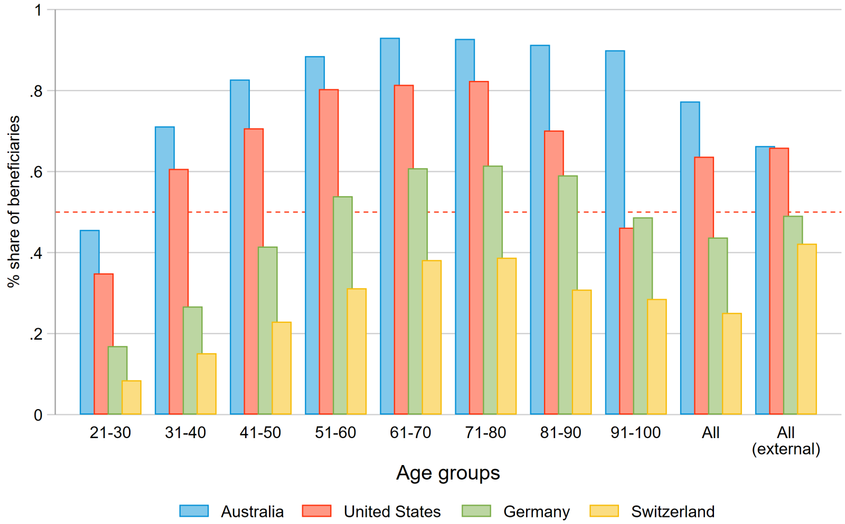
Figure 3 depicts the proportion of respondents with positive IR based on the age of the household head or partner. The right-most bar in Figure 3 shows the national homeownership rates from alternative data sources. The rates for Australia or the United States are well above 50%; in contrast, the German or Swiss rates are below 50%. These stark distinctions in the homeownership rates confirm that each country's rental market performs differently and, as discussed in Section 2.1, different methods of calculating IR should be chosen for each country. This further affirms that choosing a custodian-selected and institution-specific method of imputing is likely the best approach to reflect the economic content of IR optimally.

In CNEF PSID, IR is constructed with the user-cost method according to the formula $6\% \times (P_H - \textit{remaining mortgage principal})$. The percentage reflects the United States' cost of capital and other ownership costs such as taxes, maintenance and depreciation. Since this method ignores subsided tenancy, the share of IR beneficiaries in the data for the United States should correspond to homeownership. This correspondence is evident from the equal heights of the two right-most bars representing the United States in Figure 3.

The HILDA dataset employs a similar approach for homeowners, though with the use of 4% instead of 6%, reflecting differences in cost. However, HILDA also considers IR for subsidised tenants. In cases of public housing, a comparison method is used to estimate the market value of a similar property. Then the actual payment is subtracted from this estimate, resulting in $\hat{Y}_i - \textit{Actual}$. As a result, the right-most bar in Figure 3 is slightly lower than the share of IR beneficiaries. This observation reflects the role of public housing in Australia's housing market, which is not as pronounced in the United States.

It's worth noting that despite a smaller fraction of homeownership being treated as IR (4% instead of 6%) in the Australian data, the share of IR in income (as shown in Table 1) is still almost twice as high in Australia (31%) compared to the United States

FIGURE 3. Share of IR beneficiaries



Notes: Share of respondents with positive household imputed rent by the age of household head or partner. The right-most bar shows homeownership rates taken from alternative databases. The red dashed line shows 50%. To account for years, the shares are calculated by regressing age group dummies on the indicator function for positive IR with the year effects included but intercept excluded. The coefficients on age group dummies are shown in the figure. For all age groups, the age group dummies are replaced with intercept. For all (external), the annual average for 2001-2007 is taken from the official web pages.
Source : Waves 2001 and 2007 of HILDA, PSID, SOEP and SHP.
Source (external): For Switzerland and Germany, the EU Statistics on Income and Living Conditions; For the US, the U.S. Census Bureau; For Australia, the Australian Bureau of Statistics.

(18%). This could be attributed to Australia's uniquely high housing prices relative to wages. Additionally, an interesting pattern emerges in the Australian data in Figure 3, where the fraction of people with positive IR does not decrease for older age groups. Moreover, the difference in the number of beneficiaries between Australia and other countries becomes more pronounced as age increases, as evidenced by the increasing height of the blue bar relative to the other bars along the horizontal axis.

User-cost methods are also preferred for Anglo-Saxon countries because it is common to use housing as a financial asset in these countries. This, in turn, contributes to (or originates from) excessive volatility of the housing cycle (Voigtländer 2014). Using methods that allow for negative values is thus important for Australia and the United States. In contrast, the Swiss and German housing markets are among the most stable globally, with developed rental markets.

As a result, SOEP and SHP use hedonic regression to estimate IR. The hedonic characteristics vary slightly by country, but the general procedure is similar. For owners, the hedonic regression is used to estimate a hypothetical rental income, then costs associated with owning are subtracted from it, $\hat{Y}_i - (\textit{user costs} + \textit{interest rate})$. If costs exceed the income advantage, which is common at the beginning of the mortgage repayment period, zero is recorded. The IR for subsidised tenancy is just a difference between the estimated market and the actual rents, $\hat{Y}_i - Actual$. These qualities of this IR method are also reflected in Figure 3, where the share of IR beneficiaries is slightly lower than the raw homeownership rate.

