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Rent Divergence within Neighborhoods

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Housing rents are a large share of household budgets and make a large contribution to overall inflation. Rent inflation rates for different types of housing units sometimes diverge, even in the same neighborhoods. We estimate during 2013 to 2016 apartment rents outpaced rents for detached housing in the United States by 0.76 percentage points annually after controlling for location effects. These rent dynamics imply a segmented housing market. They also suggest rent indexes need to be based on data structurally representative of their measurement objective. In particular, indexes based on professionally managed apartment complexes mismeasure the rents for housing generally. Even indexes based on careful geographical sampling, such as the Consumer Price Index's Owners' Equivalent Rent component, may be biased by using an unrepresentative mix of apartments and houses.

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1 Introduction

Housing rents have a huge expenditure weight in household budgets – particularly for low-income renters or renters in cities with growing labor markets – so rent inflation often raises affordability concerns. Rent inflation is a large component of overall inflation, underscoring the importance of its measurement. Rent inflation for different types of housing units sometimes diverge.

In the United States, apartment rents outpaced detached house rents throughout the 2010s (see Section 2.2). If the influence of location entirely explains the difference, then a sample of units of any type in a representative mix of neighborhoods can provide the basis for an accurate rent index. But if rents diverge by more than the influence of location can explain, then an unrepresentative mix of housing types (even in a random mix of neighborhoods) gives rise to an incorrect measurement of rent dynamics.

The rental unit microdata from the US Bureau of Labor Statistics (BLS) contain enough information to allow estimation of location and structure type effects. The BLS tracks rents for about 40,000 units for its rent and owners' equivalent rent (OER)[1] indexes. The survey randomly selects small neighborhoods from within a city, then samples a half-dozen rental units from each selected neighborhood (irrespective of structure type or management structure). This procedure ensures that the sample contains units from all parts of the rental market and contains some competing rental units within every selected neighborhood. These sample-randomizing features contrast with the procedures underlying other rental data sources, which often omit significant portions of the rental market or which contain no location data.

We find that, controlling for location and for other observable characteristics, rent dynamics differ in a statistically and economically significant manner across structure types, over long periods. For instance, between 2013 and 2016, after controlling for the effects of location, multiunit rent growth exceeded single-family detached rent growth by 0.76 percentage points annually.

The different price movements based on structure type imply market segmentation within rental housing. This might be expected from both demand and supply considerations. Burns (2015), Drew (2015), and Lerner (2016) document that tenants who seek apartment rentals differ from tenants who

[1]OER measures the value of housing services consumed in owner-occupied units. Movements in this implicit rent is imputed from price changes in nearby rental units.

seek single-unit homes in preferences and family situation. In 2013, 43 percent of renters of single-family detached units were families with children, compared to 27 percent of multifamily rentals. Young adults and high-income urban dwellers are less likely to want to live in older single-family detached suburban homes. Different preferences and family characteristics may give renters of single-family detached homes different outside options and different demand elasticities. On the supply side, there are differences as well. Most single detached homes are not professionally managed (though this has been changing). Detached-unit user costs differ from multifamily-unit user costs. For example, Coulson and Fisher (2015) and Halket, Nesheim and Oswald (2020) find that maintenance costs are systematically different and the land-unit ratio differs as well. Apartment complexes have economies of scope and scale, and different management structures can lead to different bargaining strategies and outcomes (Gallin and Verbrugge, 2019). Detached units can easily move into and out of the rental market. Supply changes (such as the surge in supply of single-family detached rentals since 2006) could well result in differential rent dynamics.

However, these considerations do not necessarily imply differential rent dynamics across structure type. Even cost differentials need not map into rent differentials, since rents do not seem to be that closely related to costs; user costs and rents can diverge markedly over extended periods. (Verbrugge, 2008; Braga and Lerman, 2019) Location has been well-established as the chief determinant of rent growth (see, e.g. Verbrugge, Dorfman, Johnson, Marsh, Poole and Shoemaker, 2017). What matters for pricing is the marginal renter, not average characteristics or demand elasticities. So long as some pool of renters views apartments and nearby rentable detached homes as substitutes, market forces might be expected to ensure a close relationship between the rent dynamics of different structures.

We show that despite the importance of location for determining rent growth, structure type is also an economically and statistically significant driver of rent growth.

