THE
QUARTERLY JOURNAL
OF ECONOMICS

Vol. 130 February 2015 Issue 1

VERY LONG-RUN DISCOUNT RATES*

Stefano Giglio
Matteo Maggiori
Johannes Stroebel

We estimate how households trade off immediate costs and uncertain future benefits that occur in the very long run, 100 or more years away. We exploit a unique feature of housing markets in the United Kingdom and Singapore, where residential property ownership takes the form of either leaseholds or freeholds. Leaseholds are temporary, prepaid, and tradable ownership contracts with maturities between 99 and 999 years, while freeholds are perpetual ownership contracts. The price difference between leaseholds and freeholds reflects the present value of perpetual rental income starting at leasehold

*We thank Sumit Agarwal, Robert Barro, Arthur van Benthem, Jules van Binsbergen, Andrea Caggese, John Campbell, Jaime Casassus, Raj Chetty, John Cochrane, Xavier Gabaix, Bob Goldstein, Christian Gollier, Bob Hall, Barney Hartman-Glaeser, Anil Kashyap, Amir Kermani, Mervyn King, Ralph Koijen, Martin Lettau, Lars Lochstoer, Hanno Lustig, Tim McQuade, Stijn van Nieuwerburgh, William Nordhaus, Monika Piazzesi, Tomasz Piskorski, Tarun Ramadorai, Esteban Rossi-Hansberg, Florian Schulz, Amit Seru, Andrei Shleifer, Nick Stern, Martin Weitzman, four anonymous referees, as well as seminar participants at the NBER (Public Economics, Asset Pricing, Financing Housing Capital, Real Estate), Harvard (HBS and Department), Princeton (JRC Conference), Chicago Booth, NYU Stern, Columbia GSB, Wharton, MIT Sloan, the University of Minnesota, Federal Reserve Board, Adam Smith Conference (LBS/LSE/Oxford), UBC Winter Finance Conference, ASU Sonoran Winter Finance Conference, HULM Conference, Universitaet Regensburg, Fordham University, Notre Dame, and the ASSA 2015 meetings for helpful discussions and comments. We gratefully acknowledge the generous research support from the NYU Stern Center for the Global Economy and Business as well as from the Fama-Miller Center and the Initiative on Global Markets at the University of Chicago Booth School of Business, and from the Harvard Weatherhead Center for International Affairs. We thank iProperty and Rightmove for sharing part of their data. We thank Miguel de Faria e Castro for excellent research assistance.

The Quarterly Journal of Economics (2015), 1–53. doi:10.1093/qje/qju036.
Advance Access publication on November 25, 2014.

expiration, and is thus informative about very long-run discount rates. We estimate the price discounts for varying leasehold maturities compared to freeholds and extremely long-run leaseholds via hedonic regressions using proprietary data sets of the universe of transactions in each country. Households discount very long-run cash flows at low rates, assigning high present value to cash flows hundreds of years in the future. For example, 100-year leaseholds are valued at more than 10% less than otherwise identical freeholds, implying discount rates below 2.6% for 100-year claims. *JEL Codes:* G11, G12, R30.

I. Introduction

Long-run discount rates play a central role in economics and public policy. For example, much of the debate around the optimal response to climate change centers on the trade-off between the immediate costs and the very long-term benefits of policies that aim to reduce global warming. Similar cost-benefit analyses are required of all U.S. government agencies prior to proposing and adopting regulation. Unfortunately, there is little direct empirical evidence on how households discount payments over very long horizons, because of the scarcity of finite, long-maturity assets necessary to estimate households' valuation of very long-run claims. For regulatory action with "intergenerational benefits or costs," the U.S. Office of Management and Budget (2003) therefore recommends a wide range of discount rates (1–7%), lamenting that while "private markets provide a reliable reference for determining how society values time within a generation, for extremely long time periods no comparable private rates exist."

We provide direct estimates of households' discount rates for payments very far in the future. We exploit a unique feature of residential housing markets in the United Kingdom and Singapore, where property ownership takes the form of either very long-term leaseholds or freeholds. Leaseholds are temporary, prepaid, and tradable ownership contracts with maturities ranging from 99 to 999 years, and freeholds are perpetual ownership contracts. The price difference between leaseholds and freeholds for otherwise identical properties captures the present value of perpetual rental income starting at leasehold expiration, and is thus informative about households' discount rates over that horizon.

Our empirical analysis is based on proprietary information on the universe of residential property sales in the United Kingdom (2004–2013) and Singapore (1995–2013). These data contain information on transaction prices, leasehold terms, and property characteristics such as location and structural attributes. We estimate long-run discount rates by comparing the prices of leaseholds with different maturities to each other and to the price of freeholds across otherwise identical properties. We use hedonic regression techniques to control for possible heterogeneity between leasehold and freehold properties. This allows us to identify price discounts associated with differences in lease length. We find that 100-year leaseholds are valued 10–15% less than otherwise identical freeholds; leaseholds with maturities of 125–150 years are valued 5–8% less than freeholds. There are no price differences between leaseholds with maturities of more than 700 years and freeholds. Our results show that households apply annual discount rates of below 2.6% to housing cash flows more than 100 years in the future.

While these housing markets provide a unique and important setting for understanding very long-run discount rates, they are not frictionless markets. We therefore address a number of possible concerns that the observed price differences between leaseholds of different maturity and freeholds might be driven not only by the different maturity of the claims, but also by other differences between the two contracts or frictions specific to housing markets. We first show that the empirical results are consistent across the United Kingdom and Singapore, two housing markets with otherwise very different institutional settings. In addition, we provide direct evidence that the leasehold discounts are not related to either systematic unobserved structural heterogeneity across different properties, differences in the liquidity of the properties, or a different clientele for the different ownership structures, and are unlikely to be explained by contractual restrictions in leasehold contracts.

We first address the concern that our estimated price discounts are driven by systematic differences in unobserved property characteristics. To do this, we analyze how annual rents differ across freehold and leasehold properties with different maturities. Conditional on observable property characteristics, rental prices of leasehold properties with different maturities and freeholds are essentially identical. This suggests that the observed transaction price differences are not related to differences in the flow utility from these properties but to the maturity of the contracts. Consistent with this, the absence of a price difference between leaseholds with 700 or more years remaining and freeholds provides further confidence that our hedonic control variables allow us to partial out all important sources of structural heterogeneity between leasehold and freehold properties.

We next consider whether covenants in leasehold contracts might explain the estimated price discounts. Since leaseholds with more than 700 years' maturity trade at the same price as freeholds, the price discounts between shorter leaseholds and freeholds are unlikely to be explained by fundamental differences in the general contract structure (e.g., a significant control premium or a maturity-independent preference of households for freehold ownership), since these should show up along the entire term structure, including the very long leases. This is consistent with a robustness check that estimates discounts only within leaseholds of different maturity, and finds them to be similarly large. We also show that our estimated price discounts persist even after controlling for the initial lease length of contracts, addressing concerns that restrictive covenants might be more prevalent for shorter maturity leaseholds. The absence of such a correlation is confirmed by a direct analysis of the covenants in a subset of leasehold contracts. Finally, since most important restrictive covenants on short-maturity leaseholds would have to be passed on to possible renters of the property, the absence of differential rental prices for leaseholds of different maturity further confirms the absence of important restrictive covenants that differ by lease length.

We then address concerns that differences in the behavior of residual freeholders, who hold the rights to the property after the expiration of the lease, might affect a leaseholder's incentives and ability to extend an existing leasehold. This could generate an endogenous correlation between remaining lease length and freeholder characteristics and behavior, since leaseholds with less attractive freeholders may be extended less frequently and could therefore sell both for less and with fewer years remaining on the lease. To consider whether this can explain our estimates, we further homogenize our estimation sample by only exploiting differences in the remaining lease length of flats in the same building, all of which have the same freeholder. Our estimates are unchanged in that sample, suggesting that the price differences across leaseholds of differential remaining maturity are not related to systematic differences in freeholder characteristics. These estimates also further address some of the other concerns already discussed. First, flats in the same building are even less likely to differ systematically on unobservable property characteristics. Second, covenants, contracts for maintenance and servicing, and restrictions on property redevelopment generally do not vary within a building. Therefore, our estimates from this sample are robust to concerns that the price differences between leaseholds of different maturity are driven by correlation of remaining lease length with any of these factors.

We also document that price differences are not driven by differential liquidity of leasehold contracts with different maturities and freehold contracts, by showing that the time on market does not vary systematically across the term structure of remaining lease length. We then consider whether the presence of a different clientele for leasehold and freehold properties can explain the price differences, but find evidence that buyers of these contracts are essentially identical on observable characteristics. Our estimates also cannot be explained by potential financing frictions that might be important for short-maturity leasehold properties (50–70 years), since leasehold discounts remain substantial even for maturities of 200 years, for which the effects of potential financing frictions are too far away to matter quantitatively.

Overall, our findings show that a sizable fraction of the value of residential properties comes from cash flows that occur hundreds of years away. To interpret the economic magnitude of the observed leasehold discounts and implied discount rates, we consider the predictions from a simple valuation model with constant discount rates across maturities. In the simplest constant-discount-rate model, rental income D_t grows at rate g and is discounted at a constant rate r. The prices for the freehold P_t, and the T-maturity leasehold P_t^T are given by:

$$P_t = \frac{D_t}{r-g}; \qquad P_t^T = \frac{D_t}{r-g}(1 - e^{-(r-g)T}).$$

The first formula is the Gordon (1982) growth valuation for infinitely lived assets, the second formula corrects the freehold price for the shorter maturity of the leasehold to obtain the leasehold price. Both are derived from first principles in Section V. In this valuation model, the price discount between leaseholds and freeholds is:

$$Disc_t^T \equiv \frac{P_t^T}{P_t} - 1 = -e^{-(r-g)T}.$$

To match the observed discounts, this formula requires the net discount rate for housing cash flows, $r - g$, to be approximately

1.9%. Together with an estimate of long-run real rent growth of $g = 0.7\%$, this shows that households use low discount rates ($r = 2.6\%$) for very distant housing cash flows. Combining this with a separate estimate of the real long-run risk-free discount rate of 1% obtained from the real U.K. yield curve, we also find that housing cash flows have a relatively low long-run risk premium of 1.6%.

Our estimates of very long-run discount rates are of direct interest to a large theoretical literature that has lamented the absence of estimates for these very long maturities, not only for real estate but for any asset. This literature has often been motivated by the importance of long-run discount rates in particular applications: the analysis of climate change (Arrow et al. 1996; Weitzman 2001, 2013; Nordhaus 2007; Stern 2007; Gollier 2012; Barro 2013; Pindyck 2013; Farmer et al. 2014), the study of the term structure and long-run properties of risky assets (Alvarez and Jermann 2005; Binsbergen, Brandt, and Koijen 2012; Binsbergen et al. 2013), and macroeconomics and fiscal policy (Auerbach, Gokhale, and Kotlikoff 1994; Hall 2014). In the conclusions, we discuss the possible implications of our finding of low long-run discount rates for these fields.

II. Housing Markets in the United Kingdom and Singapore

In this section we discuss the relevant institutional details of housing markets in the United Kingdom and in Singapore, highlighting the distinguishing characteristics of freeholds and leaseholds.[1] Online Appendix A.1 and A.2 provide detailed additional information.

II.A. Leaseholds and Freeholds in the United Kingdom

Property contracts in England and Wales come in two forms: permanent ownership, called a freehold, and long-maturity, temporary ownership, called a leasehold. A leasehold is a grant of exclusive possession for a clearly defined, temporary period of

1. This contract structure is not unique to the United Kingdom and Singapore. Other papers in the real estate literature have studied the pricing of leasehold and freehold contracts in a variety of settings and countries (e.g., Capozza and Sick 1991; Wong et al. 2008; Iwata and Yamaga 2009; Tyvimaa, Gibler, and Zahirovic-Herbert 2014; Bracke, Pinchbeck, and Wyatt 2014; Gautier and van Vuuren 2014).

time (Burn, Cartwright, and Cheshire 2011). Common initial leasehold maturities are 99, 125, 150, 250, or 999 years. During this period, ownership of the leasehold entitles the lessee to similar rights as the ownership of the freehold, including the right to mortgage and rent out the property. Unlike for commercial leases, the vast majority of the costs associated with a residential leasehold come through the upfront purchase price; annual payments, the so-called ground rents, are small to nonexistent and do not significantly affect the prices paid for leaseholds. Leasehold properties are traded in liquid secondary markets, where the buyer purchases the remaining term of the lease.

Once the leasehold expires, the ownership reverts back to the freeholder. However, it is common for leaseholders to purchase lease extensions ahead of lease expiration. Over time, a number of laws have regulated the rights of leaseholders in the United Kingdom to extend their lease terms, and have codified the bargaining process between leaseholders and freeholders. For our sample period, the law states that leaseholders had the right to request a lease extension from the freeholder in exchange for paying a premium. The valuation on which the premium is based does not include the value of improvements to the property paid for by the leaseholder. If leaseholder and freeholder cannot agree on the premium, they can appeal to a government-run Leasehold Valuation Tribunal (LVT) with the power to set the prices for extensions. In Section IV.H, we discuss the effects of LVT decisions and lease extension regulation on the interpretation of our results and show that the particular institutional setting of the United Kingdom tends to balance the potentially stronger negotiating power of freeholders with laws and court decisions that might be favorable toward the leaseholder.

Some leaseholds contain covenants that might, for example, restrict the type of commercial activity that can be operated on the land. In Section IV.B we provide empirical evidence that differential covenants across contracts are unlikely to explain the observed differences in prices between leaseholds and freeholds, and between leaseholds of different maturity. Finally, management fees and service charges that are sometimes levied on leaseholders for the maintenance of the property primarily cover expenses also faced by freeholders, and do not significantly confound our analysis.

II.B. Leaseholds and Freeholds in Singapore

Residential properties in Singapore are also sold as freeholds or leaseholds, where the latter almost always have initial terms of 99 or 999 years. By far the largest freeholder is the government of Singapore, represented by the Singapore Land Authority (SLA). As in the United Kingdom, there is a vibrant private secondary market for leaseholds, where buyers purchase the remaining term of the original lease.

At the expiration of the lease, the ownership interest reverts to the SLA. Leaseholders may apply for a renewal of the lease before expiration. The granting of an extension is decided on a case-by-case basis; considerations include whether the development is in line with the government's planning intentions, and results in land use intensification or the mitigation of property decay. Between 2007 and 2010 about 60% of lease extension applications were approved. If the extension is approved, the chief valuer determines the "land premium" that will be charged. Prior to 2008 an additional "building premium" was charged, based on the value the chief valuer puts on the building sitting on the land with an expiring lease; the building premium was payable if a lease extension was sought and the building not demolished. The new lease cannot exceed the original and might be shorter if otherwise not in line with the Urban Redevelopment Authority's (URA) planning intention.

III. Empirical Analysis

The estimation of the relative prices of leaseholds and freeholds is potentially challenging because the underlying properties are heterogeneous assets. Since leasehold and freehold properties could differ on important dimensions such as property size and location, comparing prices across properties requires us to control for these differences. We use hedonic regression techniques (Rosen 1974), which allow us to consider the variation in price over time and across lease terms for different properties while controlling for key characteristics of each property.