I now present the empirical results. The empirics is presented in three separate sections. In the Positional evidence section, I begin with evidence that IR influences the households' positions in income distribution. This holds true at both the distribution level and for individual households. Additionally, I examine the socio-economic predictors of transitions. In the subsequent Index-based evidence section, I present the ER indexes, illustrating that including IR leads to greater stability in country rankings. In the last section, Regression-based evidence, I provide the regression results based on the ER indexes.

3. Positional evidence

Here, I delve into individual household data and show that IR affects distribution (shown by kernel density estimates) and the relative placements of individual households in income distribution (shown by transition matrices). Moreover, by approximating the socio-economic profile of households, I show that older households are moved higher in distribution while younger ones are lower.

3.1 *Measuring income group transitions and their predictors*

A transition matrix is a tool that characterises changes in the relative position of individual households. There are two ways to formally introduce the matrices (Jäntti and Jenkins 2015).

A transition matrix, P, is constructed by first dividing $X^{\overline{r}}$ and X^r into three tertiles and cross-tabulating the relative frequencies of observations with each matrix cell: element p_{ij} is the relative frequency of observations of X^r in tertile i and $X^{\overline{r}}$ in tertile j. The graphical representation of the discrete joint probability density function is the bivariate histogram. Although, I opt for a tabular presentation.

Alternatively, the transition matrix and two distributions may represent the transition process. To maintain a reasenable number of households per matrix cell, I specify 3 income ranges, with the relative number of observations in a tertile group k in X^r is q_r^k for $k = 1, 2, 3$, and correspondingly in $X^{\overline{r}}$. The marginal distribution X^r is summarized by the vector $q_r = \left(q_r^1, q_r^2, q_r^3\right)$ and correspondingly for X^r. Hence,

$$q_r^k = q_{\overline{r}}^k A \tag{1}$$

Each group contains one-third of the population. The transition matrix is then bistochastic. The relative change in income distribution is then entirely characterised by the transition matrix A.

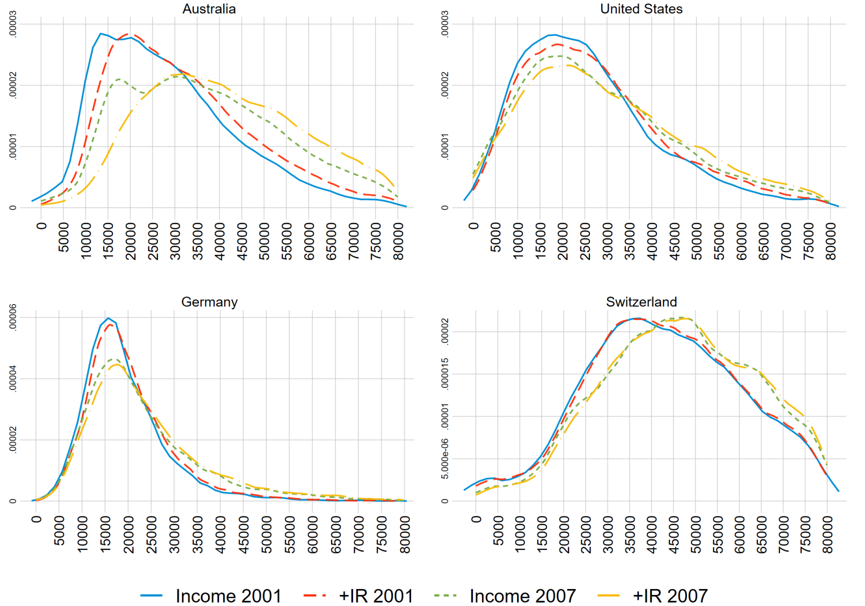
The transition matrices show that IR has the most significant influence on the relative positions of middle-income households, fully supporting the paper's claims. This raises the question of whether there are more granular characteristics that define individuals whose positions change due to the inclusion of IR. To understand the characteristics that predict an income-group transition, I calculate the rank of household i in year t in the distributions $X^{\overline{r}}$ and X^r and then subtract the latter rank from the former. This gives a household-level measure of change in relative income position denoted by R_{it}. To offer further insight into these transitions, I specify the following model:

$$R_{it} = g_1(\mathit{Age}_{it}) + g_2(\mathit{Schooling}_{it}) + \varepsilon_{it}, \tag{2}$$

where R_{it} represents the change in rank due to IR, and g_i are unknown functions estimated using a third-order B-spline basis. Cross-validation selects the optimal number of terms in the basis function (Chen 2007). This approach is chosen because the task here is inherently descriptive, and using a nonparametric serial estimation ensures that data patterns are fully appreciated. Schooling is chosen as a rough proxy for socio-economic status, while age is chosen as it is unambiguously exogenous and consistently shown to be a reliable predictor of homeownership (e.g., Malmendier and Steiny 2017).

Other variables commonly considered predictors of homeownership include marital status, presence of children, and employment status (e.g., Fu 2014). However, these variables are not included in our analysis due to their likely correlational relationships with transition. By construction, homeownership is the immediate causal predecessor of IR. Investigating the causal effect beyond this relationship would require a deeper understanding of complex lifecycle decision-making. This level of causal reasoning demands a more robust estimation design and purposeful theoretical analysis, which falls outside the scope of this paper (but could be a promising area for future research).

FIGURE 4. Income distribution



Notes: The densities are estimated using Epanechnikov kernel with a width that minimises the mean integrated squared error if the data were Gaussian. See Table 1 for descriptive statistics of the data used.
Source: Waves 2001 and 2007 of HILDA, PSID, SOEP and SHP.