Thus, sample representativeness is an utmost concern when measuring housing pricing movements. Our findings challenge the conclusions of several prominent studies that have criticized CPI shelter inflation measurement on the basis of rent inflation estimated using different data sources. Accurate shelter inflation measurement requires a rental housing sample that is geographically and structurally representative. In the United States, several new rental data sources are based on rents of professionally managed apartment complexes. They would need to be supplemented (and not merely reweighted) to become representative of the whole rental market. Alone they are of limited use for drawing implications about the accuracy of BLS rent indexes (contra Ambrose, Coulson and Yoshida, 2015; Ambrose, Coulson and Yoshida, 2018; Nothaft, 2018).

Second, our findings have implications for understanding how living standards have changed across income groups. For most in the bottom quintile of the income distribution, housing expenditures take more than 40 percent of income (OECD, 2019), so accurate accounting for housing costs is critical. Comparing income growth to average rent growth can give rise to misleading conclusions, since both location and structure type vary systematically with other socioeconomic indicators. For instance, our results indicate that over the past decade, shelter cost inflation has been overestimated for house-dwellers, in turn leading to an underestimation of growth in their living standards.

Third, our findings enhance our understanding of rental market dynamics. Despite differences in demand and supply influences across structure types, what matters for pricing is the marginal renter. Differential rent dynamics across structure types (after controlling for location) implies important market segmentation.

Finally, our findings have important implications for inflation measurement: OER inflation may have been notably mismeasured over the period of our study; section 4 estimates OER inflation to have been overstated by 0.34 percentage points between 2013 and 2016, for instance. Quantitatively, a deviation of this magnitude from the measurement goal is large enough to shift the headline CPI by almost 0.1 percentage point, of larger estimated magnitude than lower-level substitution bias and as large as new outlets bias (Moulton, 2018). Mismeasurement arises because our findings imply that the BLS rental sample is not representative for homeowners. Most homeowners live in detached houses, so we consider the change in the value of the implicit flow of rental services from owned housing is better proxied by the rent changes of nearby detached rental units.[2] But the BLS rental housing sample

[2]The service flow that houses yield to owners might diverge in important ways from the rent commanded by superficially similar houses, for at least two reasons. First, the findings of Halket et al. (2020) imply that owned houses have higher unobserved quality features that are more delicate (such as rose gardens or hot tubs), features that might deteriorate rapidly under the tenure of a renter. (Heston and Nakamura (2009) and Aten (2018) both provide evidence that contract rents understate the flow of rental services to

over the course of our study (and currently) is representative of the rental housing stock, not of the owned housing stock: for the latter, the percentage of detached units in the sample is too low. This implies that apartment rents receive too large a weight in the OER index, compared to detached units. Because apartment rents over our study rose more rapidly than detached unit rents, we argue OER inflation was overestimated. Section 4 also discusses how rent indexes could account for segmentation by structure type.

2 Data

2.1 Data Source

The BLS's Consumer Price Index Housing Survey asks the owners, property managers, and renters the rent charged for approximately 40,000 housing units in the United States. Each unit is surveyed every six months to create a panel data set of rents.

The BLS selects its sample by first selecting approximately 80 areas to be representative of all urban areas in the United States.[3] Each area is divided into contiguous regions labeled "strata," and Census block groups[4] are randomly selected from each stratum, using probability-proportional-to-size

the typical homeowner; Aten and Heston (2020) suggest a data-based method to estimate a premium to rental-equivalence estimates of OER for use in the national accounts.) Second, it may be argued that since most detached homes are not professionally managed, this might lead to mispricing. Detached homes feature far stickier rents than do apartments, and management structure may well influence rent dynamics (see Verbrugge and Gallin, 2017; Gallin and Verbrugge, 2019). However, regarding OER, the measurement goal is essentially this: how did the answer to the question "What would your home rent for?" change over the past six months. (In other words, what is the change in the market value of the flow of services your house provided over this period?) We find it difficult to believe that the rent movements in nearby apartments more closely proxy this unobserved change than do the rent movements in nearby detached homes. In other contexts, differences in observables raise questions about comparability (see the vast literature on causal inference, for example Athey and Imbens (2017)). For exactly this reason, between 1987 and 1998, BLS sampling procedures specified that particular owner units were matched to particular rental units with similar structural attributes.

[3]More precisely, the areas are what the BLS terms primary statistical units (PSUs). They are core based statistical areas following Office of Management and Budget definitions, except with less frequent revision.