III.A. U.K. Residential Housing Data

We begin by analyzing data from England and Wales. We obtained transaction-level administrative data on all residential housing sales between 1995 and 2013 from the Land Registry.

The data include the price paid as well as some characteristics of the house: whether the property is a flat (apartment) or a house, the type of house (detached, semi-detached, or terraced), the full address, and a "new construction" indicator. In addition, the Land Registry provided us with an indicator of whether the transaction was for a freehold or a leasehold property, as well as previously unexplored information on leasehold characteristics such as the origination date and lease length.

Two limitations affect the combined transaction-leasehold data. First, until the Land Registry Act 2002, which was enforced from October 2003, leasehold registration was not mandatory for all leases, and leases were valid and enforceable even if they were not registered with the Land Registry (see Online Appendix A.1.2). This limited the incentive to register a lease. Therefore, data before October 2003 are subject to measurement error because transactions that occurred before that date may have been erroneously associated with out-of-date leaseholds. For this reason, we focus our analysis on the period 2004–2013.

The second limitation is that the leasehold associated with each transaction in the Land Registry data set is the lease registered with the Land Registry at the time of the transaction. This is not a problem for our analysis except when the freeholder and leaseholder agree to a lease extension. A problem occurs if the lease extension happened before the transaction but is only registered afterward. We have manually detected a number of such instances in a subsample of leasehold transactions. In those cases, the data erroneously report the terms of the older (and shorter) lease, while the price paid pertains to the new (and longer) lease. This biases our analysis against finding a large price discount for short leases because a higher price (corresponding to a longer lease) would be mistakenly associated with a leasehold with fewer years remaining. When we can identify lease extensions (because we observe transactions that occur under both the old and the new lease), we observe that around 84% of extensions occur for leaseholds of less than 80 years remaining (see Online Appendix Figure A.1). We therefore focus on estimating price discounts for leaseholds with maturities above 80 years, where extensions are rare and which are particularly informative about very long-run discount rates.[2]

2. We also exclude the 3.1% of transactions for flats in properties for which we observe both a freehold and a leasehold transaction. This is because when the same

For 80% of the properties, we have obtained additional proprietary property characteristics, such as the number of bedrooms, bathrooms, the size and age of the property, and information on the condition, type of heating, and availability of parking. These are collected by Rightmove.co.uk from "for sale" listings and other data sources. Rightmove also provided information on the time on the market for most properties, as well as rental data for about 29,000 flats that were listed in London in 2011 and 2012, which allows us to compare rental prices across leasehold and freehold properties.

III.B. U.K. Data: Summary Statistics

Our final data set contains more than 8 million property transactions between 2004 and 2013. Table I provides an overview of the U.K. transaction sample. Houses in the United Kingdom are mainly transacted as freeholds, with some very long leases and very few shorter leases. Flats are mainly traded as leaseholds. Since the market for flats and the market for houses are relatively segmented and differ in their propensity to use freehold and leasehold contracts, we study them separately. In our main analysis we focus on flats, which have significant variation in terms of remaining lease length. Online Appendix A.3 shows that the estimated price discounts are, if anything, larger in the sample of house transactions for which we have less variation across contracts.

The top panel of Figure I shows the distribution of remaining lease length for flats at the time of sale. There are many transactions with remaining lease length below 300 years and above 700 years, allowing us to trace out the term structure of leasehold discounts across long horizons. To reduce noise in our estimation, we pool leaseholds into a number of buckets with similar remaining lease length at the time of transaction, as shown in Table I. The top panel presents the composition of our sample of flats,

person purchases both the freehold and the leasehold, it is unclear what the division of price between the two titles captures. This also removes transactions of flats in buildings where the leaseholders have jointly purchased the freehold and now own a "leasehold with a share of the freehold." This procedure does not remove "freehold flats," which are different contracts from leaseholds with a share of the freehold, because they have clearly assigned individual freehold titles for each flat, as opposed to a fraction of a common freehold title.

TABLE I
U.K.: SAMPLE OVERVIEW

		Share of Transactions by Contract					
	N	80–99	100–124	125–149	150–300	700+	Freehold
Flats							
2004	183,599	0.19	0.37	0.08	0.05	0.28	0.03
2005	168,435	0.16	0.39	0.09	0.07	0.27	0.03
2006	212,734	0.14	0.39	0.10	0.08	0.27	0.03
2007	219,402	0.13	0.40	0.11	0.08	0.25	0.03
2008	116,048	0.12	0.41	0.11	0.10	0.24	0.03
2009	93,861	0.11	0.42	0.10	0.08	0.26	0.03
2010	99,663	0.13	0.41	0.09	0.08	0.27	0.02
2011	97,733	0.13	0.40	0.09	0.08	0.27	0.02
2012	98,464	0.14	0.39	0.09	0.08	0.29	0.02
2013	83,444	0.15	0.37	0.09	0.09	0.28	0.02
Total	1,373,383	0.14	0.39	0.09	0.08	0.27	0.03

	N	80–99	100–124	125–200	700+	Freehold
Houses						
2004	955,112	0.005	0.005	0.002	0.05	0.94
2005	803,983	0.005	0.005	0.002	0.05	0.94
2006	1,000,714	0.004	0.005	0.002	0.04	0.94
2007	942,575	0.004	0.006	0.002	0.05	0.94
2008	470,987	0.005	0.007	0.003	0.04	0.94
2009	480,827	0.004	0.005	0.002	0.04	0.95
2010	510,342	0.003	0.005	0.002	0.04	0.95
2011	513,179	0.004	0.004	0.002	0.04	0.95
2012	511,817	0.002	0.003	0.002	0.04	0.96
2013	438,598	0.002	0.003	0.002	0.03	0.96
Total	6,628,134	0.004	0.005	0.002	0.04	0.95

Notes. Table shows the data sample for the U.K. analysis. The top panel is for flats, the bottom panel is for houses. For each year we show the number of transactions (N), as well as the share of transactions in each bucket of remaining lease length at the point of transaction.

comprising almost 1.4 million transactions. About 3% of transactions are for freeholds, and 27% are for extremely long leaseholds (700 or more years remaining). The rest of the transactions are for shorter maturity leaseholds.

Although our data set covers all of England and Wales, it is important to verify that all types of contracts are present in most locations. We focus on the variation in lease length within three-digit postal codes; these relatively small geographical units correspond to the level of geographic fixed effects used in our hedonic analysis. Overall, flats have significant variation across contract

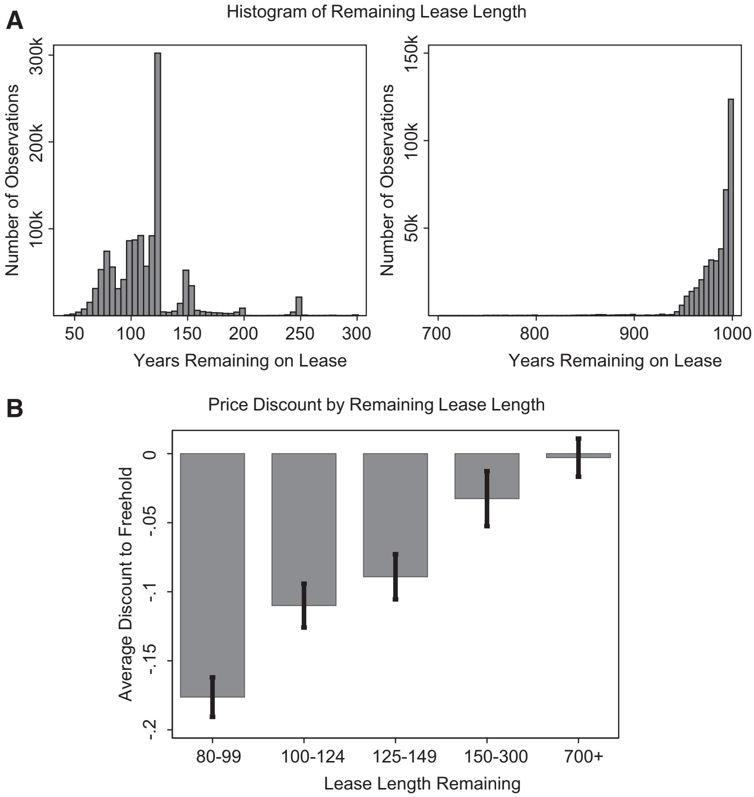


FIGURE I

U.K. Flats: Sample and Price Discounts

Panel A shows the distribution of remaining lease length at the point of sale for flats in our U.K. transaction sample. Panel B plots β_j coefficients from regression (1). The dependent variable is log price for flats sold in England and Wales between 2004 and 2013. Price discounts are relative to freeholds, and correspond to column (1) in Table III. We include three-digit postal code by transaction month fixed effects. We also control for property size, the number of bedrooms, the number of bathrooms, property age, property condition, whether there is parking, and the type of heating. The bars indicate the 95% confidence interval of the estimate using standard errors double clustered by three-digit postal code and by quarter.

types (freehold versus leasehold), within leaseholds (by number of years remaining), and across geographic areas.[3]

3. For each of the 2,375 three-digit postal codes in the United Kingdom, we compute the fraction of transactions that occur with each remaining lease length as

Table II shows summary statistics for the main hedonic variables in our analysis of flats in the United Kingdom, displayed by remaining lease length and pooled across all properties. The median flat in the United Kingdom has two bedrooms and one bathroom, has an area of 65 m^2, and is in a building that is 36 years old. The median price for a flat in the United Kingdom is £123,000. Property characteristics display some variation between freeholds and leaseholds and across leaseholds of different remaining lease length. The patterns, however, differ across characteristics. For example, all leaseholds have a very similar number of bedrooms and bathrooms; freeholds tend to have more bedrooms but fewer bathrooms than leaseholds do. Shorter leaseholds and freeholds tend to be on older buildings than leaseholds of intermediate lease length.

These summary statistics do not condition on the geographic location of the properties. To better understand the differences in observable characteristics between leaseholds and freeholds, Figure II shows the residuals of a regression of each hedonic characteristic on postal code fixed effect for each group of remaining lease length. Within each three-digit postal code we observe little systematic difference in these observable characteristics across leaseholds with different remaining lease length, which will be our main source of price variation. For the case of property age, there is some variation across leaseholds of different length, but it is not systematically related to the number of years remaining (for example, freeholds and shorter leaseholds with 80–99 years remaining tend to be on older buildings than leaseholds with maturities above 99 years). Although there is little variation within leaseholds, there is evidence that even within three-digit postal codes, freehold properties are somewhat larger than leasehold properties.

III.C. Price Variation by Lease Length Remaining in the United Kingdom

In this section we estimate the relative prices paid for leaseholds of varying maturity and freeholds for flats in England and Wales. Given the support of the "remaining lease length"

well as the fraction of freeholds. Online Appendix Table A.1 presents the distribution of the shares of contracts across postcodes. To visualize the geographic variation of freeholds and leaseholds, Online Appendix Figures A.3–A.26 provide maps of the shares of freeholds and leaseholds of different lease length remaining by postcode. The maps show significant geographic dispersion for freeholds and leaseholds in the case of flats.

TABLE II

U.K. Flats: Summary Statistics

Variable	Lease length (years)	Mean	Std. dev.	Percentile p1	p5	p25	p50	p75	p95	p99
Price (£'000)	80–99	121.1	125.7	18	29	57	91	149	290	545
	100–124	155.0	145.0	21	36	80	130	190	350	610
	125–149	177.6	183.6	25	52	103	145	205	380	750
	150–300	175.2	146.7	26	46	103	146	210	385	650
	700+	176.0	242.9	20	33	75	125	202	460	950
	Freehold	140.9	191.6	15	27	59	105	163	359	780
	Total	155.6	178.1	20	34	73	123	185	371	712
Bedrooms	80–99	1.66	0.65	1	1	1	2	2	3	3
	100–124	1.79	0.66	1	1	1	2	2	3	4
	125–149	1.83	0.60	1	1	1	2	2	3	4
	150–300	1.80	0.58	1	1	1	2	2	3	3
	700+	1.84	0.65	1	1	1	2	2	3	4
	Freehold	2.33	0.98	1	1	2	2	3	4	5
	Total	1.79	0.66	1	1	1	2	2	3	4
Bathrooms	80–99	1.08	0.29	1	1	1	1	1	2	2
	100–124	1.17	0.40	1	1	1	1	1	2	2
	125–149	1.29	0.50	1	1	1	1	2	2	3
	150–300	1.27	0.46	1	1	1	1	2	2	2
	700+	1.21	0.44	1	1	1	1	1	2	3
	Freehold	1.17	0.47	1	1	1	1	1	2	3
	Total	1.17	.40	1	1	1	1	1	2	2
Size (m^2)	80–99	66.3	48.2	29	35	49	60	73	103	161
	100–124	71.9	55.0	30	40	54	66	79	108	180
	125–149	74.0	52.4	33	43	57	67	79	115	200
	150–300	71.1	42.9	31	41	55	66	78	111	162
	700+	75.6	62.7	30	39	54	67	82	127	212
	Freehold	94.0	45.0	42	49	71	96	99	152	237
	Total	72.2	54.9	30	39	53	65	80	115	190
Age (years)	80–99	60.3	48.4	0	3	15	56	101	127	165
	100–124	44.8	44.1	0	0	10	35	67	121	158
	125–149	37.4	49.4	0	0	1	9	69	123	160
	150–300	39.4	48.9	0	0	1	21	73	123	162
	700+	52.2	60.0	0	0	10	35	97	144	205
	Freehold	61.2	56.7	0	2	19	45	100	146	253
	Total	50.3	48.7	0	0	10	36	95	128	179

Notes. Table shows summary statistics for the main hedonic variables for the sample of U.K. flats. For each characteristic, we report the statistics separately for different buckets of remaining lease length, as well as for the pooled sample.