3.2 *Results*

Figure 4 gives the visual representation of the income and +IR variables. For comparability, the same fixed range is shown. IR value tends to shift the distributions to the right with ambiguous effects on the spread and centre of the distributions. Visual inspection of densities allows anticipating the effect of IR value on the polarisation and inequality measurements. Inequality indexes capture the spread; in contrast, polarisation indexes capture shifts at the centre.

IR values are unlikely to change measured inequality substantially, as the spreads are largely unaffected. The most considerable change is seen for Australia, suggesting that Australia's Gini coefficient is likely to be most responsive to IR. As for the shape of the centre of the distributions, IR values affect all countries, but to a different extent. The Australian distribution again stands out.

Without IR values, the middle of income distribution develops a notch in 2007 (dashed green line), which is also somewhat visible in 2001 (solid blue line). This notch is gradually getting more prominent throughout the housing cycle. Therefore, the polarisation measures are likely to reduce substantially for Australia following the inclusion of IR values. Noteworthy is that this change is driven by the indentation in the income variable and not by the shape of the +IR variable. That is, the same results would be true if other methods of IR value imputation were used for Australia.

Densities are sensitive to bandwidth and kernel choice and non-informative about the effect of IR on individual households. Transition matrices can further investigate whether IR preserves the households' ranking in the income distribution.

TABLE 2. Income group transitions

+IR	Income Low	Income Middle	Income High	Income Low	Income Middle	Income High	
	Australia			United States			
Low	88.385%	10.768%	0.847%	95.018%	4.982%	0.000%	Wave 2001
	85.929%	13.304%	0.766%	93.620%	5.816%	0.564%	Wave 2007
	-2.455%	2.536%	-0.081%	-1.398%	0.833%	0.564%	Difference
Middle	11.593%	79.257%	9.150%	4.984%	89.305%	5.711%	Wave 2001
	14.071%	75.226%	10.704%	6.380%	87.558%	6.061%	Wave 2007
	2.478%	-4.031%	1.554%	1.397%	-1.747%	0.350%	Difference
High	0.000%	9.975%	90.025%	0.000%	5.711%	94.289%	Wave 2001
	0.000%	11.473%	88.527%	0.000%	6.627%	93.373%	Wave 2007
	0.000%	1.498%	-1.498%	0.000%	0.916%	-0.916%	Difference
	Germany			Switzerland			
Low	91.659%	8.341%	0.000%	97.061%	2.939%	0.000%	Wave 2001
	92.393%	7.558%	0.049%	97.238%	2.762%	0.000%	Wave 2007
	-0.734%	0.783%	-0.049%	-0.177%	0.177%	0.000%	Difference
Middle	8.323%	85.202%	6.475%	2.979%	95.090%	1.930%	Wave 2001
	7.608%	86.641%	5.751%	2.723%	93.771%	3.507%	Wave 2007
	0.714%	-1.438%	0.724%	0.257%	1.320%	-1.576%	Difference
High	0.000%	6.436%	93.564%	0.000%	1.890%	98.110%	Wave 2001
	0.000%	5.801%	94.199%	0.000%	3.548%	96.452%	Wave 2007
	0.000%	0.635%	-0.635%	0.000%	-1.658%	1.658%	Difference

Notes: The table shows the cross-tabulation of the relative frequencies of income with imputed rent included (+IR) and excluded (Income). For descriptive statistics of the data used, refer to Table 1. For a formal introduction to transition matrices, see Section 3.1. The calculations were performed using the package developed by Savegnago (2016).
Source: Wave 2001 and 2007 of HILDA, PSID, SOEP and SHP.

Table 2 reports transition matrices for all countries for waves 2001 and 2007. For example, the top left part of the table reports two transition matrices for Australia - one matrix for wave 2001 and another for 2007. The bottom of each cell also reports the difference between the matrices - the change between waves. If IR did not affect household standing in the income distribution, all values would be concentrated at the diagonal. That is not the case.

As shown in the middle cell of the Australia matrix for wave 2001, only 79.257% of the households belonging to the middle-income group when IR is included maintain their position when IR is excluded – 11.593% transition into the bottom and 9.15% into the top-income tertile. In 2007, when the housing prices reach a new level, the share of middle-income groups that maintain their position without IR is further dropped by 4.031%. As for other income groups, a significantly larger fraction of households maintains their position in the distribution.