[4]The block groups are labelled "segments" in BLS databases. In rare instances, a segment is an amalgamation of neighboring block groups.

procedures. Housing units are randomly selected from a list of probable rental units. (See Ptacek and Baskin (1996) for details.) Housing characteristics (including age, type of structure, exact location, number of bedrooms, and what utilities are included in rental payment) are recorded along with rents. The housing survey thus includes apartments and single-family homes, individually managed and corporately managed units, suburban and urban units. It is the most representative and diverse panel of rental housing available for the United States.

In much of what follows, we use three-year periods. We do so for two reasons. First, some structure types – particularly large apartment complexes – have more flexible rents than detached units; Genesove (2003) and Gallin and Verbrugge (2019) document that both tenant turnover and rent changes upon lease renewal vary notably by structure type. Hence, rents in large apartments will respond more rapidly to market developments; thus, a differential might simply reflect speed of response, rather than a truly different underlying inflation rate. After three years, however, most units will have experienced a rent change, mitigating this responsiveness differential. Second, a three-year differential will be unambiguously important – users of the Consumer Price Index are certain to see a differential over such a lengthy period as essential to correct. Unfortunately, using three-year periods has notable implications for sample sizes. Because of panel rotation and frequent non-response, only around a third of units in any period also have a rent quote three years later. We often highlight the second half of 2013 to the second half of 2016, because it is last three-year period before an acceleration in panel rotation.

The Consumer Price Index uses its housing survey to compute a rent index and an owners' equivalent of rent index. The rent measures used in rent and OER index construction are not the tenant- or landlord-reported "sticker price" or nominal rents, but instead receive various adjustments necessary for index accuracy. For instance, units age over time, and the BLS corrects for this using an "aging-bias" correction; reported rents often depart from the true market rent of the units, because tenant receive rent discounts in exchange for services rendered to a landlord.[5] One such adjustment applies only to OER rents. OER is a price-of-shelter concept that does not include

[5]The rents entering the index may receive other adjustments. An important case is vacancy: rents that are missing owing to vacancy are imputed. We do not include any imputed rents from this study.

utilities, since utilities are measurable out-of-pocket expenses for homeowners. OER index movements are based upon inflation in market rents; but since these rents often include utilities – and utilities costs often greatly exceed 10% of the rent – the BLS must estimate the utilities part of each rent, and remove it, before using this rent in constructing OER (see Verbrugge 2012). The resultant (post-utilities-adjusted) rent measure is termed "economic rent." However, in our tables and results, we use nominal rents (except where otherwise noted) for comparability with previous studies.[6] The two indexes also have different weights, which change monthly based on response rates and rent movements. This article's tables and regressions equally weight observations (unless otherwise noted). We drop observations which record rent as $0 or $1.[7]

2.2 Data Patterns

Recently, rents in multiunit buildings have increased faster than rents for single-family detached houses. Table 1 shows that a divergence in rents occurred in every seminannum for a decade, though by varying amounts. Table 2 demonstrates that this pattern is present widely, across geography, rental unit size, and rent levels.

3 Regression Analysis

3.1 Location Indicator Regressions

Regressions can help separate structure type effects from neighborhood effects. Rent quotes were collected in both the second half of 2013 and the second half of 2016 for 3,390 single-family detached homes, 7,005 condominiums or apartments in multiunit buildings, 2,166 single-family attached homes, and 320 mobile homes or units in other structure types. Table 3 presents coefficients from a regression where the annualized percentage rent

[6]See Verbrugge and Poole (2010) for a study detailing the importance of these weights and other differences between the rent and OER indexes. Our main results do not hinge on the particular rent measure used.

[7]Such observations are not uncommon, but typically reflect a rent discount offered to certain tenants in exchange for services provided to the landlord. The BLS data do not contain public housing units.