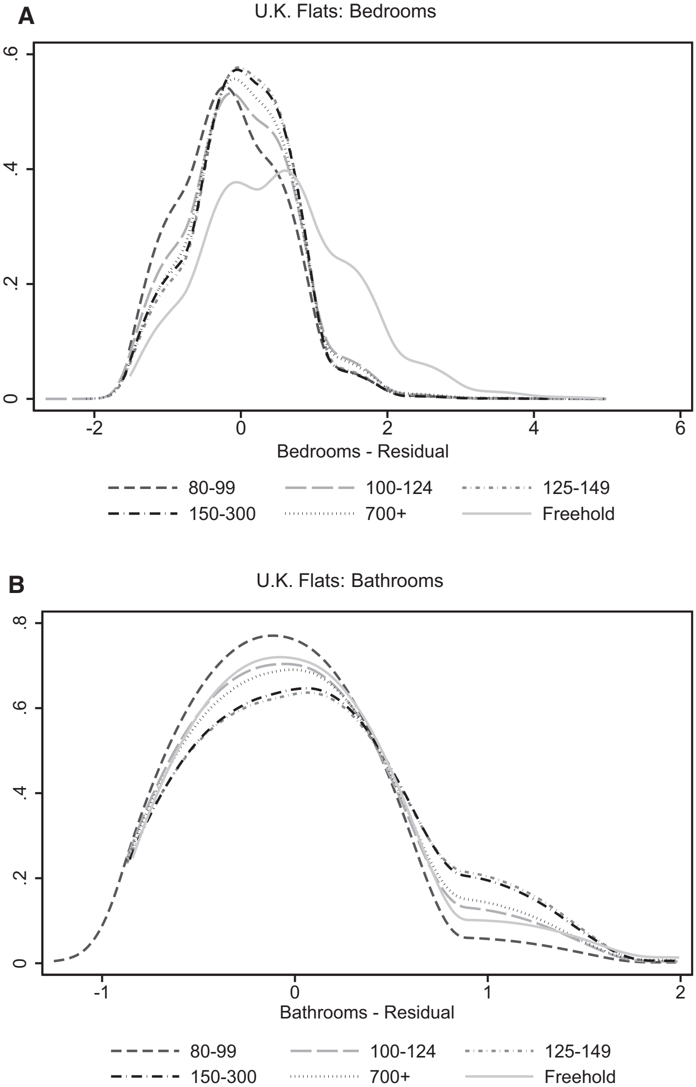


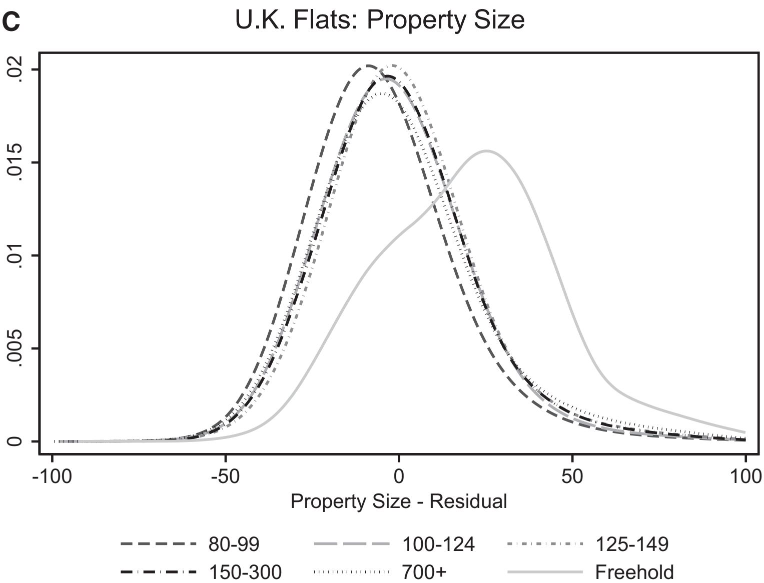
FIGURE II

Hedonic Characteristics by Lease Type

Panels A–D show the distribution of residuals from a regression of property characteristics on three-digit postal code fixed effects for freeholds and leaseholds with different remaining maturity. The sample is U.K. flats, the characteristics plotted are number of bedrooms, number of bathrooms, property size (square meters), and property age (years). Panels E and F show the distribution of residuals from a regression of property characteristics on property type × title type (strata or land) × five-digit postal code fixed effects for freeholds and leaseholds with different remaining maturity. The sample is Singapore, the characteristics plotted are property size (square meters) and property age (years).

(continued)

C



D

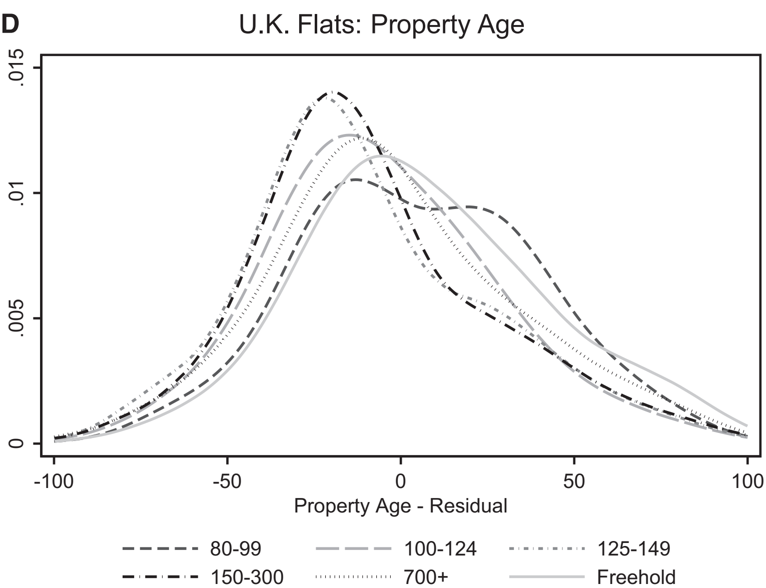


FIGURE II

(continued)

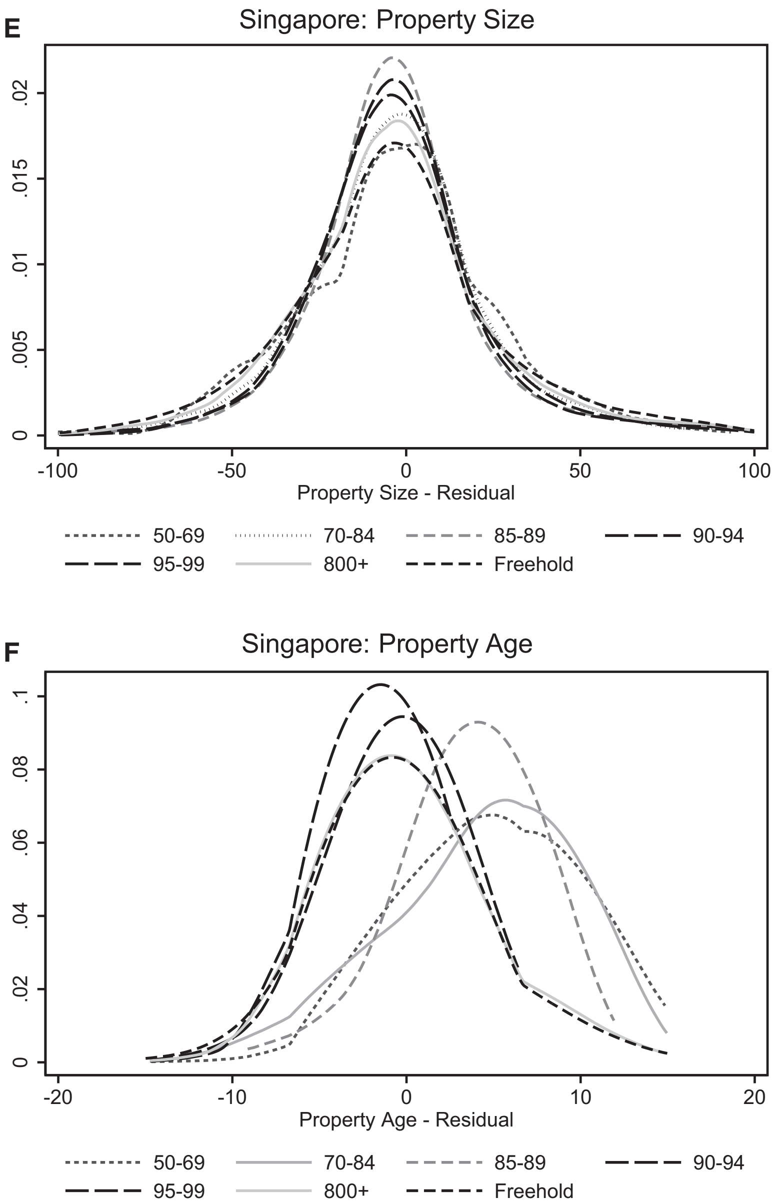


FIGURE II

(continued)

distribution (see top panel of Figure I), we construct a set *MaturityGroup* with five buckets for different remaining lease length: 80–99 years, 100–124 years, 125–149 years, 150–300 years, and 700+ years. We then estimate regression (1). The unit of observation is a transaction i of a property in three-digit postal code h at time t. We assign each leasehold transaction with remaining maturity at time of sale $T_{i,t}$ to one of the *MaturityGroup*. The β_j coefficients capture the log-discount of leaseholds with maturity in bucket j of the *MaturityGroup* relative to otherwise similar freeholds.

$$log(Price_{i,h,t}) = \alpha + \sum_{j=1}^{5} \beta_j \mathbf{1}_{\{T_{i,t} \in MaturityGroup_j\}} + \gamma Controls_{i,t} + \xi_h \times \psi_t + \epsilon_{i,h,t}. \tag{1}$$

We control for average prices in a property's geography by including three-digit postal code (ξ_h) by time of sale (ψ_t) fixed effects. We also include dummy variables for whether the property is a new construction, as well as for the number of bedrooms, bathrooms, property condition, whether there is parking, and the type of heating. We further control for the size and age of the property in a flexible way by including dummy variables for 50 equally sized groups of these characteristics. Standard errors are clustered at both the quarter and three-digit postal code level, following the procedure in Petersen (2009).

Table III shows the results from regression (1) for flats; the different columns test the robustness of our estimates to different samples and specifications. In column (1) we control for the time of sale in the interacted fixed effects by using the month of sale, in column (2) by using the quarter of sale, and in column (3) by using the year of sale. In column (4) we interact all our controls with year dummies, to allow for time variation in the effect of property characteristics. In column (5) we report the results obtained by winsorizing prices at the 1% level to reduce the possible effect of outliers. In column (6), we report the results obtained using only flats with nonmissing property characteristics.[4] In column (7) we report the results excluding the London postal district.

4. In our baseline regressions we treat missing characteristics in a different way. For each characteristic X we add a variable X_{mis} that is equal to 1 for all observations for which X is not observed, and 0 otherwise. We then set $X = 0$ whenever $X_{mis} = 1$. We estimate our regressions including both X and the dummy X_{mis}, which controls for any average difference between transactions for properties with

TABLE III

U.K. FLATS: IMPACT OF LEASE TYPE ON PRICES

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Lease length remaining							
80–99 years	−0.176***	−0.178***	−0.179***	−0.178***	−0.170***	−0.157***	−0.175***
	(0.007)	(0.007)	(0.007)	(0.002)	(0.007)	(0.008)	(0.009)
100–124 years	−0.110***	−0.109***	−0.106***	−0.110***	−0.105***	−0.111***	−0.073***
	(0.008)	(0.007)	(0.007)	(0.002)	(0.007)	(0.008)	(0.008)
125–149 years	−0.089***	−0.088***	−0.086***	−0.090***	−0.083***	−0.086***	−0.060***
	(0.008)	(0.008)	(0.007)	(0.002)	(0.008)	(0.008)	(0.009)
150–300 years	−0.033***	−0.035***	−0.034***	−0.034***	−0.028***	−0.027***	−0.012
	(0.010)	(0.009)	(0.009)	(0.002)	(0.010)	(0.009)	(0.011)
700+ years	−0.003	−0.005	−0.005	−0.005	−0.005	−0.012	−0.004
	(0.007)	(0.006)	(0.006)	(0.002)	(0.007)	(0.008)	(0.007)
Fixed effects	PC × M	PC × Q	PC × Y	PC × M	PC × M	PC × M	PC × M
Controls	✓	✓	✓	✓, × year	✓	✓	✓
Restrictions	.	.	.	.	Winsorize price	Nonmiss. hedonics	Exclude London
R-squared	0.729	0.721	0.712	0.731	0.738	0.776	0.616
N	1,373,383	1,373,383	1,373,383	1,373,383	1,373,383	953,660	1,028,031

Notes. Table shows results from regression (1). The dependent variable is the log price for flats sold in England and Wales between 2004 and 2013. We include three-digit postal code by transaction time fixed effects. In columns (2) and (3) the transaction time is the transaction quarter and year, respectively, in the other columns the transaction month. We also control for property size, the number of bedrooms, the number of bathrooms, property age, property condition, whether there is parking, and the type of heating. In column (4) we interact the controls with the transaction year. In column (5) we winsorize the price at the 1st and 99th percentiles. In column (6) we only include properties for which characteristics are not missing. In column (7) we exclude transactions in London. Standard errors are double clustered by three-digit postal code and by quarter. Significance levels: * $p < .10$, ** $p < .05$, *** $p < .01$.

The coefficients β_j of our baseline estimate, column (1), are also plotted in the bottom panel of Figure I. Freeholds and leaseholds with maturities of more than 700 years trade at approximately the same price: the coefficient on β_{700+} is small and statistically indistinguishable from 0. This suggest that the present value of rents starting in 700 years is negligible. Leaseholds with shorter maturities trade at significant discounts to otherwise identical freeholds: leaseholds with 80–99 years remaining trade at an approximately 16% discount to freeholds.[5] For the median flat, this corresponds to a price difference of approximately £20,000. The discount decreases to 10% for leaseholds with 100–124 years remaining, 8% for 125–149 years remaining, and 3% for 150–300 years remaining. The results are robust to the various specifications reported in Table III.

To our knowledge, this is the first extensive analysis of the relative valuation of leaseholds and freeholds using the universe of transactions and lease terms in England and Wales, combined with an extensive set of hedonic property characteristics. The analysis reveals substantial discounts for shorter leaseholds compared to longer leaseholds and freeholds. Interestingly, when informally investigating the priors of participants in this market (home buyers, valuers, real estate agents) we found them to be very dispersed. In particular, a number of valuers believe the discounts to be smaller than those we found in our systematic analysis, and a number of home buyers believe them to be bigger. As discussed in Online Appendix A.1.5, the priors appear to be based on either little data or introspection.[6] This dispersion is consistent with significant heterogeneity of properties, segmentation of the housing market, and the absence of a large-scale systematic empirical analysis of market valuations. In Section IV, we show that our estimated price differences are not driven by a number of frictions that could differentially affect

and without missing characteristics and allows us to keep the observations with missing values for X in our estimation. See Dickens and Katz (1987) for a description of this procedure and a discussion of different approaches to dealing with missing characteristics. As the robustness check in column (6) of Table III shows, our results are robust to various ways of dealing with missing characteristics.

5. The β_j coefficients are log-discounts. To convert into percentage discounts, compute $e^{\beta_j} - 1$.

6. Valuers at most look at about 200 transactions scattered over a number of years, 10 or more, and often use subjective judgment and exponential discounting to fill in gaps in the valuations.

the flow utility between leaseholds of different maturity and freeholds. Instead, the price differences suggest a significant present value attached by buyers to rents 100 or more years in the future, and, therefore, a relatively low discount rate over those horizons.

III.D. Singapore Residential Housing Data

We obtained transaction-level price data for all private residential transactions in Singapore from the URA. We do not use transaction prices for property sales by the Housing Development Board (HDB), which usually happen at below-market value (see Online Appendix A.2). We observe approximately 380,000 private market transactions between 1995 and 2013. For each transaction there is information on the transaction price and date, the lease terms, property characteristics and the precise location of the property. Table IV provides an overview of the transaction sample used in the regressions. There are between 10,000 and 40,000 transactions per year. Between 30% and 60% of all transactions each year are of freehold properties. We observe substantial dispersion in the lease length remaining at the time of sale. Figure III plots the remaining lease length at the time of sale for leaseholds with initially 99 years (left panel) and initially 999 years maturity (right panel). There are no transactions of leaseholds with 100 to 800 years remaining on the lease, and relatively few transactions with lease lengths less than 70 years.