TABLE 3. Predictors of change in income position

	(1)		(2)		(3)		(4)	
	Dependent variable (DV): Increase in rank							
	Australia		United States		Germany		Switzerland	
	Estimate	SE	Estimate	SE	Estimate	SE	Estimate	SE
	Average marginal effects							
Age	0.241***	(0.00238)	0.0916***	(0.00176)	0.0723***	(0.000964)	0.0383***	(0.000951)
	Marginal effects at various age points							
20	-6.455***	(0.0906)	-1.703***	(0.0619)	-1.107***	(0.0648)	-0.651***	(0.0468)
30	-5.007***	(0.0582)	-2.145***	(0.0265)	-1.599***	(0.0219)	-0.769***	(0.0202)
40	-2.247***	(0.0720)	-1.080***	(0.0324)	-1.264***	(0.0159)	-0.521***	(0.0195)
50	0.195**	(0.0750)	-0.139***	(0.0380)	-0.305***	(0.0194)	-0.182***	(0.0255)
60	4.185***	(0.113)	1.673***	(0.0631)	1.001***	(0.0208)	0.380***	(0.0422)
70	6.802***	(0.137)	3.984***	(0.103)	2.154***	(0.0330)	1.474***	(0.0624)
80	7.807***	(0.171)	5.142***	(0.152)	2.611***	(0.0505)	1.528***	(0.0914)
	Average marginal effects							
Age	0.257***	(0.00261)	0.0893***	(0.00188)	0.0724***	(0.00100)	0.0379***	(0.000977)
Schooling	0.822***	(0.167)	0.162***	(0.0251)	-0.00176	(0.00978)	-23.81	(16.88)
	Marginal effects at various age points							
20	-6.484***	(0.0960)	-1.416***	(0.0731)	-1.182***	(0.0703)	-0.543***	(0.0551)
30	-5.411***	(0.0615)	-2.194***	(0.0277)	-1.583***	(0.0229)	-0.769***	(0.0213)
40	-2.534***	(0.0739)	-1.084***	(0.0333)	-1.236***	(0.0168)	-0.537***	(0.0203)
50	0.0394	(0.0781)	-0.169***	(0.0384)	-0.293***	(0.0201)	-0.190***	(0.0256)
60	4.462***	(0.116)	1.649***	(0.0640)	1.002***	(0.0213)	0.378***	(0.0422)
70	7.436***	(0.143)	4.057***	(0.104)	2.140***	(0.0337)	1.484***	(0.0625)
80	8.914***	(0.182)	5.301***	(0.153)	2.561***	(0.0513)	1.548***	(0.0920)
	Marginal effects at various schooling points							
5	-2.853***	(0.237)	-1.162***	(0.145)	-0.00904	(0.103)	.	.
9	-1.360***	(0.0807)	-0.976***	(0.0577)	0.0685*	(0.0332)	-0.202***	(0.0362)
12	0.117*	(0.0543)	-0.235***	(0.0319)	0.0678**	(0.0214)	0.0584***	(0.0172)
18	1.084***	(0.145)	-0.712*	(0.354)	-0.242***	(0.0275)	-0.0374	(0.0318)
N	74124		44990		125983		49401	
DV mean	0.000		-0.320		0.000		0.000	
DV SD	9.487		4.981		4.648		2.779	
DV min	-10.54		-6.303		-5.098		-2.000	
DV max	98.30		87.30		57.63		34.37	

Notes: The table reports results from Model (2) estimated with cubic B-spline. The top table reports results when only age is included, while the bottom table additionally includes schooling.
Source: Waves 2001-2007 of HILDA, PSID, SOEP and SHP.

The same pattern can be seen in other countries (the middle cells on the diagonals are always the smallest and get smaller over the study period). IR values universally affect the standing of middle-income groups. This directly indicates that polarisation, as opposed to inequality measures, should respond the most to including IR values in the income.

It is also interesting to note that the richest households occasionally become the poorest, but none transition in the opposite direction. In Australia, the transition from the richest to the poorest occurs in both 2001 and 2007 (0.847% and 0.766%). For other countries (except for Switzerland), this occurs only in 2007. We see this even when the imputation methods for Germanic countries do not permit negative values. This dynamic is present because adding IR to income changes the relative position of other households.

Table 3 presents the predictors of transition. I begin by conducting a regression, using a nonparametric function to specify the relationship between rank increase and age (schooling is omitted for now). The average marginal effects for four countries are reported at the top line of the table. On average, a marginal increase in age leads to a much higher change in relative position in Australia compared to other countries.

Subsequently, I utilise the estimated nonparametric function to predict the effects for various ages, and the results are shown in the bottom half of the table's top section. A young age predicts a substantial reduction in position, whereas older age predicts an increase. Interestingly, for all countries, the coefficient changes the sign at the age of 60, while in Australia, this change occurs at the age of 50 years.

At the bottom of Table 3, I repeat the analysis, but this time, households are stratified by the number of years of education of the households' heads and partners. The new estimates on education for Australia stand out once again in magnitude. Interestingly, stratification by education produces no changes in the effects of age. Notably, for Australia, age 50 appears to be an exception. When schooling is included in the analysis, the effect becomes statistically insignificant.

It is interesting to observe that schooling has no predictive power for Germanic countries on average. When schooling is investigated on various levels, the pattern shows that more educated individuals increase their position at the expense of those with less education, but only in Australia and Switzerland. In Germany, this redistribution along the levels of education goes in the opposite direction. Meanwhile, for the United States, higher education is generally protective against reduction in position (as in Australia and Switzerland); however, for all levels, IR reduces position. The Swiss prediction for 5 years of schooling is missing, as the panel reports 9 as the minimum number.

What is clear is that the positional effects of IR are directly linked to the lifecycle, while the role of socio-economic status, roughly proxied by the number of years of education, remains ambiguous. The ambiguity observed in the data may indeed reflect genuine distinctions across countries. Or it could be attributed to the fact that education is a choice subject to various sources of country-specific endogeneities, which my approach does not resolve.

4. Index-based evidence

In this section, I apply the ER index to national panels to construct an annual country-level measure of inequality or polarisation. I then demonstrate the increased stability of rankings when the IR is incorporated into the income measure.