Table 1: Average rent inflation by structure type

Interval	Single detached	Single attached	Multiunit
2010h1 – 2010h2	0.55	0.57	1.01
2010h2 – 2011h1	0.35	0.96	0.89
2011h1 – 2011h2	0.88	0.91	1.78
2011h2 – 2012h1	0.56	0.82	1.20
2012h1 – 2012h2	0.84	1.02	1.65
2012h2 – 2013h1	0.73	0.87	1.22
2013h1 – 2013h2	0.97	1.06	1.75
2013h2 – 2014h1	0.73	1.08	1.29
2014h1 – 2014h2	1.01	0.90	1.66
2014h2 – 2015h1	0.86	1.15	1.45
2015h1 – 2015h2	0.97	1.23	1.88
2015h2 – 2016h1	1.14	1.12	1.51
2016h1 – 2016h2	1.26	1.40	1.89
2016h2 – 2017h1	0.99	1.10	1.59
2017h1 – 2017h2	1.30	1.58	1.99
2017h2 – 2018h1	1.19	1.32	1.52
2018h1 – 2018h2	1.48	1.65	1.83
2018h2 – 2019h1	1.46	1.43	1.50
2019h1 – 2019h2	1.35	1.53	1.97
2019h2 – 2020h1	0.76	1.10	1.47
2020h1 – 2020h2	2.14	0.79	1.17

Percentages are equally-weighted arithmetic averages of the increase in nominal rents for units in the CPI Housing Survey with quotes for both beginning and end periods. Percentages are not annualized.

Table 2: Average rent inflation, first half 2017 - first half 2020

	Single detached	Single attached	Multiunit
Overall	2.74	3.17	3.37
	(0.10)	(0.14)	(0.06)
County population density			
$< 200/mi^2$	1.66	2.00	2.12
	(0.26)	(0.30)	(0.17)
200 to 5,000/mi^2	3.03	3.42	3.58
	(0.11)	(0.15)	(0.07)
5,000 to 20,000/mi^2	1.84	1.90	3.36
	(0.36)	(0.91)	(0.46)
$> 20,000/mi^2$	2.27	1.93	2.96
	(0.52)	(0.88)	(0.32)
In a state with rent control			
No	2.73	3.00	3.30
	(0.12)	(0.14)	(0.07)
Yes	2.79	4.11	3.62
	(0.23)	(0.38)	(0.15)
Numbers of bedrooms			
2 bedrooms	2.73	3.06	3.29
	(0.16)	(0.18)	(0.09)
3 bedrooms	2.62	3.41	2.75
	(0.21)	(0.30)	(0.23)
4 or more bedrooms	2.36	2.61	4.56
	(0.29)	(0.68)	(2.81)
Initial rents			
\$1 to \$649	3.48	3.36	3.69
	(0.39)	(0.32)	(0.18)
\$650 to \$899	2.87	3.45	3.79
	(0.14)	(0.18)	(0.09)
\$900 to \$1299	3.08	2.88	3.38
	(0.17)	(0.27)	(0.10)
\$1300 or more	1.92	2.53	2.39
	(0.16)	(0.25)	(0.12)
Observations	2,869	1,173	3,928

Percentages are annualized, equally weighted arithmetic averages of the increase in nominal rents for units in the CPI Housing Survey with quotes for both Jan-Jun 2017 and Jan-Jun 2020. Standard errors in parenthesis.

Table 3: Regression of annualized percent rent change, second half 2013 - second half 2016

	(1)	(2)	(3)	(4)	(5)
Intercept	1.91***	2.00***	1.06	0.78	0.51
	(0.08)	(0.32)	(0.72)	(2.28)	(2.26)
Single-family detached	-	-	-	-	-
Single-family attached	0.38***	0.21	0.61***	0.35*	0.29
	(0.12)	(0.13)	(0.13)	(0.18)	(0.19)
Multiunit	1.10***	0.78***	1.24***	0.76***	0.59***
	(0.09)	(0.12)	(0.10)	(0.17)	(0.19)
Mobile home and other	-0.13	-0.16	0.30	0.54	-0.92
	(0.27)	(0.99)	(0.28)	(0.43)	(1.32)
Built before 1990		-			-
Built after 1990		0.39***			0.32*
		(0.11)			(0.20)
Studio or 0-bedroom unit		-			-
1-bedroom unit		0.42			0.55
		(0.31)			(0.37)
2-bedroom unit		0.02			0.24
		(0.31)			(0.37)
3-bedroom unit		-0.1744			0.0532
		(0.32)			(0.39)
4-bedroom unit		-0.36			-0.28
		(0.36)			(0.44)
5- or more bedroom unit		-0.46			-0.43
		(0.50)			(0.63)
Location indicators	none	none	county	block group	block group
Observations	12,881	12,389	12,881	12,881	12,389
R^2	0.012	0.015	0.126	0.512	0.522

Standard errors in parenthesis. Significant at *10%, **5%, ***1%.