For Singapore we observe fewer hedonic characteristics than for the United Kingdom; the primary characteristics are property size, development size, and property age. Table V shows summary statistics for age and size of the property in our sample. Although there is some heterogeneity in size and age across leaseholds and between leaseholds and freeholds, there do not seem to be clear patterns across maturities. For example, leaseholds with 70–84 years remaining on the lease tend to be smaller both relative to shorter leases (50–69 years) and to longer leases (800+), as well as freeholds. Age of the building is correlated with maturity, but freeholds and 800+ leaseholds tend to be in buildings that are approximately the same age as leasehold with maturity 85–94 years. As we do for the United Kingdom, to further study the differences in characteristics across contracts we estimate residuals conditional on five-digit postal code fixed effects. Figure II (bottom row) shows that there is no systematic difference in property size by lease length remaining; older properties,

TABLE IV
SINGAPORE: DATA SAMPLE

		Share of transactions by contract						
	N	50–69	70–84	85–89	90–94	95–99	800+	Freehold
1995	11,719	0.001	0.036	0.030	0.004	0.268	0.088	0.573
1996	17,514	0.001	0.025	0.021	0.024	0.303	0.145	0.481
1997	12,354	0.001	0.045	0.003	0.023	0.455	0.073	0.399
1998	13,002	0.001	0.029	0.002	0.029	0.577	0.052	0.310
1999	23,231	0.002	0.044	0.002	0.064	0.306	0.082	0.499
2000	12,387	0.007	0.050	0.005	0.094	0.269	0.085	0.492
2001	11,521	0.005	0.036	0.015	0.108	0.408	0.040	0.389
2002	17,549	0.003	0.033	0.013	0.133	0.338	0.060	0.420
2003	9,702	0.006	0.056	0.035	0.141	0.264	0.061	0.436
2004	11,203	0.006	0.049	0.050	0.123	0.192	0.058	0.522
2005	16,758	0.014	0.039	0.057	0.111	0.134	0.070	0.575
2006	24,236	0.008	0.038	0.074	0.101	0.132	0.061	0.587
2007	39,182	0.013	0.040	0.126	0.083	0.138	0.079	0.521
2008	13,911	0.015	0.056	0.159	0.084	0.138	0.073	0.475
2009	32,961	0.011	0.056	0.106	0.064	0.194	0.078	0.490
2010	34,475	0.011	0.083	0.097	0.051	0.225	0.066	0.466
2011	25,221	0.009	0.083	0.070	0.040	0.370	0.049	0.378
2012	36,633	0.016	0.085	0.039	0.040	0.444	0.047	0.329
2013	15,209	0.014	0.067	0.026	0.049	0.535	0.042	0.267
Total	378,768	0.009	0.054	0.060	0.068	0.285	0.069	0.455

Notes. Table shows the data sample for the Singapore analysis. For each year we show the number of transactions (N), as well as the share of transactions in each bucket of remaining lease length at the point of transaction.

unsurprisingly, tend to transact with fewer years remaining on the lease.

III.E. Price Variation by Lease Length Remaining in Singapore

To analyze the relative price paid for leaseholds and freeholds in Singapore, we run regression (2). The unit of observation is a property i of type g (e.g., apartment, condominium, detached house, executive condominium, semi-detached house and terrace house), of title type s (either "strata" or "land"),[7] in geography h,

7. Residential properties in Singapore are classified into land or strata titles. Land title properties occupy land that is exclusive to the owner (e.g., a detached house), whereas a strata title comprises units in cluster housing (e.g., an apartment). Owners of strata properties enjoy exclusive title only to the airspace of their individual unit. The land that the development is built on is shared by all the owners

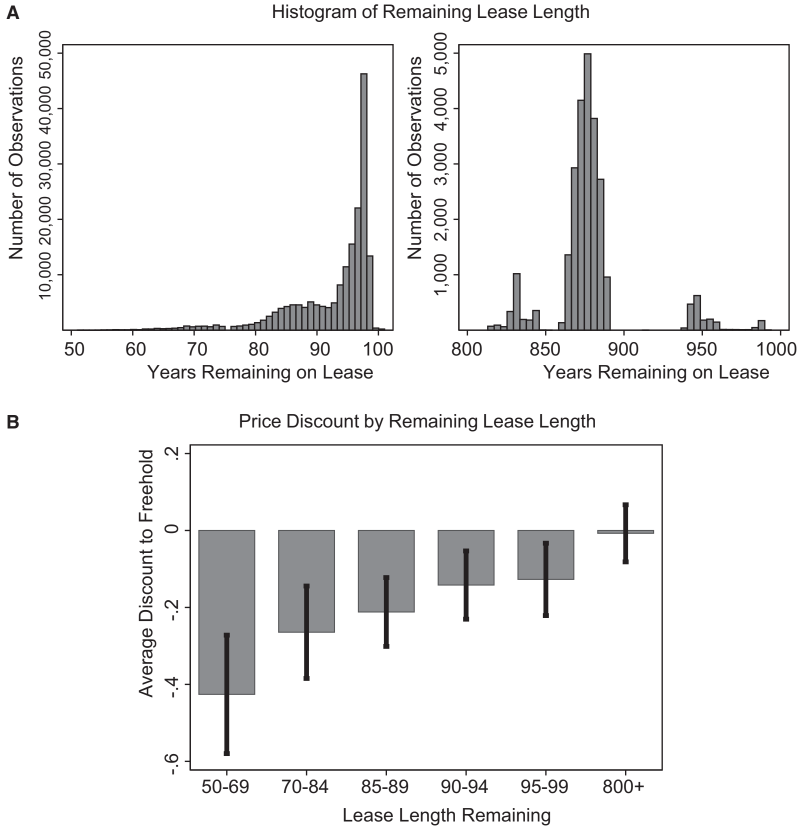


FIGURE III

Singapore: Sample and Price Discounts

Panel A shows the distribution of remaining lease length at the point of sale in our Singapore transaction sample. Panel B plots β_j coefficients from regression (2). The dependent variable is the log price for properties sold by private parties in Singapore between 1995 and 2013. Price discounts are relative to freeholds, and correspond to column (2) in Table VI. We include five-digit postal code by property type (apartment, condominium, detached house, executive condominium, semi-detached house, and terrace house) by title type (strata or land) by transaction month fixed effects. We also control for property size, property age, and the total number of units in the property. The bars indicate the 95% confidence interval of the estimate using standard errors double clustered by five-digit postal code and by quarter.

TABLE V

Singapore: Summary Statistics

Variable	Lease length (years)	Mean	Std. dev.	Percentile p1	p5	p25	p50	p75	p95	p99
Price (S$'000)	50–69	1,005.8	693.0	210	295	565	875	1,272	2,149	3,900
	70–84	929.9	567.5	305	390	590	808	1,070	1,920	3,200
	85–89	933.9	644.4	350	418	600	788	1,080	1,800	3,280
	90–94	1,038.2	1,220.1	380	435	610	783	1,100	2,203	5,030
	95–99	977.1	855.1	408	471	618	789	1,054	1,982	4,069
	800+	1,401.6	1,292.9	450	580	820	1,083	1,543	3,080	6,200
	Freehold	1,583.1	1,807.9	411	535	788	1,114	1,732	4,046	8,129
	Total	1,281.9	1,431.9	382	478	685	922	1,378	3,135	6,500
Size (m^2)	50–69	149.0	84.3	38	66	104	139	163	304	429
	70–84	139.4	64.9	68	81	103	122	157	244	355
	85–89	131.1	59.0	66	82	106	116	137	241	351
	90–94	128.5	69.4	52	64	97	114	133	266	357
	95–99	116.4	55.6	43	55	90	111	127	202	311
	800+	175.5	147.3	42	64	107	134	193	396	705
	Freehold	173.1	197.0	37	51	99	129	182	394	941
	Total	149.7	147.0	40	55	96	119	156	320	651
Age (years)	50–69	23.0	12.5	0	0	23	27	31	36	39
	70–84	13.6	6.4	0	0	12	13	17	24	28
	85–89	7.9	2.3	0	3	7	8	10	11	12
	90–94	2.4	2.0	0	0	0	2	4	6	7
	95–99	.04	0.2	0	0	0	0	0	0	1
	800+	4.5	7.4	0	0	0	0	7	21	31
	Freehold	5.2	8.8	0	0	0	0	8	24	37
	Total	4.2	7.6	0	0	0	0	6	21	32

Notes. Table shows summary statistics for the main hedonic variables for the Singapore sample. For each characteristic, we report the statistics separately for different buckets of remaining lease length, as well as for the pooled sample.

sold at time t. For leaseholds the variable $T_{i,t}$ captures the number of years remaining on the lease at the time of sale. We split the 99-year leases into five buckets with different groups of lease length remaining (50–69 years, 70–84 years, 85–89 years, 90–94 years, and 95–99 years). We also include a dummy variable for 999-year leases, all of which have at least 800 years remaining when we observe the transaction. The excluded category are

of the project, based on the share of the strata title unit owned by each owner. See Online Appendix A.2 for details.

the freeholds. The key dependent variable is the log of the price paid in the transaction.

$$\ln(Price)_{i,h,s,g,t} = \alpha + \sum_{j=1}^{6} \beta_j \mathbf{1}_{\{T_{i,t} \in MaturityGroup_j\}} + \gamma Controls_{i,t}$$
$$+\xi_h \times \rho_s \times \phi_g \times \psi_t + \epsilon_{i,h,s,g,t} \tag{2}$$

The results from this regression are shown in Table VI. In column (1) we control for five-digit postal code (ξ_h) by title type (ρ_s) by property type (ϕ_g) by transaction quarter (ψ_t) fixed effects. Beyond these 94,700 fixed effects, our other control variables include property age, size, and type, as well as the total number of units in a development. Standard errors are double clustered by five-digit postal code and by quarter.

The results are consistent with our findings for the United Kingdom: the price paid for freeholds and otherwise similar leaseholds with more than 800 years remaining is economically and statistically identical. Leases with maturities of 99 years or less sell at a significant discount to otherwise identical freeholds. For example, a leasehold with 95–99 years remaining maturity trades at an 11.8% discount, which corresponds to a S$108,000 price discount for the median flat. A leasehold with 70–84 years remaining trades at a 24% discount.[8] In column (2) we control for the transaction month rather than the transaction quarter. In column (3), rather than controlling for the age of the property directly, we focus only on the sale of newly built properties. The estimates for 95–99 year leases are unaffected. For leases with shorter maturities, the estimates of the discount increase somewhat. However, since most leases get topped up to 99 years when the property gets rebuilt, there are few observations to estimate the discount of new properties with 80 years' lease length remaining. In column (4) we restrict transactions to those where the buyer is not the HDB. The results are essentially unchanged, suggesting that sales to the HDB generally happen at market

8. The regression has a high adjusted R^2 of above 95%. This suggests that there remains no significant variation in prices that is not explained by our control variables, and that our discounts are thus unlikely to be driven by unobserved heterogeneity between freehold and leasehold properties. The adjusted R^2 remains at 95% if we exclude those instances where we only observe one transaction for a particular fixed effect, in which the fixed effects perfectly explains the transaction price. This is consistent with properties in Singapore being very homogeneous, in particular conditional on our tight geographic fixed effects.

TABLE VI

SINGAPORE: IMPACT OF LEASE TYPE ON PRICES

	(1)	(2)	(3)	(4)	(5)	(6)
Lease length remaining						
50–69 years	−0.409***	−0.426***	−0.464***	−0.487***	−0.424***	NA
	(0.069)	(0.079)	(0.034)	(0.085)	(0.074)	
70–84 years	−0.275***	−0.264***	−0.469***	−0.273***	−0.262***	−0.328**
	(0.054)	(0.061)	(0.047)	(0.074)	(0.058)	(0.139)
85–89 years	−0.215***	−0.212***	−0.111**	−0.216***	−0.210***	NA
	(0.040)	(0.046)	(0.049)	(0.054)	(0.043)	
90–94 years	−0.148***	−0.142***	−0.169***	−0.146***	−0.142***	−0.179
	(0.040)	(0.045)	(0.044)	(0.052)	(0.042)	(0.675)
95–99 years	−0.125***	−0.127***	−0.127**	−0.132***	−0.129***	−0.213
	(0.042)	(0.048)	(0.055)	(0.049)	(0.045)	(0.621)
800+ years	−0.010	−0.007	0.019	−0.002	−0.008	0.006
	(0.033)	(0.038)	(0.049)	(0.043)	(0.036)	(0.127)
Fixed effects	PC × Q × Prop type × Title type	PC × M × Prop type × Title type	PC × M × Prop type × Title type	PC × M × Prop type × Title type	PC × M × Prop type	PC × M × Prop type
Controls	✓	✓	✓	✓	✓	✓
Restrictions	.	.	New only	Private buyer	Strata only	No strata
R-squared	0.977	0.979	0.981	0.978	0.977	0.985
N	378,768	378,768	223,810	220,044	333,684	45,084

Notes. Table shows results from regression (2). The dependent variable is the log price for properties sold by private parties in Singapore between 1995 and 2013. We include five-digit postal code by property type (apartment, condominium, detached house, executive condominium, semi-detached house and terrace house) by title type (strata or land) by transaction time fixed effects. In column (1), the transaction time is transaction quarter, in columns (2)–(6) it is transaction month. We also control for property age, property size, and the total number of units in the property. In column (3) we focus on properties that were built within three years of the transaction date. In column (4) we focus on properties that were bought by private individuals (and not the HDB). In columns (5) and (6) we conduct the analysis for strata and nonstrata titles separately. Standard errors are double clustered by five-digit postal code and by quarter. Significance levels: * $p < .10$, ** $p < .05$, *** $p < .01$.

value. In column (5) we restrict our analysis to strata properties, which comprise the majority of all title types; in column (6) we restrict the analysis to land titles. Leaseholds of 999 years and freeholds trade at the same price. There are very few land title properties trading on 99-year leases, making it hard to estimate the lower end of the term structure of leasehold discounts. Nevertheless, while the estimates are very noisy (and there are not sufficient data to estimate every bucket), the point estimates for the land title and strata regressions are similar.

IV. Frictions and Leasehold Discounts

In Section III we estimated significant price differences between 100-year leaseholds and otherwise similar freeholds or leaseholds with longer maturity. In Section V we will show that these leasehold discounts are informative about very long-run discount rates. Before discussing the interpretation of our results in terms of discount rates, however, we explore several ex ante plausible alternative interpretations of the estimated price discounts, all of which involve the flow utility being different across the different contracts: (i) the presence of unobserved structural heterogeneity across properties, (ii) the impact of leasehold covenants, (iii) the impact of differences across freeholders, (iv) a different clientele for freeholds and leaseholds, (v) differential liquidity of freehold and leasehold properties, (vi) financing frictions for short leaseholds, (vii) the impact of taxation, and (viii) hold-up problems at lease extension. Overall, there is no evidence that the flow utility from owning properties differs across leaseholds of different maturity and freeholds. In terms of the model in the introduction, thus, we argue that conditional on our observable control variables, D is the same across all properties. This will allow us to interpret the estimated discounts in terms of net discount rates, $r-g$.