4.1 *Measuring polarisation*

The initial components of the analysis are two distributions of income drawn from each country. One distribution excludes IR value and is denoted by $X^{\overline{r}} \in \mathbb{R}^n$, where n is the number of households. The second distribution includes the IR value and is denoted by $X^r \in \mathbb{R}^n$. The polarisation is then calculated using the ER polarisation index developed by Esteban and Ray (1994):

$$Y_{a,m}(X_m) = \frac{1}{2\mu_m} \sum_{i=1}^{n} \sum_{j=1}^{n} |x_{i,m} - x_{j,m}| (\pi_{i,m})^{1+a} \pi_{j,m} \qquad m \in \{\overline{r}, r\}, \;\; a \in \{0, 1.6\}. \quad (3)$$

In the index, a represents a polarisation sensitivity factor (chosen by the researcher), x_m represents the individual value of X_m normalised by twice its mean value (by the scalar in front of the expression, which has no bearing on ranking), π represents the proportion of households with two or more observations with the same value multiplied by cross-sectional weights (neglected in the formula for expositional purposes). Thus, the number of income groups equals the number of distinct values of X^m. When $a = 0$, and because of the scaling chosen, the index yields a sample-weighted Gini coefficient. When $a > 0$, the index captures the clustering of X^m around income groups.

For $a \in [0, 1.6]$, the ER index satisfies a collection of axioms founded upon a so-called Identification–Alienation nexus wherein notions of polarisation are fostered jointly by an agent's sense of increasing within-group identity and between-group distance or alienation. The four axioms may be loosely summarised as follows (Anderson 2016). Axiom

1: A mean preserving reduction in the spread of a distribution cannot increase polarisation; Axiom 2: Mean preserving reductions in the spread of sub-distributions at the extremes of a density cannot reduce polarisation; Axiom 3: Separation of two sub-densities towards the extremes of the distributions range must increase polarisation; Axiom 4: polarisation measures should be population-size invariant.

The calculation of ER indexes and their ordering would already demonstrate the key finding of this paper. IR influences polarisation measures to the extent of rank-reversing the cross-country comparison. In contrast, the measures of inequality are barely affected.

TABLE 4. Results: ranking of countries by polarisation

		2001	2002	2003	2004	2005	2006	2007
Ranking by ER index with polarisation factor 0 (Gini coefficient)								
Income	Australia	2	2	2	2	2	2	2
	United States	1	1	1	1	1	1	1
	Germany	4	4	4	4	3	3	3
	Switzerland	3	3	3	3	4	4	4
+IR Value	Australia	2	2	2	2	2	2	2
	United States	1	1	1	1	1	1	1
	Germany	4	4	4	4	3	3	3
	Switzerland	3	3	3	3	4	4	4
Ranking by ER index with polarisation factor 1.6								
Income	Australia	3	2	1	1	1	1	1
	United States	2	3	3	2	3	3	3
	Germany	4	4	4	4	4	4	4
	Switzerland	1	1	2	3	2	2	2
+IR Value	Australia	3	3	3	3	2	2	2
	United States	2	2	2	2	3	3	3
	Germany	4	4	4	4	4	4	4
	Switzerland	1	1	1	1	1	1	1

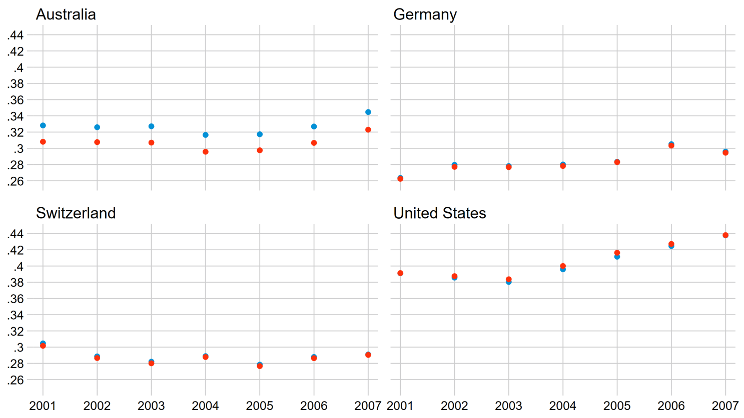
Notes: Table reports ranking of estimates reported in Figure 5. First place refers to the largest estimate (i.e., most polarised or unequal).
Source: Waves 2001-2007 of HILDA, PSID, SOEP and SHP.

4.2 *Results*

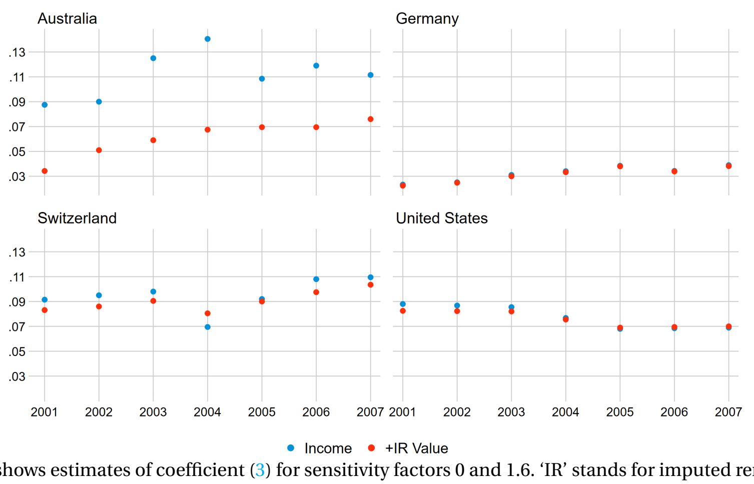
Figure 5 visualises the estimates of the ER index. The top figure visualises the estimates with a polarisation factor of 0 – the Gini coefficient. The bottom reports the estimates with a polarisation factor of 1.6. For example, the ER index with a polarisation factor of 1.6 for Australia in 2007 with income measure without IR is 0.1115; the inclusion of IR drops it by 31.84% to 0.0760.