Table 4: Regression of annualized rent change, first half 2017 - first half 2020

	(1)	(2)	(3)	(4)	(5)
Intercept	2.74***	3.47***	2.85***	6.96***	6.52***
	(0.10)	(0.37)	(0.82)	(1.95)	(1.94)
Single-family detached	-	-	-	-	-
Single-family attached	0.43**	0.39**	0.32*	0.34	0.20
	(0.17)	(0.17)	(0.17)	(0.27)	(0.27)
Multiunit	0.63***	0.45***	0.53***	0.60***	0.31
	(0.12)	(0.13)	(0.12)	(0.22)	(0.23)
Mobile home and other	0.47	3.82***	0.52	-0.68	0.49
	(0.43)	(0.96)	(0.44)	(0.70)	(1.98)
Built before 1990		-			-
Built after 1990		-0.09			0.13
		(0.13)			(0.24)
Studio or 0-bedroom unit		-			-
1-bedroom unit		-0.39			0.65
		(0.36)			(0.43)
2-bedroom unit		-0.65*			0.39
		(0.36)			(0.43)
3-bedroom unit		-0.85**			-0.02
		(0.38)			(0.46)
4-bedroom unit		-0.98**			-0.60
		(0.46)			(0.58)
5- or more bedroom unit		-1.79**			1.63
		(0.83)			(1.15)
Location indicators	none	none	county	block group	block group
Observations	8,241	7,929	8,241	8,241	7,929
R^2	0.004	0.007	0.126	0.538	0.553

Standard errors in parenthesis. Significant at *10%, **5%, ***1%.

growth is regressed onto indicator variables for structure type and location. Thus, the regression model is

$$r_i = a_{j(i)} + b_{\ell(i)} + \beta X_i + \epsilon(i), \tag{1}$$

where

$$r_i = 100\left[(rent_{(i,2016)}/rent_{i,2013})^{1/3} - 1\right], \tag{2}$$

i indexes rental units, $j(i)$ indicates the structure type of unit i, $\ell(i)$ indicates its location, a is the fixed effect for structure type $j(i)$, b is fixed effect for location $\ell(i)$, X_i is vector of other characteristics of unit i, and ϵ_i is an idiosyncratic econometric error term. Column 1 of Table 3 presents indicator coefficients in a specification without any location controls. An ordinary least squares regression with one set of category indicators is equivalent to taking averages, so the coefficients in Column 1 are the differential average rent changes by structure type. Nominal rents increased by 1.91 percentage points for detached single-family homes and 1.91+1.10=3.01 percent for apartments in multiunit buildings. The 1.10 percentage point differential is both economically and statistically significant.

Columns 3 and 4 of Table 3 report specifications that add location indicators. If location drove the entire structure-type rent divergence, then coefficients on structure type indicators would become statistically insignificant as location controls are added. But as the indicator variables for location represent finer and finer geography, the coefficients on structure type often remain significantly different from zero and so reveal the significant influence of structure type. During this period, rent growth for multiunit exceeded that of single-family detached housing by 0.76 percentage points (annualized) even after for controlling for location with block group indicators (Column 4, Table 3).

The divergence in rents between multiunit and single-family housing units is not explained by the units being in different neighborhoods, nor is it fully explained by the units being of different sizes or ages. Columns 2 and 5 of Table 3 report regressions that add indicators for the number of bedrooms or recent construction as explanatory variables. (Some housing units had unknown construction dates, so these columns are run on a subset of the data used in the other columns.) Most of the room count indicators have statistically insignificant coefficients. The coefficient of multiunit remains statistically and economically significant.

Table 4 reports the same regression specifications on data from the next triannum, the first half of 2017 to the first half of 2020.[8] As before, the coefficient on the multiunit indicator remains statistically significant even after adding location controls. Multiunit rent increased 0.60 percentage points (annualized) more than single-family units after accounting for block group variation in rent changes. The estimated structure type effect is reduced by the inclusion of more unit characteristics in the regression in Columns 2 and 5. Smaller units appear to have greater rent increases in this period, and apartments are disproportionally smaller units. Yet, the structure type remains associated with large and statistically significant differences in rent appreciation.