IV.A. Unobserved Structural Heterogeneity

Even after controlling for all observed differences across properties (such as property size and age), there potentially remains unobserved structural heterogeneity that could vary systematically by lease length. Although an econometrician is unable to observe all characteristics that might affect the flow utility from a property, to the extent that these characteristics affect the relative prices paid for leaseholds and freeholds they should also affect the annual market rents for these properties. Conversely, if our control variables correctly capture all sources of systematic heterogeneity across properties, rents should not differ systematically between freeholds and leaseholds of different maturity.

To test this, Rightmove provided us with a sample of around 29,000 rental listing prices for flats with a full set of property characteristics listed in London during 2011 and 2012. Figure IV, Panel A shows the price discounts in “for sale” transactions for our full sample as well as the subsample for which we observe rental prices; this regression is identical to column (1) in Table III. Price discounts are very similar in both samples, suggesting that our sample with rental data is representative on this important dimension.

In columns (1)–(3) of Table VII we estimate different specifications of regression (1) using the log of annual rents as the dependent variable. There is no significant and systematic difference between rental rates of freeholds and leaseholds of different maturity.[9] Figure IV, Panel B shows the rental discounts graphically. These results provide support to the assumption that our controls are correctly capturing the main heterogeneity across properties. This is consistent with the observation that conditional on geography, observable characteristics did not vary significantly across leasehold maturity. Finally, an additional piece of evidence that our hedonic regression allows us to control for all important structural differences is that once we control for those observables, there is no observed price difference between 700+ year leaseholds and freeholds.

IV.B. Leasehold Covenants or Contract Structure

A second alternative interpretation of the results is that buyers might perceive an intrinsic difference between owning a leasehold and owning a freehold (for example, because of restrictions on leaseholders to redevelop the property, or because of a pure psychological preference for freehold ownership). To demonstrate that this is not the case, we show that the price discounts

9. Only the rental prices for the 700+ year bucket are statistically different to the rents for the other buckets. However, Figure IV, Panel A shows that in this subsample, properties in that bucket also sell for a higher price than freeholds, which they do not in the full sample.

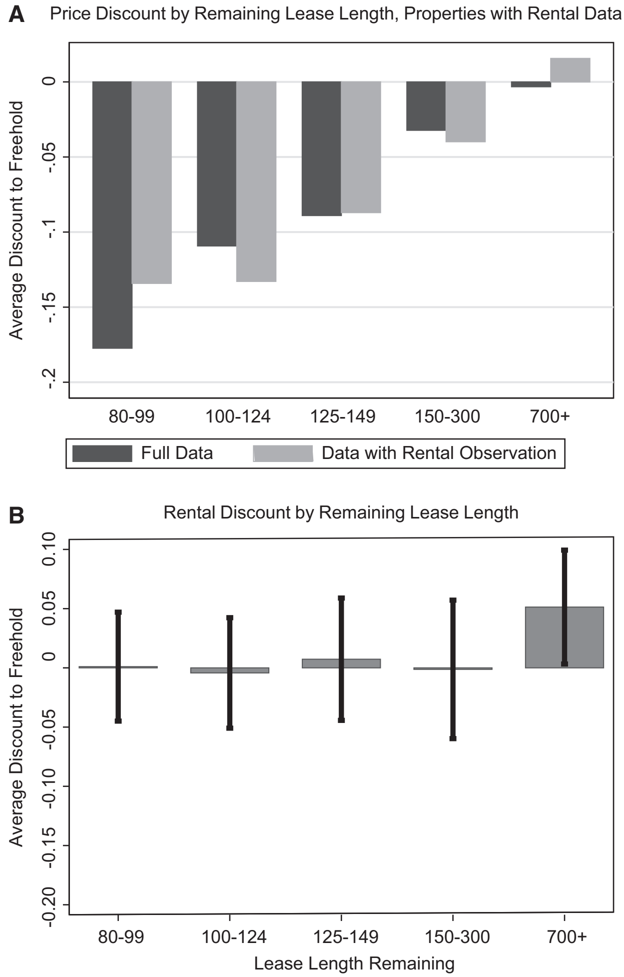


FIGURE IV

Rental Data Analysis

Panel A shows β_j coefficients from regression (1) as in column (1) of Table III, allowing for a different coefficient for properties for which we observe rental data (second bar). Panel B shows β_j coefficients from regression (1), as in column (2) of Table VII. The dependent variable in Panel B is the log of monthly rents. The sample for rent data is the universe of London flats for which rent and hedonic characteristics are available on Rightmove.co.uk during 2011 and 2012. We control for three-digit postal code by listing month fixed effects, as well as for property size, the number of bedrooms, the number of bathrooms, property age, property condition, whether there is parking, and the type of heating. The bars indicate the 95% confidence interval of the estimate using standard errors double clustered by three-digit postal code and by listing month.

TABLE VII

RENTS AND TIME ON MARKET ANALYSIS

Dependent variable:	Log(rent)			Log(time on market)		
	(1)	(2)	(3)	(4)	(5)	(6)
Lease length remaining						
80–99 years	−0.021	0.001	0.008	0.059***	0.060***	0.047***
	(0.028)	(0.023)	(0.018)	(0.010)	(0.010)	(0.014)
100–124 years	−0.024	−0.004	0.003	0.048***	0.048***	0.022*
	(0.028)	(0.024)	(0.019)	(0.008)	(0.008)	(0.013)
125–149 years	−0.005	0.007	0.014	0.063***	0.060***	0.059***
	(0.030)	(0.026)	(0.021)	(0.015)	(0.015)	(0.020)
150–300 years	−0.020	−0.001	0.009	0.080***	0.076***	0.071***
	(0.033)	(0.030)	(0.024)	(0.011)	(0.010)	(0.017)
700+ years	0.037	0.051**	0.057***	0.028***	0.028***	0.017***
	(0.029)	(0.025)	(0.018)	(0.005)	(0.005)	(0.006)
Fixed effects	PC and M	PC × M	PC × M	PC × M × Prop type	PC × M × Prop type	PC × M × Prop type
Controls	✓	✓	✓	✓	✓, × year	✓
Restrictions	Nonmiss. Hedonics	Nonmiss. hedonics	Nonmiss. hed, winsor.	·	·	Nonmiss. hedonics
R-squared	0.674	0.746	0.766	0.070	0.092	0.073
N	29,020	29,020	29,020	2,409,181	2,409,181	1,290,825

Notes. Table shows results from regression (1). The dependent variable is the log of monthly rents in columns (1)–(3), and the log of time on market between listing and sale in days in columns (4)–(6). The sample for rent data are the universe of London flats for which rent and hedonic characteristics are available on Rightmove.co.uk during 2011 and 2012. The sample for the time on market analysis is restricted to the sales of properties (both houses and flats) for which Rightmove observes a "for sale" listing. In column (3) we winsorize the dependent variable at the 1st and 99th percentiles. We control for property size, the number of bedrooms, the number of bathrooms, property age, property condition, whether there is parking, and the type of heating. Standard errors are double clustered by three-digit postal code, and by transaction month (columns (1)–(3)) of transaction quarter (columns (4)–(6)). Significance levels: * $p < .10$, ** $p < .05$, *** $p < .01$.

remain the same when we exclude freeholds and use the longest leaseholds as the excluded category (700+ years).[10]

This analysis, however, does not yet rule out that covenants or other contract characteristics might be more restrictive only on shorter leases, which could affect the relative valuation of leaseholds of different remaining maturity. Ideally, we would like to estimate the price differences between leaseholds with different remaining maturity at the time of transaction, controlling for the initial lease length, which would capture differential covenants attached to contracts with different initial maturity. In practice, this is complicated by the high degree of collinearity between initial and remaining lease length.[11]

To deal with this collinearity, we conduct a number of empirical tests, none of which suggest that our results are explained by covenants that vary systematically with initial lease length. In a first test, we run regression (1) while including fixed effects for the 10 most common initial lease lengths (which represent 92% of all transactions).[12] Table VIII shows the results. The inclusion of these fixed effects does not significantly affect the estimated price discounts. The discount for leaseholds with 80–99 years remaining relative to freeholds drops from about 16% to 13%; a small part of the initial 16% discount is now attributed to the fixed effects for contracts with 99 years' initial maturity. In addition, the coefficients on the initial lease length fixed effects do not systematically vary with contract maturity. Overall, we find no evidence that differential covenants are an important explanation for the estimated discounts.

10. See Online Appendix Table A.3 and the top panel of Appendix Figure A.27. We conduct a similar analysis for Singapore, using leasehold with 95–99 years remaining maturity as the excluded category. See Online Appendix Table A.4 and the bottom panel of Appendix Figure A.27.

11. The identification of the initial lease length effects—separate from the effect of the years remaining at the time of transaction—relies on observing transactions with approximately the same number of years remaining (say, 125–150) but very different initial lease length. Unfortunately, for all buckets except 80–99 years, there are very few transactions of contracts with higher initial lease length. For example, only 3% of the properties trading with 125–150 years remaining have an initial lease length of above 150 years.

12. We include fixed effects for initial lease lengths of 99, 120, 125, 150, 155, 199, 200, 250, 800, and 999 years. Results are robust to also including fixed effects for the 10 next most common initial lease lengths. In addition to the controls in Table III, we include an indicator for whether transactions occur as a new contract is started.

TABLE VIII

U.K. Flats: Analysis with Contract Type Fixed Effects

	(1)	(2)	(3)	(4)	(5)	(6)
Lease length remaining						
80–99 years	−0.123***	−0.124***	−0.123***	−0.126***	−0.118***	−0.120***
	(0.010)	(0.011)	(0.011)	(0.003)	(0.010)	(0.011)
100–124 years	−0.096***	−0.098***	−0.100***	−0.095***	−0.105***	−0.070***
	(0.010)	(0.010)	(0.011)	(0.002)	(0.009)	(0.011)
125–149 years	−0.083***	−0.086***	−0.088***	−0.083***	−0.089***	−0.069***
	(0.010)	(0.010)	(0.011)	(0.003)	(0.009)	(0.012)
150–300 years	−0.055***	−0.054***	−0.053***	−0.057***	−0.051***	−0.049***
	(0.011)	(0.011)	(0.011)	(0.003)	(0.008)	(0.013)
700+ years	0.013	0.013	0.013	0.012***	−0.006	0.010
	(0.020)	(0.021)	(0.023)	(0.003)	(0.012)	(0.010)
Contract type: initial lease length						
99 years	−0.065***	−0.064***	−0.063***	−0.062***	−0.055***	−0.067***
	(0.010)	(0.010)	(0.011)	(0.002)	(0.009)	(0.009)
120 years	−0.036***	−0.031**	−0.026**	−0.035***	−0.019*	−0.071***
	(0.012)	(0.012)	(0.013)	(0.003)	(0.011)	(0.015)
125 years	−0.006	−0.008	−0.008	−0.008***	−0.003	−0.001
	(0.009)	(0.009)	(0.010)	(0.002)	(0.008)	(0.009)
150 years	0.068***	0.070***	0.073***	0.066***	0.063***	0.070***
	(0.010)	(0.010)	(0.010)	(0.002)	(0.008)	(0.011)
155 years	0.044***	0.047***	0.051***	0.044***	0.043***	0.051***
	(0.013)	(0.013)	(0.014)	(0.003)	(0.011)	(0.015)
199 years	0.038*	0.035	0.032	0.039***	0.027	0.046*
	(0.022)	(0.023)	(0.025)	(0.003)	(0.017)	(0.026)
200 years	0.039**	0.038**	0.038**	0.040***	0.036**	0.031*
	(0.016)	(0.017)	(0.018)	(0.004)	(0.014)	(0.017)
250 years	0.035**	0.029*	0.028*	0.040***	0.042***	0.059***
	(0.015)	(0.016)	(0.017)	(0.003)	(0.015)	(0.015)
800 years	0.111	0.070	0.067	0.103**	0.272**	0.137
	(0.138)	(0.134)	(0.138)	(0.044)	(0.103)	(0.139)
999 years	−0.011	−0.011	−0.009	−0.012***	−0.005	−0.010
	(0.018)	(0.019)	(0.021)	(0.003)	(0.010)	(0.009)
Fixed effects	PC × Y	PC × Q	PC × M	PC × Y	PC × Y	PC × Y
Controls	✓	✓	✓	✓, × year	✓	✓
Restrictions	.	.	.	.	Nonmiss. hedonics	Exclude London
R-squared	0.715	0.724	0.732	0.717	0.766	0.586
N	1,373,383	1,373,383	1,373,383	1,373,383	953,660	1,028,031

Notes. Table shows results from regression (1), including fixed effects for the most common initial lease lengths. The dependent variable is the log price for flats sold in England and Wales between 2004 and 2013. We include three-digit postal code by transaction time fixed effects. In columns (2) and (3) the transaction time is the transaction quarter and month, respectively, in the other columns the transaction year. We also control for property size, the number of bedrooms, the number of bathrooms, property age, property condition, whether there is parking, and the type of heating. In column (4) we interact the controls with the transaction year. In column (5) we only include properties with nonmissing characteristics. In column (6) we exclude transactions in London. Standard errors are double clustered by three-digit postal code and by quarter. Significance levels: * $p<.10$, ** $p<.05$, *** $p<0.01$.

Since we observe significant variation in initial lease length for the group of transactions with 80–99 years remaining at the time of sale, we analyze this group in more detail in Table IX. Columns (1) and (2) show results from a variant of regression (1) where the price discount for this group is allowed to differ between contracts with initial lease length above or below 99 years. Consistent with Table VIII, we only find a small difference between the two groups of initial lease length.[13] A different specification, presented in columns (5) and (6), restricts the sample to transactions of leaseholds with 80–99 years remaining, and includes initial lease length fixed effects for the eight most common initial lease lengths in this window. This avoids having to estimate the level of the price discount, and only looks at the differential prices of different contracts that all trade with roughly the same number of years remaining. There is no systematic pattern in pricing across initial lease lengths.

Overall, the analysis in Tables VIII and IX suggests that even after controlling (to the extent possible) for the initial lease length, the discounts related to the remaining leasehold maturity are large and significant. In addition, the term structure of leasehold discounts we estimate for Singapore between 50 and 99 years (see Online Appendix Table A.4) keeps the initial length constant (all are 99-year contracts), and thus cannot be explained by differential initial lease length effects. Two additional pieces of evidence confirm that leasehold covenants are unlikely to have a significant confounding impact on our analysis. First, to the extent that restrictive covenants affect the flow utility from the property (for example, because they require a certain configuration of the flat), these restrictions should be passed onto renters of the property. The absence of differential rents across leaseholds of different maturity makes it unlikely that there are significant differences in restrictive covenants. Second, a manual inspection of covenants on 801 leasehold properties with different lease lengths in postal code E16 (East London) suggests that the type of covenants included does not vary with lease length (see Online Appendix A.1.7.1).

13. Columns (3) and (4) show that in addition, there is essentially no difference in discounts for leaseholds with 100–124 years remaining between contracts with original lease length below or above 125 (though only 2% of the transactions are in the latter group).