FIGURE 5. Results: estimates of polarisation

(a) ER index with polarisation factor 0 (Gini coefficient)



(b) ER index with polarisation factor 1.6



Notes: The figure shows estimates of coefficient (3) for sensitivity factors 0 and 1.6. 'IR' stands for imputed rent, 'Income' stands for household disposable income, and '+IR' stands for household disposable income plus imputed rental value. For United States data for 2002, 2004, and 2006, the polarisation estimates are the average between two adjacent years. The estimates of the ER index with a sensitivity factor of 1.6 are multiplied by 100,000. See Table 1 for descriptive statistics of the data used. A package by Gradin (2014) is used in calculations.
Source: Waves 2001-2007 of HILDA, PSID, SOEP and SHP.

Table 4, instead of the estimates, reports the ranking of countries for a given year and the income measure. For the United States, data for 2002, 2004, and 2006 are not collected. To rank countries for all years continuously, polarisation estimates for missing years are taken as an average between two adjacent years. For example, for 2002, the value is a simple average between 2001 and 2003. This corresponds to an imputation method performed with parameter linear regression for continuous variables.

According to the Gini coefficient, Germany is the most equal society, followed by Australia, Switzerland, and the United States. IR values do not affect these countries' equality rankings. The inclusion of an IR value decreases inequality in Australia, the same result has been shown before by Saunders and Siminski (2005) and Yates (1994). A reduction is also seen for Germany and Switzerland. For Germany, Frick, Goebel and Grabka (2014) found a similar decreasing effect using the 2003 wave of SOEP. I could not find previous papers that I could use to compare my results for Switzerland. Similar to the results of Garner and Short (2009), IR marginally increases inequality in the United States.

The bottom of the table reports the polarisation results. Germany is the least polarised country, whether IR values are present or not. The United States is the following least polarised country without IR values. The results reflect those of Duclos, Esteban and Ray (2004), who show that density in the United States is unique for its flatness in the middle. The inclusion of IR values changes the position of the United States, but only because of large changes in estimates for Australia and Switzerland.

The changes for Australia stand out not only in size but also because the ranking changes before and after 2003. Without IR and before 2003, Australia is almost the least polarised society; however, after 2003, it becomes the most polarised. This change coincides with the Australian housing prices reaching a new historically high level. With IR, the Australian polarisation ranking does not change from one of least polarised to the most polarised in one year. Including IR allows offsetting the effects of the housing cycle on the ranking.

5. Regression-based evidence

In this section, I further analyse the data presented in Figure 5. The data varies by country, year, measure (inequality or polarisation) and IR (included or excluded), which makes it suitable for regression-based analysis. I demonstrate that polarisation bears a greater impact of IR than inequality. Using this finding, I apply a TD framework to connect the estimates with the housing cycles.

5.1 *Establishing significance*

The following two-way fixed effect model is used to reaffirm the effect of IR

$$Y^a_{ctm} = \delta_c + \gamma_t + \beta^a D^r_m + \varepsilon_{ctm} \qquad a \in \{0, 1.6\}. \tag{4}$$

Here, Y^a_{ctm} is the ER index with factor a (superscripted a indicates that polarisation and inequality are estimated separately), for country c, in year t, with income measure m. The variable D^r_m is an indicator function that takes on value 1 for outcomes that include IR. The parameters δ_c and γ_t are, respectively, country and year fixed effects.

The parameter β is of interest, and it shows the effect of including IR into income on calculated ER indexes. The country fixed effects are important as they absorb differences in the country-specific calculation of IR and all other time-invariant confounders. The time fixed effects, in addition, allow leveraging the information across different years.

The correct inference should account for serial correlations. One way to achieve this is to use a heteroskedastic- and cluster-robust variance matrix, where clusters are countries. This allows for autocorrelation within countries (including across income with IR and without IR). In my setting, the number of countries is too small for cluster-robust inference. Instead, I follow the advice of Cameron and Miller (2015) and estimate Equation (4) with a generalised least squares (GLS) estimator.[3]

I use two different error structures when specifying GLS. Firstly, I set the variance to be proportional to the inverse of each panel's sample size. The reason behind this choice is that the ER measurements denoted by Y^a_{ctm} are computed from samples of varying sizes, leading to differences in precision across panels. Secondly, I set the panel units at the level of country-measure (e.g., Australia with and without IR considered separate units) and specify an autocorrelation structure unique to each panel unit. This allows for distinct autocorrelation for measurements with IR and without IR.

Finally, I also demonstrate that the results remain robust, albeit with reduced precision, when using standard errors based on three types of heteroskedasticity-consistent covariance matrices (HC1, HC2, and HC3). While these variance structures are robust to heteroskedasticity (permit arbitrary elements on the main diagonal of the variance matrix), it is important to note that they may compromise the estimates' consistency due to potentially unaccounted autocorrelation (may incorrectly assume that elements off the main diagonal are zeros).

To foreshadow the finding, Model (4) shows that IR does affect polarisation but not inequality. These estimates establish the effect of IR on measured polarisation but leave the link to the housing cycle unexplored.

[3]In his practitioner's guide to cluster-robust inference, Cameron and Miller (2015) notes that 'It is remarkable that current econometric practice with clustered errors ignores the potential efficiency gains of FGLS.'