3.2 Robustness

The rent survey has outliers. The standard deviation of rent change between the second halves of 2013 and 2016 was 16.7 percent (not annualized). An eight-fold rent increase and 95 percent rent reduction are the most extreme movements observed. Of the 12,881 observations, 452 had rent changes more than 2 standard deviations from the average. Yet, dropping these 452 outliers had no significant effect on regression results, as reported in Table 5. Column 1 copies column 4 of Table 3. Column 2 repeats the same regression, only dropping the 452 outliers. The structure type coefficients are still the difference in rent growth by structure type after controlling for location with block group indicators. The estimated difference between single-family detached and multiunit rent increases is 0.91 annualized percentage points (instead of 0.76 percentage points) when outliers are excluded.

All regressions reported thus far use the nominal rent that respondents report. But the rent divergence by structure type is also seen in the rent measures that enter the rent index ("economic rent," which corrects for subsidies and work reductions, and includes adjustments for aging and other quality adjustments), and that enter the OER index ("pure rent," which adjusts economic rent by removing the utilities portion of the rent; see Verbrugge (2012)). The different inflation rates are not driven by to the presence of utilities or other adjustments differing by structure type. Column 3 repeats the regression of Column 1, except using economic rent change instead of nominal rent change as the dependent variable. Column 4 uses pure rent

[8] Table 2 uses the same data, minus the units categorized as Mobile homes and other.

Table 5: Regression of annualized percent rent change, second half 2013 - second half 2016

	(1)	no outliers (2)	economic rent (3)	pure rent (4)
Intercept	0.78	0.74	0.43	0.42
	(2.28)	(1.54)	(2.34)	(2.57)
Single-family detached	-	-	-	
Single-family attached	0.35*	0.31**	0.36*	0.38*
	(0.18)	(0.13)	(0.19)	(0.20)
Multiunit	0.76***	0.91***	0.81***	0.93***
	(0.17)	(0.12)	(0.18)	(0.19)
Mobile home and other	0.54	0.19	0.97**	1.15**
	(0.43)	(0.30)	(0.44)	(0.48)
Observations	12,881	12,429	12,881	12,881
R^2	0.512	0.543	0.516	0.506

Standard errors in parenthesis. Significant at *10%, **5%, ***1%.

rate change as the dependent variable. The coefficient on multiunit indicates a 0.76 to 0.93 percentage point difference from single-family detached rents, depending on the rent measure.

4 Implications for Housing Indexes

The divergence in rents by structure type would not be a problem for an index calculated from a sample that was both geographically and structurally representative for the index in question. The CPI sample and its mix of structure types is fairly representative of urban rental housing, so differences in rent inflation between housing types cause no major bias in the CPI rent index. However, the CPI's OER calculations use fewer detached houses and more multiunit buildings than would be representative of owner-occupied housing. This is not a problem unique to the BLS sample. Most other rent indexes and rent datasets incorporate an even smaller proportion of single detached homes, and none is both geographically and structurally representative.

What is the implication of the not-fully-representative sample for OER inflation? To estimate this, suppose, as in the regression reported in Table 3, rent growth in expectation (denoted g_i for unit i) is the sum of a structure-type effect (a_h for structure type h) and a neighborhood effect (b_ℓ for location ℓ). Thus, $g_i = a_j + b_\ell$. Expected OER growth in location ℓ (G_ℓ) is the sum of rent growth for each structure type, weighted by the structure type's share of OER in that location (denoted $s_{j,\ell}$ for structure type j in location ℓ): $G_\ell = \sum_{i \in \ell} g_i = \sum_j s_{j,\ell}(a_j + b_\ell)$. Overall rent inflation (G) is weighted sum over locations, where w_ℓ is the weight for location ℓ:

$$G = \sum_\ell w_\ell \cdot G_\ell \tag{3}$$

$$G = \sum_\ell \sum_j w_\ell s_{j,\ell}(a_j + b_\ell) \tag{4}$$

$$G = \sum_j a_j \sum_\ell w_\ell s_{j,\ell} + \sum_\ell w_\ell b_\ell \underbrace{\sum_j s_{j,\ell}}_{=1} \tag{5}$$