TABLE IX

U.K. FLATS: INITIAL LEASE LENGTH ANALYSIS

	Panel A					Panel B	
	(1)	(2)	(3)	(4)		(5)	(6)
Contract type					Initial lease length		
80–99 years,	−0.133***	−0.133***	−0.133***	−0.133***	99 years	−0.065***	−0.067***
Init. lease >99	(0.008)	(0.010)	(0.008)	(0.010)		(0.013)	(0.018)
80–99 years,	−0.185***	−0.183***	−0.185***	−0.183***	100 years	−0.138***	−0.141***
Init. lease ≤99	(0.008)	(0.008)	(0.008)	(0.008)		(0.020)	(0.026)
100–124 years,			−0.096***	−0.097***	105 years	0.025	0.054
Init. lease >125			(0.014)	(0.017)		(0.067)	(0.085)
100–124 years,			−0.102***	−0.107***	120 years	−0.052***	−0.037*
Init. lease ≤125			(0.007)	(0.008)		(0.016)	(0.022)
100–124 years	−0.102***	−0.107***			124 years	0.013	0.003
	(0.007)	(0.008)				(0.027)	(0.036)
125–149 years	−0.056***	−0.060***	−0.056***	−0.060***	125 years	−0.036**	−0.041*
	(0.007)	(0.008)	(0.007)	(0.008)		(0.018)	(0.026)
150–300 years	−0.024**	−0.023**	−0.024**	−0.023**	126 years	−0.074***	−0.088***
	(0.009)	(0.010)	(0.009)	(0.010)		(0.025)	(0.034)
700+ years	0.002	0.004	0.002	0.004	130 years	−0.025	−0.016
	(0.006)	(0.007)	(0.006)	(0.007)		(0.024)	(0.032)
Fixed effects	PC × Y	PC × M	PC × Y	PC × M		PC × Y	PC × M
Controls	✓	✓	✓	✓		✓	✓
Restrictions	.	.	.	.		80–99 Years Remaining	
R-squared	0.714	0.731	0.714	0.731		0.772	0.787
N	1,373,383	1,373,383	1,373,383	1,373,383		197,341	197,341

Notes. Panel A shows results from regression (1), where the dummies for buckets of remaining lease length are interacted with indicators for the initial lease length of the contract. The dependent variable is log price, for flats sold in England and Wales between 2004 and 2013. Columns (1) and (2) interact the indicator for the group of transactions with 80–99 years remaining with an indicator of whether the initial length of the leasehold transacted was below or above 99 years. Columns (3) and (4) also interact the indicator for the group of transactions with 100–124 years remaining with an indicator of whether the contract had initial lease length above or below 125 years. Panel B runs the same hedonic regression restricted to leaseholds that transact with 80–99 years remaining, and includes fixed effects for the eight most common initial lease lengths. In all regressions, we include three-digit postal code by transaction time fixed effects. In columns (1), (3), and (5) the transaction time is the transaction year, in the other columns the transaction month. We also control for property size, the number of bedrooms, the number of bathrooms, property age, property condition, whether there is parking, and the type of heating. Standard errors are double clustered by three-digit postal code and by quarter. Significance levels: * $p < .10$, ** $p < .05$, *** $p < .01$.

IV.C. Heterogeneous Freeholders

One further possible explanation for the estimated price discounts of leaseholds with relatively short remaining maturities relates to differences across freeholders. If freeholders meaningfully differed in their treatment of leaseholders, this might affect leaseholders' incentives and ability to extend an existing lease. In equilibrium, this could generate an endogenous correlation between remaining lease length and freeholder characteristics, since leaseholds with more problematic freeholders would be extended less frequently and would thus trade with fewer years remaining. If leaseholds with less attractive freeholders also sold for a lower price, this might generate a correlation between the observed discounts and the years remaining on the lease at the time it transacts that is not directly related to differences in the maturity of cash flows.

There are a number of reasons such differences across freeholders cannot explain our estimated price discounts. First, if such differences were economically important, they would affect the incentives of leaseholders to maintain the property: a leaseholder who is unlikely to extend the lease because of disagreements with the freeholder will invest less in maintenance. Similarly a prospective renter would be willing to pay less to live in a building that is not smoothly run or the common areas of which are not well maintained due to open conflicts between leaseholders and the freeholder. The fact that annual rents do not differ across remaining lease lengths therefore already suggests that differences across freeholders are likely to be small.

In addition, to directly address this concern, we further homogenize the estimation sample by only exploiting variation in remaining lease length of flats within the same building, which are generally owned by the same freeholder. We focus on flats with remaining lease lengths of less than 300 years, since in buildings with initial lease lengths of 999 years all properties are part of the 700+ years remaining bucket (and therefore there is no within-building variation to exploit). There are about 40,000 transactions of flats that sell in the same building and quarter as at least one other flat in a different remaining lease length bucket. The remaining lease length generally differs because of a differential history of lease extensions, but can also be caused by differential initial lease lengths. Figure V shows residuals of a regression of different property characteristics on

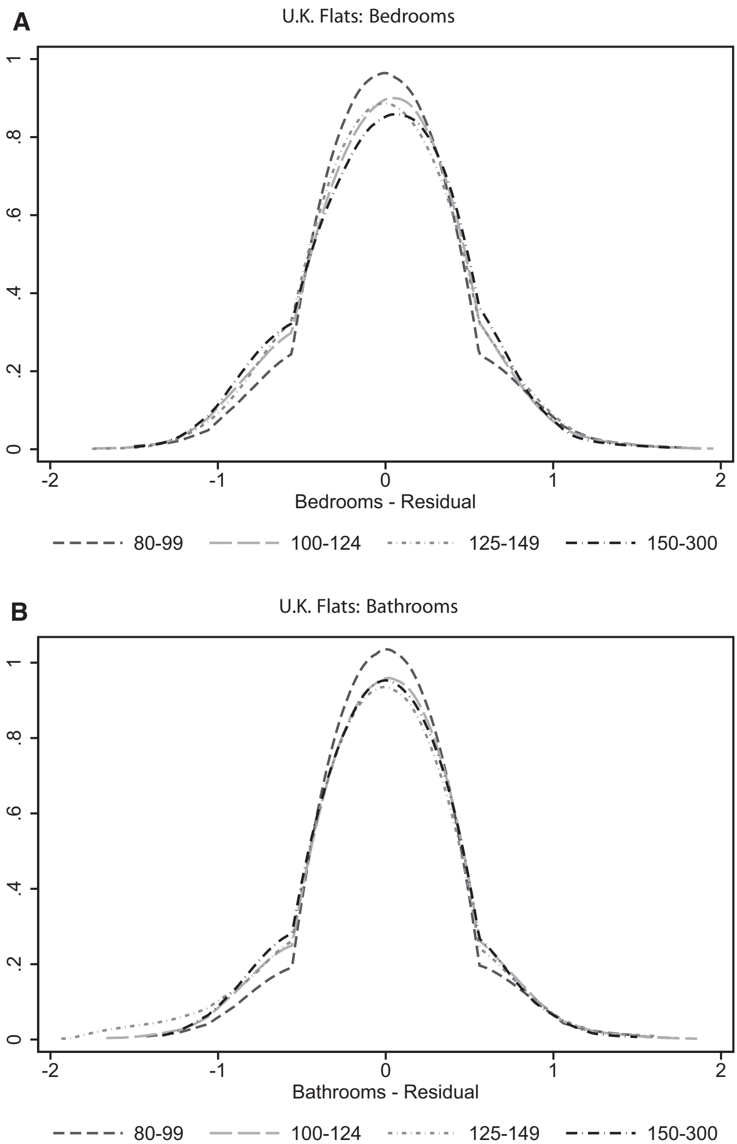


FIGURE V

Hedonic Characteristics by Lease Type: Within-Building Analysis

Figure shows the distribution of residuals from a regression of property characteristics on building fixed effects for leaseholds with different remaining maturity. The sample consists of all U.K. flats with less than 300 years' maturity remaining that sell in the same building and quarter as at least one other flat in a different remaining lease length bucket. The characteristics plotted are number of bedrooms (Panel A), number of bathrooms (Panel B), and property size in square meters (Panel C).

(continued)

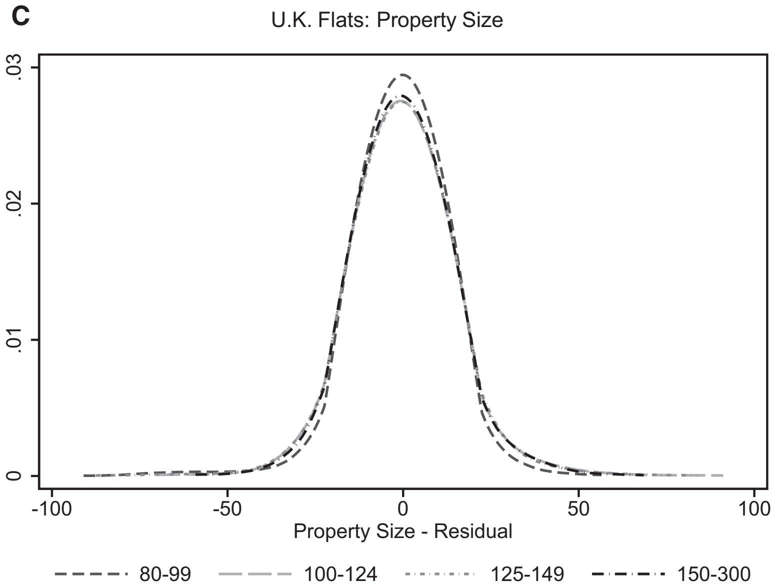


FIGURE V

(continued)

building fixed effects by remaining lease length and would allow us to detect any difference in characteristics across flats with different lease lengths in the same building. Within a building, differential remaining lease length is not systematically related to characteristics of the property, and, by construction, orthogonal to characteristics of the freeholder.

Column (1) of Table X shows coefficients of regressions of transaction price on building $\times$ transaction quarter fixed effects, other hedonic control variables, and dummy variables capturing different remaining lease length buckets. The excluded category are leaseholds with 150–300 years remaining. The evidence is highly consistent with our previous estimates; for example, leaseholds with 80–99 years remaining trade at about a 14% discount to leaseholds with 150–300 years remaining. In column (2) we include building $\times$ transaction month fixed effects, and in column (3) we winsorize the the dependent variable. In columns (4)–(6) we expand the sample to include all properties with less than 300 years remaining lease length. Although the coefficients on remaining lease length buckets continue to be identified by the

TABLE X

U.K. FLATS: IMPACT OF LEASE TYPE ON PRICE LEVEL, WITHIN BUILDING

	(1)	(2)	(3)	(4)	(5)	(6)
Lease length remaining						
80–99 years	−0.156***	−0.158***	−0.151***	−0.159***	−0.132***	−0.154***
	(0.025)	(0.030)	(0.025)	(0.046)	(0.041)	(0.046)
100–124 years	−0.080***	−0.053***	−0.078***	−0.084**	−0.057*	−0.082**
	(0.020)	(0.019)	(0.020)	(0.036)	(0.030)	(0.036)
125–149 years	−0.064***	−0.044**	−0.061***	−0.065**	−0.049*	−0.063*
	(0.018)	(0.019)	(0.018)	(0.033)	(0.030)	(0.033)
Fixed effects	Building × quarter	Building × month	Building × quarter	Building × quarter	Building × month	Building × quarter
Controls	✓	✓	✓	✓	✓	✓
Restrictions	.	.	Winsorize	.	.	Winsorize
R-squared	0.798	0.836	0.803	0.887	0.890	0.890
N	40,010	40,010	40,010	971,965	971,965	971,965

Notes. Table shows results from regression (1). The dependent variable is log price, for flats sold in England and Wales between 2004 and 2013 with less than 300 years remaining lease length. Columns (1)–(3) only include transactions of flats that sell in the same building and quarter as at least one other flat in a different remaining lease length bucket. We include building by transaction time fixed effects. In columns (2) and (5) the transaction time is the transaction month, in the other columns the transaction quarter. We also control for property size, the number of bedrooms, the number of bathrooms, property age, property condition, whether there is parking, and the type of heating. In columns (3) and (6) we winsorize the price at the 1st and 99th percentiles. Standard errors are double clustered by three-digit postal code and by quarter. Significance levels: * $p < .10$, ** $p < .05$, *** $p < .01$.

40,000 flats that experience within-building variation in remaining lease length, the additional transactions help better identify the coefficients on the control variables. The estimates are highly robust across specifications.

In addition to alleviating concerns about differences across freeholders, these results also further address some of the other concerns already discussed. First, this sample is even less likely to differ systematically on unobservable property characteristics—for example, by definition the condition and maintenance of the structure do not vary within building. Second, covenants, contracts for maintenance and servicing, and restrictions on property redevelopment generally do not vary within a building. Therefore, our estimates are robust to concerns that the price differences between leaseholds of different maturity are driven by correlation of remaining lease length with any of these factors.

For Singapore, a similar analysis is not possible, because the leaseholds for all flats in a building usually get renewed at the same time. However, it is important to bear in mind that the vast majority of leaseholders in Singapore have the same freeholder, the SLA. While the SLA's willingness to extend leases might differ depending on planning intentions in that area (see Section II.B), within our five-digit postal codes, which generally do not include more than 10 buildings, such planning intentions are likely to be very similar. This means that the possible endogeneity of lease extensions is significantly mitigated by our very tight geographic fixed effects.

Finally, we also want to reemphasize that our results are very consistent across the United Kingdom and Singapore, despite many differences in the institutional settings. In this case, for example, the U.K. features many different freeholders, whereas in Singapore the SLA is the main freeholder. At a basic level, finding very similar discounts across these two setups is an indication that freeholder heterogeneity is unlikely to be an important driver of our results.

IV.D. Heterogeneous Buyers

We observe no characteristics of the buyers in our transaction sample. Consequently, there might be a concern that the clientele for leasehold and freehold properties is different, which could help explain the price differences that we observe.

To address this concern, we analyze data from the Survey of English Housing, an annual household-level survey conducted between 1994 and 2008 that allows us to compare characteristics of owners of freeholds and leaseholds. We focus on the sample of 201,933 owner-occupiers.

Table XI presents the results of a variant of regression (1) using each individual characteristic of the owners as a dependent variable. The coefficients on the leasehold indicator then represent the average difference in each characteristic between leaseholders and freeholders, controlling for property type by region fixed effects (columns (4) and (5)) and property characteristics (column (5)).[14] Households owning freeholds and leaseholds are very similar. For example, the weekly income of households owning leasehold properties is between £3 less and £8 more than the income of households owning freeholds; this difference is insignificant and small relative to a sample mean and standard deviation of £350 and £450, respectively. The lack of major differences across buyers conditional on observable characteristics makes it unlikely that our results are driven by clientele effects related to, for example, differential bequest motives.

IV.E. Differential Market Liquidity

Leasehold and freehold properties could potentially be differentially liquid in the resale market, in which case our estimated price differences might capture a liquidity discount that increases as lease length declines. To test whether this hypothesis explains the estimated price discounts, Rightmove provided us with for-sale listing information for about 2.4 million transactions of flats and houses. For these transactions we calculate the time between first listing and sale, that is, the time on the market (see Online Appendix Figure A.28 for the distribution), which provides a proxy for the liquidity of the asset (see Genesove and Han 2012; Piazzesi, Schneider, and Stroebel 2013).