5.2 *Establishing causality*

To establish the causal link with the cycle, I compare the late, more intense stage of the cycle with the earlier, less intense stage. This model compares trends shown in Figure 5. In this figure, the rise in intensity is vividly demonstrated by Australia, where the housing cycle shows an increase in polarisation around 2003. This jump is driving the

Using Model (4), described above, I will know that inequality, unlike polarisation, is the outcome that is not affected by IR. In total, I have the treatment (inclusion of IR) that affects one outcome (polarisation) in the affected period (2003 onwards); thus, the following TD framework can be used to pin down the causal link. Namely,

$$Y_{atcm} = \gamma_{at} + \nu_{ac} + \delta_{ct} + \beta D^r_m \cdot D^{\geq 03}_t \cdot D^p_a + \varepsilon_{atcm}, \tag{5}$$

where Y_{atcm} is the measure a, for country c, in year t, with IR excluded or included m. The variables D^r_m and $D^{\geq 03}_t$ are the indicator functions for, respectively, outcomes that include IR and for outcomes for 2003 and after. The variable D^p_a is an indicator function that takes on the value 1 for polarisation (ER index with factor 1.6) and zero for inequality (ER index with factor 0). Parameters γ_{at}, ν_{ac} and δ_{ct} are interacted fixed effects, which are, respectively, measure-year, measure-country and country-year.

This model estimates the effect of the cycle on the effect of including IR in polarisation. More preciously, the model estimates the Average Treatment Effect on Treated, which, on the one hand, underscores the limitation of the results' external validity; on the other, these estimates are adequate for establishing the desired link between the cycle and the effect of IR on polarisation.

Although two contrasts are used, TD requires only one common trend assumption in the subsumed double difference (Olden and Møen 2022). As Figure 5a shows, this assumption is warranted in my setting, at least for the Gini coefficients along the dimension that includes or excludes IR (red and blue dots). Rephrasing Gruber (1994), an alternative way to think about the identifying assumption is that there is no year-country-specific shock in 2003 that affects polarisation and inequality differentially. One of the key benefits of TD is that the interacted fixed effects absorb spillovers. An alternative to TD could be a double difference where the data on polarisation is only used. Then, the variation across measures with IR and without is used to measure the effect. A potential issue is that for the same country, there are likely to be spillovers across polariton measures with and without IR. TD takes better care of that (Olden and Møen 2022).

The choice of using 2003 as the boundary between the cycle's late (more intense) and early (less intense) stages could potentially be seen as arbitrary. As a robustness check, I demonstrate that the sought effect remains in place when I model the treatment as a dose-response function (Callaway, Goodman-Bacon and Sant'Anna 2021). In this approach, I replace the dummy variable for the period from 2003 onwards with a trend variable. This trend variable takes on values of 1 for 2001, 2 for 2002, and so on. This approach is informed by the notion that the intensity of the housing cycle is escalating over time and, consequently, the effect should follow a similar trend.

5.3 *Results*

Table 5 reports the econometric model estimates. Column (1) corresponds to Model (4) when data is limited to Gini indexes. Column (2) refers to the same model applied to polarisation indexes. All regressions are based on the indexes shown in Figure 5. The results further reaffirm the above conclusion: IR influences polarisation indexes but not inequality. The benefit of the regression model is that country-specific IR calculation methods are accounted for with country-fixed effects.

Column (3) reports estimates of our TD framework, formalised in Equation (5). Indeed, the housing cycle is the reason why IR affects polarisation. The bottom of the table also reports the estimated effect relative to the dependent variable's mean value. The effect of IR on inequality is not only statistically insignificant but also small, whereas the effect on polarisation is both economically and statistically significant.

Using 2003 as a boundary between the cycle's late (more intense) vs. early (less intense) stage is potentially arbitrary. Then, instead of this conventional discreet specification, I estimate a dose-response TD (Callaway, Goodman-Bacon and Sant'Anna 2021). Specifically, a dummy for 2003 onwards is replaced with a trend variable that assigns 1 for 2001, 2 for 2002, and so on. The idea is that the cycle is exacerbating over time; thus, the effect should be, too.

The assumption for a dummy TD is that the common trend holds in at least one dimension. Figure 5 shows, for example, that Gini coefficients follow the same trend prior to 2003. The assumptions for continuous TD are stronger. The common trend holds true for every increment in dose. What is important is that either specification supports the main conclusion. Coupled with transition matrices and logical reasoning (that homeownership is a hallmark of the middle class), the regression results should not be surprising.

TABLE 5. Results: the effect of IR on ER

	(1)	(2)	(3)	(4)
	Dependent variable (DV): ER index			
	Effect on		Triple difference	
	inequality	polarisation	dummy	dose-response
	OLS - robust variance			
Estimate	-0.00522	-0.0144	-0.0142	-0.00217
HC1	(0.0145)	(0.00421)**	(0.00535)**	(0.000779)**
HC2	(0.0145)	(0.00421)**	(0.00544)**	(0.000711)**
HC3	(0.0156)	(0.00470)**	(0.00790)+	(0.00128)+
	GLS - variance inversely proportional to panels' sample size			
Estimate	-0.00401	-0.0127**	-0.0131**	-0.00199**
SE	(0.00214)	(0.00432)	(0.00542)	(0.000773)
	GLS - country-measure-specific autocorrelation structure			
Estimate	-0.00331	-0.0131**	-0.0133*	-0.00189**
SE	(0.00298)	(0.00501)	(0.00599)	(0.00069)
N	56	56	112	112
Adj. R^2	-0.109	0.708	0.992	0.992
DV mean	0.323	0.0718	0.197	0.197
% change	-2%	-20%	.	.