$$G = \sum_j a_j \sum_\ell w_\ell s_{j,\ell} + \sum_\ell w_\ell b_\ell \tag{6}$$

Let $W_j = w_\ell s_{j,\ell}$. This weight is mismeasured; let $\tilde{W}_j$ denote the incorrect value of W_j used. Then, the resulting measurement error for rent growth is $\sum_j a_j(W_j - \tilde{W}_j)$. The regression coefficients for structure type indicators in Table 3 give an estimate for a_j over the period from the second half of 2013 to the second half of 2016. The structure-type weight W_j should be the share of owner-occupied housing services produced by that housing type j. For calculations here, the implied expenditures by housing type from the Consumer Expenditure Surveys (CE) will be assumed to measure W_j accurately. The shares measured by CE differ greatly from the OER weights. CE estimates 86.6 percent of owner-occupied rental equivalence came from single-family detached homes in the second half of 2016. (Similarly, the Census's American Community Survey estimated 82.6 percent of owner-occupied housing units are single unit detached.) However, single-family detached housing had only 33.6 percent of the weight in the CPI's OER calculations. Multiunit housing accounted for 5.2 percent of owner-occupied housing services in the CE data (combining the building type categories of 3-plex or 4-plex, garden, high-rise, and apartment or flat) but represented 44.6 percent of OER

Table 6: Mismatch of structure types in OER weights

	Rental equivalence share (Jul-Dec 2016 CE %) (1)	Share of OER weight (Jul-Dec 2016, %) (2)	Structure effect estimate (3)	Estimated contribution to OER mismeasurement (%) (4)
Single detached	86.6	33.6	0	0
Single attached	5.6	18.4	0.35	0.04
Multiunit	5.2	44.6	0.76	0.30
Other	2.6	3.4	0.54	0.00
Total	100	100		0.34

weight (combining the structure type categories multiunit with elevator and multiunit without elevator).

Table 6 presents our calculations. The difference between columns 1 and 2 gives an estimate of $W_j - \tilde{W}_j$. Column 4 is that difference, multiplied by a_j from Column 3. The under-weighting of single detached units and the over-weighting of all other housing resulted in an overestimate of OER inflation by 0.34 percent annually from 2013 to 2016.

OER has a relative importance in the CPI of 0.23, so sampling that accounts for structure type effects would have decreased the all items CPI by $0.34 \times 0.23 = 0.08$ percentage points annually. To give a sense of its significance, this is roughly the same magnitude as the aging bias adjustment in the CPI shelter indexes, universally thought to be far too large to ignore. (Randolph, 1988; Gallin and Verbrugge, 2007) It is bigger than Moulton (2018) estimates for the CPI bias from lower-level substitution, as big as the bias from new outlets, and a quarter of the size as from new products and quality change, all price index measurement issues to which great attention is given.

Avoiding this bias in an OER index is challenging. A thorough comparison of the options is beyond the scope of this study. Instead, we simply note three of the leading possibilities (that are not mutually exclusive). Each has notable drawbacks. First, a sample that is both geographically and structurally representative for owner-occupied housing would be ideal for an OER index, but such a sample would be expensive to construct. Rented houses can be hard to identify. Houses might enter and exit the rental market more frequently, necessitating expense to replace units as they drop out of the sample. A stratified sampling procedure that maintains enough houses in each locality is still of unknown feasibility.

Second, data sources with all types of structures, such as the BLS's, allow reweighting units to achieve a balance of structure types. However, this depends upon the sample at hand. The sample for some areas may not even have a full complement of structure types. In the BLS's data, a given block group has at most six rental units in the sample (before sample attrition reduces the number). Suppose only one detached unit is in the sample for a block group; given the owned structure types in the neighborhood, this single unit might need to receive 90 percent of the weight for the block group. Variance of the index could markedly increase.

Third, rent indexes could be constructed by combining separate geographically representative indexes for each structure type. Indeed, these structure-type indexes could even make use of different data sources. An index with only apartments or only houses by itself would give an incomplete and potentially misleading reading on rents (or owners' equivalent rents) generally.

5 Conclusion

Despite the importance of location for determining rent growth, structure type is also an economically and statistically significant driver of rent growth. There are several important implications. First, the importance of structure types implies that alternative rental data sources are of limited use for drawing implications about the accuracy of BLS rent indexes. Second, it implies that understanding changes in housing costs facing various income groups must rely upon data that are able to take into account both the locational and structural characteristics of this population's housing. Third, it is an important step forward for our understanding of rental market dynamics, since differential rent dynamics across structure types implies important market segmentation. Finally, it implies that there is a measurement problem in the CPI's OER index.

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