To test whether liquidity differs by maturity of the lease, columns (4) to (6) of Table VII repeat the analysis of regression (1) using time on the market as the dependent variable. The results show that leaseholds tend to stay a modest 3–6% longer on

14. Geographic controls here are more coarse than in previous sections, because the Survey of English Housing only reports 354 unique local authority codes. Property controls are those observed in both the Survey of English Housing and the transaction data set, such as the number of rooms and the property age.

TABLE XI

CHARACTERISTICS OF BUYERS OF LEASEHOLDS AND FREEHOLDS: U.K.

	Sample		Leasehold Δ		
	Mean (1)	Std. dev. (2)	Unconditional (3)	Conditional I (4)	Conditional II (5)
Age head of household (years)	52.30	16.01	−2.68	−1.54	−1.30
				(0.21)	(0.20)
Weekly income (£)	350.2	450.6	−48.07	−3.01	5.60
				(4.56)	(4.45)
Number of people in household	2.53	1.27	−0.48	−0.03	0.02
				(0.01)	(0.01)
Number of dependent children	0.55	0.94	−0.19	−0.01	0.02
				(0.01)	(0.01)
Head of household married	0.64	0.48	−0.21	−0.01	0.01
				(0.01)	(0.01)
First-time buyer	0.40	0.48	0.11	−0.00	−0.01
				(0.01)	(0.01)
Currently has mortgage	0.59	0.49	0.03	0.02	0.02
				(0.01)	(0.01)
Very satisfied with neighborhood	0.47	0.50	−0.06	0.00	0.00
				(0.00)	0.00

Notes. Table shows summary statistics on characteristics of owners of freeholds and leaseholds in the Survey of English Housing. The data contains information on 201,933 households in England surveyed between 1994 and 2008. The first two columns provide the sample means of the outcome variables. Column (3) provides the unconditional average difference between leasehold owners and freehold owners (e.g., on average, heads of household owning leasehold properties are 2.7 years younger than heads of households owning freehold properties). Columns (4) and (5) show the β coefficient of the following regression: $Outcome_i = \alpha + \beta Leasehold_i + \xi X_i + \phi_{PropertyType \times Region} + \varepsilon_i$. Column (4) does not include any additional controls in X_i, column (5) includes dummy variables for property age and the number of rooms (the control variables that overlap with the transaction data set). Standard errors are double clustered by local authority code and by year.

the market than freeholds, relative to a mean of about 160 days. Importantly, there is no pattern between remaining lease length and time on market that could explain the significant discounts we found in comparing short and long leases. The highest time on market is observed for leaseholds of 150-300 years remaining, followed by the groups 125–149, 80–99, 100–124, and 700+. Differences in liquidity are therefore unlikely to explain our results.

IV.F. Financing Frictions

Financing frictions have the potential to affect the relative valuation of leaseholds and freeholds. Leaseholds, in particular short dated ones, require lower upfront payments to take ownership of a property. If households have high future income that they cannot borrow against, these shorter leaseholds are more attractive than longer leaseholds or freeholds. This credit constraint makes shorter leaseholds more desirable, increasing their valuation relative to a frictionless benchmark.

On the other hand, short-maturity leaseholds are harder to finance than long-maturity ones. For example, U.K. mortgage lenders typically require a 30-year unexpired lease term to remain at the end of the mortgage (Council of Mortgage Lenders, 2013). This means that leasehold purchases have to be financed with shorter maturity mortgages once the lease length falls below 55 or 60 years. The loss in "collateral value" for these leaseholds could contribute to the large estimated discounts for leaseholds with maturities around 80 years. However, in Online Appendix A.4 we calibrate a version of the simple valuation model from the introduction to show that even under conservative assumptions for the collateral value of the house, financing frictions cannot explain discounts for leases of long maturities. Intuitively, a lease with 200 years remaining will only incur direct losses to its collateral value in 140 years, when the lease will have 60 years left: the loss of a fraction of the total value so far in the future has minimal effect on the present value of the leasehold.

IV.G. Taxation and the Stamp Duty

Purchases of property in the United Kingdom are subject to a transaction tax (stamp duty). The tax applies equally to freehold or leasehold purchases. The tax schedule is progressive: for example, a purchase of a property up to £125,000 is tax exempt, whereas a purchase of a property between £125,001 and £250,000 is taxed at 1% of the total purchase price (see Online Appendix A.1.3). This tax schedule potentially makes shorter term leases more attractive because for similar properties a shorter and thus cheaper leasehold might avoid incurring the higher tax bracket. This would bias against finding significant leasehold discounts. However, only buyers of properties that are very close to the boundaries of the tax bracket would be affected, and since the brackets are relatively large the effect on the average discount is unlikely to be quantitatively important.

IV.H. Hold-Up at Lease Extension

One other friction that might contribute to explaining our results relates to hold-up problems during lease extensions, whereby a freeholder might charge unreasonably high premiums and administrative expenses for the purchase of extra years on the lease. These concerns could potentially make leasehold contracts less attractive and therefore contribute to explaining the leasehold price discounts that we estimate in the data.

In recent years, however, U.K. legislation and court practice have systematically alleviated this concern. Legislation passed in 1993, well before the beginning of our sample, has granted virtually all leaseholders who have lived in a property for more than 2 years the statutory right to seek a lease extension by 90 years in return for paying a premium (see Online Appendix A.1.5). If a reasonable premium cannot be negotiated with the freeholder, the leaseholder can refer the matter to the LVT, which will establish the payable premium. Badarinza and Ramadorai (2014) recently argued that court-enforced settlements have occurred at estimated discounts favorable to leaseholders compared to those estimated based on market values in this article. Although the court decisions are somewhat infrequent and subject to legal and advisory costs that can run in the "tens of thousands of pounds" (Westminster City Council 2013), they alleviate the concern that our discounts could simply be due to the hold-up problem.[15]

In addition, our data provide direct evidence that hold-up frictions cannot explain our estimated price discounts. If some freeholders were more prone to hold leaseholders up at lease extension, thus resulting in the related leaseholds trading both for lower prices and on average with lower remaining lease length, we would expect the price differences between leaseholds of different maturity to decrease significantly when estimated within

15. The possibility of favorable tribunal decisions, and potentially cheaper lease extensions outside the court system as an indirect effect, would increase the ex ante valuation of leaseholds if prospective buyers were to anticipate lower future costs of extensions. To the extent that buyers take this into account, this mechanism would generate a bias against finding large discounts for leaseholds relative to freeholds. There are several reasons buyers may discount the ex ante value of this potential advantage: transaction costs related to the extension process can be significant, bargaining times are long (6–18 months), there is uncertainty about the outcome (which may involve going to court twice), and there is general low awareness by buyers about the details of the extension process. See Online Appendix A.1.5 for more details.

the same building because leaseholds in the same building have the same freeholder. In Section IV.C we showed that leasehold discounts remain identical when estimating them within the same building.

V. Discussion and Interpretation

Section III presented new facts about the relative pricing of freeholds and leaseholds of different maturities. Leaseholds with over 700 years of maturity trade at the same price as freeholds for otherwise identical properties. For leaseholds with shorter maturities the price discounts range from 10–15% at 80–100 years remaining to 5–8% at 125–150 years. This suggests that a significant fraction of the value of freehold properties comes from cash flows (rents) more than 100 years in the future.

In this section we introduce a simple pricing model to discuss the forces that drive these estimated price discounts. Following the classic valuation model of Gordon (1982), we assume that cash flows arising in each future period are discounted at a constant rate r, so that the T-period discount factor is e^{rT}. We also assume that rents are expected to grow at a constant rate g, so that expected rents follow: $E_t[D_{t+s}] = D_t e^{gs}$.[16] In this model, a claim to the rents for T periods, the T-maturity leasehold, is valued at

$$P_t^T = \int_t^{t+T} e^{-r(s-t)} D_t e^{g(s-t)} ds = \frac{D_t}{r-g}(1 - e^{-(r-g)T}). \tag{3}$$

Correspondingly, the infinite maturity claim, the freehold, is valued at: $P_t = \lim_{T\to\infty} P_t^T = \frac{D_t}{r-g}$, with $r > g$. The price discount for a T-maturity leasehold with respect to the freehold is:

$$Disc_t^T \equiv \frac{P_t^T}{P_t} - 1 = -e^{-(r-g)T}. \tag{4}$$

For any given maturity, the price discount decreases (in absolute value) the higher the discount rate r and the lower the growth rate of rents g. The first effect occurs because higher discount rates reduce the present value of future rents. The second effect occurs

16. Technically, g is the sum of the expected growth rate of rents and a Jensen inequality term. The Jensen term is very small given the variance of rent growth and in the interest of intuitive results, we ignore this term and refer to g as the expected growth rate of rents.

because higher growth rates of rents increase the residual value of the property after leasehold expiry.

This simple model shows that the value of leasehold contracts is increasing with maturity. Since we observe transactions of leaseholds with different maturities in the data, the housing markets we consider represent an equilibrium where the marginal freeholders and leaseholders are indifferent between the various contracts.

The estimated price differences between very long-run leaseholds and freeholds can be matched by a calibration of equation (4) with a net discount rate of $r - g = 1.9\%$. Figure VI visually confirms the good fit of this parameterization. To recover the implied gross discount rate, we need to separately estimate the average long-run real growth rate of rents.

Using data from rental indexes, we estimate the real long-run growth rate of rents, g, to be 0.62% and 0.17% for the United Kingdom and Singapore, respectively (see Online Appendix A.5 for all the data sources).[17] To be conservative, we set $g = 0.7\%$. These estimates of low long-run real growth rates of rents are consistent with Shiller (2006), who estimates long-run real house price growth rates to be very low, often below 1%.[18]

To sum up, our estimated price discounts between long-run leaseholds and freeholds, combined with the estimate of the real growth rate of rents of 0.7%, suggest that households apply

17. To verify our methodology and compare it to existing literature, we also estimate the average growth rate of real rental income in the United States, which we find to be 0.53%. Our estimates are in line with the median growth rate of 0.4% estimated in Campbell et al. (2009).

18. That the two long-run growth rates have to be identical is a feature necessary for rental yields to be stationary. Eichholtz (1997), Eitrheim and Erlandsen (2005), and Ambrose, Eichholtz, and Lindenthal (2013) also confirm Shiller's observation of negligible long-run real house price growth in different countries and using different data. Ambrose, Eichholtz, and Lindenthal (2013) find evidence of cointegration between house prices and rents using very long-run housing data for Amsterdam. Despite the historical evidence, one possibility is that agents might expect higher rent growth to occur in the future for a substantial period of time; in this case a higher expected growth rate g would increase the discount rate r necessary to match our estimates. For example, one might conjecture that super-star cities like Singapore or London might experience such high rent growth in the future (Gyourko, Mayer, and Sinai 2013). However, the low growth rate of rents were estimated in a period when London and Singapore were already major capitals of the world. In addition, our estimates in Table III are very similar outside of London, where it is even less likely that households are expecting major rental growth for centuries.

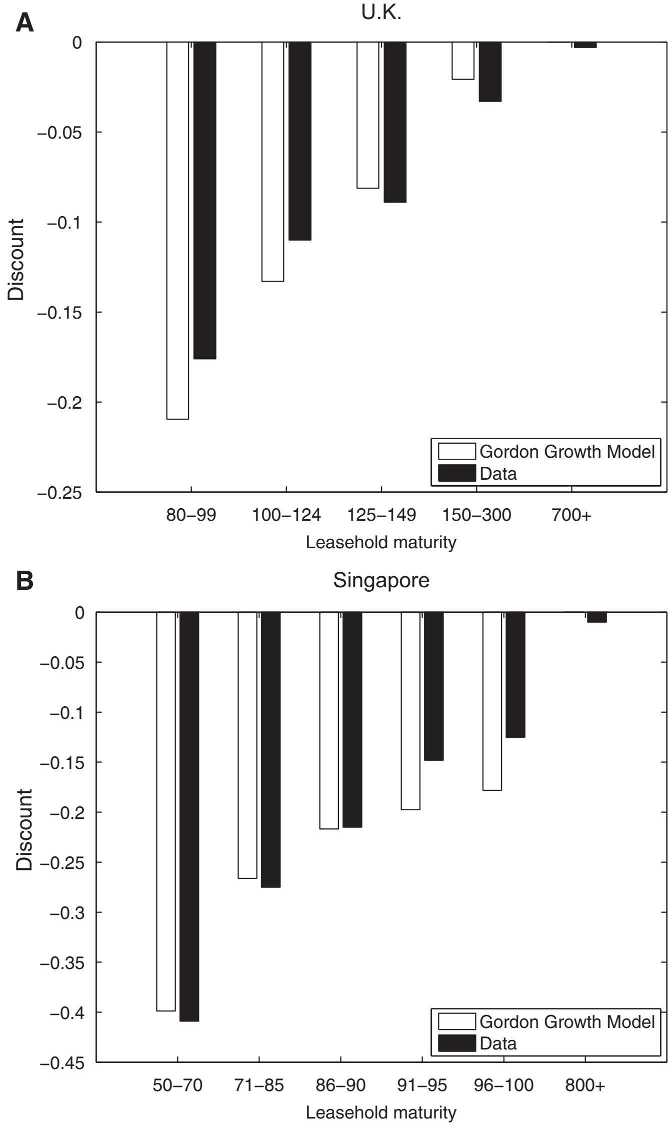


FIGURE VI

Gordon Growth Model and Estimated Discounts

Figure shows the discounts for leaseholds observed in the U.K. (Panel A) and Singapore (Panel B) together with discounts predicted by a Gordon growth model with $r - g = 1.9\%$.

discount rates of 2.6% to housing cash flows hundreds of years in the future.

V.A. Very Long-Run Risk-Free Rate and Risk Premia

Our estimated discount rate of 2.6% is appropriate for all cash flows that have the same riskiness as housing at long horizons. Estimating the stochastic process of rents and housing is difficult because of the unavailability of long, unsmoothed time series. However, despite this difficulty, our results can be used to derive useful bounds on the long-run risk-free discount rate and housing risk premium under mild assumptions.

The long-run discount rate r is composed of a risk-free component r^f, corresponding to the yield of a hypothetical very long-run real zero-coupon bond, and a risk premium r^{RP}, so that: $r = r^f + r^{RP}$. Online Appendix A.6 formalizes this decomposition. Under the assumptions that housing is risky at long horizons, so that the housing risk premium is positive, and that the risk-free rate is positive, it immediately follows that $0 < r^{RP} < 2.6\%$, and $0 < r^f < 2.6\%$.[19] Both the long-run risk-free discount rate and the housing risk premium are between 0 and 2.6% and sum to 2.6%. The assumption that housing is risky is supported by a recent literature that points to a co-movement between house prices and real economic activity (see, for example, Lustig and Nieuwerburgh 2005; Claessens, Kose, and Terrones 2009; Reinhart and Rogoff 2009; Favilukis, Ludvigson, and Nieuwerburgh 2010). Intuitively, housing performs poorly during some of the most adverse economic crises, wars, epidemics, and natural disasters. While it might be harder to quantify precisely how risky housing is, the assumption that it is at least riskier than a risk-free payment is mild. Similarly, many economists have argued for the economic and mathematical implausibility of zero or negative long-run risk-free discount rates (e.g., Koopmans 1965; Tirole 1982; Nordhaus 2007, 2013).