Notes: Table reports estimates from Model (4) and (5). Figure 5 shows the data used in the models.
+ $p < 0.10$, * $p < 0.05$, ** $p < 0.01$, *** $p < 0.001$.
Source: Waves 2001-2007 of HILDA, PSID, SOEP and SHP.

The inference in my setting is challenging because of fewer observations. Therefore, I report results using different variance estimators. The top of the regression table reports standard errors estimated with the typically used robust variance-covariance matrix (HC1). It is known that it may perform poorly with a small sample size (MacKinnon and White 1985). Therefore, I also report more robust standard errors (HC2 and HC3).

Using robust errors that permit arbitrary covariance reduces the risk of type I error (guard against false positives) at the cost of elevating type II error (may miss the true effect), particularly for smaller samples. Indeed, the results are only marginally statistically significant when HC3 is used. The alternative to robust error is more explicit modelling of the variance-covariance matrix. The middle of the table reports the results when the models are estimated with GLS, and the heteroskedasticity is assumed to be a function of the panels' sample size. Lastly, the bottom reports results that permit a more flexible autocorrelation structure.

6. CONCLUSION

This research uses four harmonised national panels from Australia, the United States, Germany, and Switzerland to provide a counterexample of using income polarisation disregarding the role of housing. Income polarisation can change throughout the housing cycle, affecting the country's ranking. Including IR stabilises the ranking, supporting the claim that the under-coverage of IR prevents correct comparison of economic outcomes across time and space (Canberra Group 2011).

Researchers who claim that the middle class is disappearing based on the evolution of polarisation indexes (Dallinger 2013; Foster and Wolfson 2010; Gornick and Jäntti 2014; Jaimovich 2020; Jenkins 1995; Pressman 2007; Simonazzi and Barbieri 2016) may need to check if the area is going through a housing cycle (or if there are steep differential treats in housing price across areas of interest). Polarisation attempts to characterise the middle class, where the transition into homeownership is most likely.

Therefore, the measures of income that do not equalise income-receiving units by rental status risk misinterpreting the evolution of measured polarisation. If this is not feasible, researchers should explicitly distinguish between pre-housing and post-housing disposable income, especially in countries facing housing affordability issues.

The finding of this paper may also be noteworthy in the context of taxation (Poterba and Sinai 2008). The results, in particular, advise against using tax data, as IR is typically exempt from taxation. Further, they remind us that non-neutral treatment of rental and owner-occupied housing for tax purposes (so-called 'homeownership bias') may be distortionary and unfair (Figari et al. 2017).

If IR is ignored, homes might become legalised 'offshore accounts' that allow hiding income. For example, Cho and Sane (2013) show that this is a common method to access the Australian means-tested government system of benefits, as occupied accommodation is exempted from those tests. As a result, housing, which is already a privilege of higher-income households, may give a tax advantage, and, as shown by Alexeev (2020), this advantage carries across generations.

I also found that in all countries, older people tend to move to higher-income positions, while younger individuals tend to move to lower-income positions when IR is included in income measures. Thus, taxing IR will, at least initially, reduce the tax burden on younger households while increasing the burden on older generations. This policy approach could potentially mitigate the ongoing Australian housing affordability crisis, as the younger population is the most affected by the crisis (Ghasri et al. 2022). However, my results are descriptive and ignore the lifecycle incentives, and more work is needed to rationalise IR taxation.

Along the same lines, ignoring IR may also lead to an unfair allocation of the tax burden. For example, Australian transition matrices show that up to 14% of households that appear to be the poorest (and, as a result, may pay less income tax) should be classified as middle- or high-income if IR is included. It is also of note that in Switzerland, the only country in this study that taxes IR, most households preserve their standing in the income distribution if IR is accounted in income. However, my work offers no evidence that taxation is the reason for this outcome (stressing again against using my finding as a rationale for more taxation).

In Australia, stamp duty is somewhat similar to IR taxation. However, how the tax is levied punishes movers, reduces labour mobility, and complicates downsizing and upsizing, distorting optimal land utilisation. Breunig (2023) makes a convincing case to replace stamp duty with an annual land tax based on the unimproved value of the land. This approach is even closer to taxing IR than stamp duty and is likely to be less distortionary.

While the role of housing is undeniable in light of the evidence in the current work, it is still unclear what the results would be if all non-monetary components were included. In particular, the IR values are high relative to income in Australia, but if the value of Australian public healthcare and education is further added to income, the IR share would be reduced, and, as a result, the effects of IR might be reduced too.

There is one explicit limitation of the work, which is that I only use the ER index to measure polarisation or inequality. This is done in the spirit of providing a counterexample. The paper does not exclude the possibility that other indexes might accommodate housing cycles better. Another potential limitation of the study is that the extreme housing cycle has been used to make the counterexample. The growing consensus is that the housing cycles of the early 2000s are historically unique (Chodorow-Reich, Guren and McQuade 2021). It is not clear whether a similar cycle will happen again. Perhaps under normal price fluctuations, the role of housing in polarisation is less important than shown in the current study. Thus, I encourage the researchers to replicate my findings using other housing cycles or indexes.

Statements

Data availability statements: The analysed dataset in the current study is publicly available and can be accessed at https://www.cnefdata.org/
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During the preparation of this work, the author used Grammarly to appear more native for English speakers. After using this tool/service, the author reviewed and edited the content as needed and took full responsibility for the publication's content.

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