To confirm this bound for the long-run risk-free rate and provide an estimate of the long-run housing risk premium, we also consider the long-run risk-free rate in the United Kingdom obtained from the U.K. real yield curve. Figure VII shows that the real yield curve is flat for maturities between 1 and 25 years with

19. Long-run housing is risky if the price of claims to long-run housing cash flows covaries positively with risk factors such as consumption, and hence has low payoffs in bad states of the world.

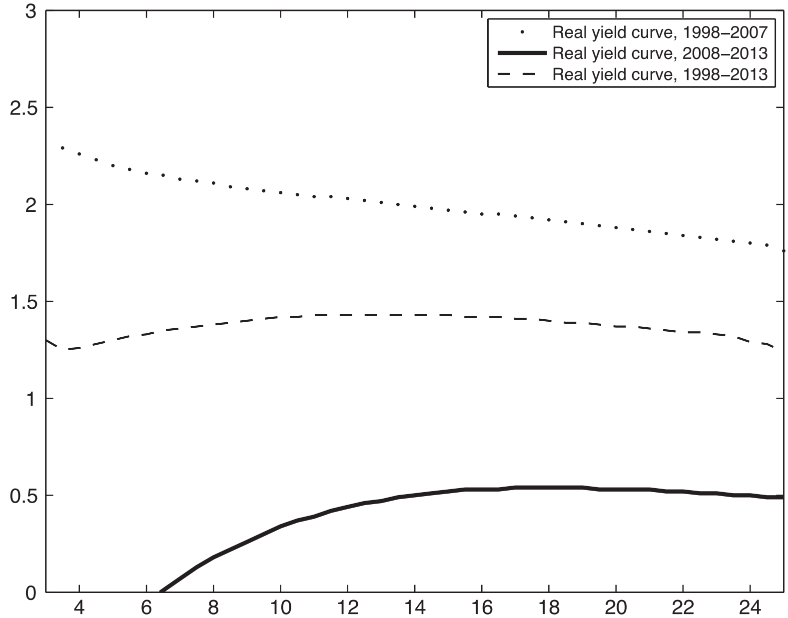


FIGURE VII

U.K. Gilts Real Yield Curve

The figure plots the real yield curve for U.K. gilts as computed by the Bank of England.

an average real yield of 1.4% for the period 1998–2013. The Bank of England also made available a 40-year real yield for the period 2006–2013; the average 40-year real yield during this period was 0.4%.[20] This latter estimate should be interpreted with caution not only because of liquidity concerns but also because the period is dominated by the global financial crisis and the European

20. The real yield curve is computed by the Bank of England and is available at http://www.bankofengland.co.uk/statistics/Pages/yieldcurve/archive.aspx. We are grateful to Zhuoshi Liu at the Bank of England for making the long-maturity average yield available to us. The U.K. government debt also includes some perpetual bonds: the War Loan and the Annuities. These bonds comprise a negligible part of the outstanding U.K. government debt (£2.6 billion out of £1.5 trillion of debt outstanding), and are classified as small and illiquid issuances by the U.K. Debt and Management Office. They are excluded from our analysis, not only because they are nominal and we only use data on U.K. real gilts, but also because their negligible size, scarce liquidity, and callability make it hard to interpret their prices in terms of discount rates.

sovereign debt crisis.[21] Nevertheless, we conclude that the U.K. real yield curve is approximately flat on average, with a real yield of 1.4% for maturities between 1 and 25 years, and that there is some evidence for a mild downward slope at longer maturities with an average 40 year yield below 1%.

Using a calibrated value of 1% for very long-run risk-free yields, we can decompose the total discount rate needed to match the estimated leasehold discounts into the risk-free component $r^f = 1\%$, and a risk adjustment of $r^{RP} = 1.6\%$. Although data from the United Kingdom real yield curve have to be interpreted cautiously given its shorter maturity and possible liquidity issues, we find the risk-free rate estimated with this independent source to be centered in the range of 0–2.6% obtained under our estimates and consistent with the assumption that rents are risky in the long run and carry a positive risk premium.

VI. Conclusions and Avenues for Future Research

We explore a unique feature of housing markets in the United Kingdom and Singapore to provide novel evidence on very long-run discount rates. We find these discount rates to be sufficiently low, at 2.6%, that more than 10% of the total value of a freehold property comes from cash flows that occur more than 100 years in the future. As such, our findings are of direct relevance for real estate economics and the ongoing effort to understand real estate prices (Flavin and Yamashita 2002; Lustig and Nieuwerburgh 2005; Piazzesi, Schneider, and Tuzel 2007; Favilukis, Ludvigson, and Nieuwerburgh 2010; Nathanson and Zwick 2012).

Our results are also informative for future research in asset pricing. Combining our discount rates with estimates of the average rate of return to housing is informative about the term structure of discount rates for housing cash flows. A recent literature has provided a wide range of estimates for the real rate of return to housing, ranging from above 6% (Flavin and Yamashita

21. Figure VII plots the average shape of the real U.K. gilts curve for the period 1998–2013, as well as for two subperiods: 1998–2007 and 2008–2013. The level of the yield curve shifted down during this latter period and the yield curve became hump-shaped.

2002; Favilukis, Ludvigson, and Nieuwerburgh 2010) to 2.5% (Piazzesi, Schneider, and Tuzel 2007). Together with our estimates of very long-run discount rates for housing cash flows, high estimates for the average return suggest a downward-sloping term structure of discount rates, while lower estimates are consistent with a flat term structure.

By specifying the dynamic process of housing cash flows, our estimate of long-run risk premiums could also be decomposed into the asset-specific quantity of risk and the economy-wide price of long-run risk. Future work based on our findings could thus provide a new testing ground for asset pricing theories, which make stark predictions for the evolution of the price of risk over long horizons (see Campbell and Cochrane 1999; Bansal and Yaron 2004; Barro 2006; Barro and Ursua 2008; Nakamura et al. 2013; Gabaix 2012, 2015; Binsbergen, Brandt, and Koijen 2012; Binsbergen et al. 2013, Belo, Collin-Dufresne, and Goldstein 2015; Andries, Eisenbach, and Schmalz 2014).

Our results also have the potential to contribute to the conduct of cost-benefit analyses that consider effects that materialize over very long horizons. For example, the literature on climate change economics has focused on the importance of long-run discount rates in assessing the benefits of policies such as reducing carbon emissions (Weitzman 1998, 2007; Gollier 2006; Nordhaus, 2007; Stern 2007; Barro 2013). The debate in this literature has often focused on the seemingly puzzling model-implied feature that agents attach very little value to cash flows far into the future; our results, instead, are supportive of the notion that households do attach significant present value to cash flows in the very long-run.

Finally, the institutional context we present in this article lends itself to a direct test for the existence of infinitely lived rational bubbles, which we pursue in a related paper (Giglio, Maggiori, and Stroebel 2014).

University of Chicago, Booth School of Business
Harvard University
New York University, Stern School of Business

SUPPLEMENTARY MATERIAL

An Online Appendix for this article can be found at QJE online (qje.oxfordjournals.org).

References

Alvarez, Fernando, and Urban J. Jermann, “Using Asset Prices to Measure the Persistence of the Marginal Utility of Wealth,” *Econometrica*, 73 (2005), 1977–2016.

Ambrose, Brent W., Piet Eichholtz, and Thies Lindenthal, “House Prices and Fundamentals: 355 Years of Evidence,” *Journal of Money, Credit and Banking*, 45 (2013), 477–491.

Andries, Marianne, Thomas M. Eisenbach, and Martin C. Schmalz, “Asset Pricing with Horizon-dependent Risk Aversion,” *Federal Reserve Bank of New York Staff Reports*, 703 (2014).

Arrow, K. J., W. Cline, K. G. Maler, M. Munasinghe, R. Squitieri, and J. Stiglitz, “Intertemporal Equity, Discounting, and Economic Efficiency,” in *Climate Change 1995: Economic and Social Dimensions of Climate Change*, J. P. Bruce, H. Lee, and E. F. Haites, eds. (New York: Cambridge University Press, 1996).

Auerbach, Alan J., Jagadeesh Gokhale, and Laurence J. Kotlikoff, “Generational Accounting: A Meaningful Way to Evaluate Fiscal Policy,” *Journal of Economic Perspectives*, 8 (1994), 73–94.

Badarinza, Cristian, and Tarun Ramadorai, “Long-Run Discounting: Evidence from the UK Leasehold Valuation Tribunal,” University of Oxford Working Paper, 2014.

Bansal, Ravi, and Amir Yaron, “Risks for the Long Run: A Potential Resolution of Asset Pricing Puzzles,” *Journal of Finance*, 59 (2004), 1481–1509.

Barro, Robert J., “Rare Disasters and Asset Markets in the Twentieth Century,” *Quarterly Journal of Economics*, 121 (2006), 823–866.

———, “Environmental Protection, Rare Disasters, and Discount Rates,” NBER Working Paper 19258, 2013.

Barro, Robert J., and Jose F. Ursua, “Macroeconomic Crises since 1870,” *Brookings Papers on Economic Activity*, 39 (2008), 255–350.

Belo, F., P. Collin-Dufresne, and R. Goldstein, “Dividend Dynamics and Term Structure of Dividend Strips,” *Journal of Finance* (2015) Forthcoming.

Binsbergen, Jules van, Michael Brandt, and Ralph Koijen, “On the Timing and Pricing of Dividends,” *American Economic Review*, 102 (2012), 1596–1618.

Binsbergen, Jules van, Wouter Hueskes, Ralph Koijen, and Evert Vrugt, “Equity Yields,” *Journal of Financial Economics*, 110 (2013), 503–519.

Bracke, P., T. Pinchbeck, and J. Wyatt, “The Time Value of Housing: Historical Evidence from London Residential Leases,” SERC Discussion Paper 0168, 2014.

Burn, Edward Hector, John Cartwright, and Geoffrey C. Cheshire *Cheshire and Burn’s Modern Law of Real Property* (Oxford: Oxford University Press, 2011).

Campbell, John Y., and John H. Cochrane, “By Force of Habit: A Consumption Based Explanation of Aggregate Stock Market Behavior,” *Journal of Political Economy*, 107 (1999), 205–251.

Campbell, Sean D., Morris A. Davis, Joshua Gallin, and Robert F. Martin, “What Moves Housing Markets: A Variance Decomposition of the Rent–Price Ratio,” *Journal of Urban Economics*, 66 (2009), 90–102.

Capozza, Dennis R., and Gordon A. Sick, “Valuing Long-term Leases: The Option to Redevelop,” *Journal of Real Estate Finance and Economics*, 4 (1991), 209–223.

Claessens, Stijn, M. Ayhan Kose, and Marco E. Terrones, “What Happens During Recessions, Crunches and Busts?,” *Economic Policy*, 24 (2009), 653–700.

Council of Mortgage Lenders, “What Minimum Unexpired Lease Term Does the Lender Accept?,” *CML Lender’s Handbook* (London: CML, 2013).

Dickens, William, and Lawrence F. Katz, “Interindustry Wage Differences and Industry Characteristics,” NBER Working Paper No. 2014, 1987.

Eichholtz, Piet, “A Long Run House Price Index: The Herengracht Index, 1628–1973,” *Real Estate Economics*, 25 (1997), 175–192.

Eitrheim, Øyvind, and Solveig K. Erlandsen, “House Prices in Norway, 1819–1989,” *Scandinavian Economic History Review*, 53 (2005), 7–33.

Farmer, J. Doyne, John Geanakoplos, Jaume Masoliver, Miquel Montero, and Josep Perelló, "Discounting the Distant Future," Cowles Foundation Discussion Paper No. 1951, 2014.
Favilukis, Jack, Sydney C. Ludvigson, and Stijn Van Nieuwerburgh, "The Macroeconomic Effects of Housing Wealth, Housing Finance, and Limited Risk-Sharing in General Equilibrium," NBER Working Paper No. 15988, 2010.
Flavin, Marjorie, and Takashi Yamashita, "Owner-Occupied Housing and the Composition of the Household Portfolio," *American Economic Review*, 92 (2002), 345–362.
Gabaix, Xavier, "Variable Rare Disasters: An Exactly Solved Framework for Ten Puzzles in Macro-finance," *Quarterly Journal of Economics*, 127 (2012), 645–700.
Gabaix, X, "Disasters and Term Structure of Risky Assets," NYU Working Paper (2015).
Gautier, Pieter A., and Aico van Vuuren, "The Estimation of Present Bias and Time Preferences Using Land-lease Contracts," CESifo Working Paper Series No. 3549, 2014.
Genesove, David, and Lu Han, "Search and Matching in the Housing Market," *Journal of Urban Economics*, 72 (2012), 31–45.
Giglio, Stefano, Matteo Maggiori, and Johannes Stroebel, "No-Bubble Condition: Model-Free Tests in Housing Markets," NBER Working Paper No. 20133, 2014.
Gollier, Christian, "An Evaluation of Stern's Report on the Economics of Climate Change," IDEI Working Papers 464, 2006.
———, "Evaluation of Long-Dated Investments under Uncertain Growth Trend, Volatility and Catastrophes," TSE Working Paper 12-361, 2012.
Gordon, Myron J. *The Investment, Financing, and Valuation of the Corporation* (Westport, CT: Greenwood Press, 1982).
Gyourko, Joseph, Christopher Mayer, and Todd Sinai, "Superstar Cities," *American Economic Journal: Economic Policy*, 5 (2013), 167–199.
Hall, Robert E., "High Discounts and High Unemployment," NBER Working Paper No. 19871, 2014.
Iwata, Shinichiro, and Hisaki Yamaga, "Land Tenure Security and Home Maintenance: Evidence from Japan," *Land Economics*, 85 (2009), 429–441.
Koopmans, Tjalling C., "On the Concept of Optimal Economic Growth," *Pontificae Academiae Scientiarum Scripta Varia*, 28 (1965), 225–300.
Lustig, Hanno N., and Stijn Van Nieuwerburgh, "Housing Collateral, Consumption Insurance, and Risk Premia: An Empirical Perspective," *Journal of Finance*, 60 (2005), 1167–1219.
Nakamura, Emi, Jón Steinsson, Robert Barro, and José Ursúa, "Crises and Recoveries in an Empirical Model of Consumption Disasters," *American Economic Journal: Macroeconomics*, 5 (2013), 35–74.
Nathanson, Charles G., and Eric Zwick, "Arrested Development: A Theory of Supply-Side Speculation in the Housing Market," Working Paper, 2012.
Nordhaus, William D., "A Review of the Stern Review on the Economics of Climate Change," *Journal of Economic Literature*, 45, 686–702.
———, *The Climate Casino: Risk, Uncertainty, and Economics for a Warming World* (New Haven, CT: Yale University Press, 2013).
Petersen, Mitchell A., "Estimating Standard Errors in Finance Panel Data Sets: Comparing Approaches," *Review of Financial Studies*, 22 (2009), 435.
Piazzesi, Monika, Martin Schneider, and Johannes Stroebel, "Segmented Housing Search," Stanford/NYU working paper, 2013.
Piazzesi, Monika, Martin Schneider, and Selale Tuzel, "Housing, Consumption and Asset Pricing," *Journal of Financial Economics*, 83 (2007), 531–569.
Pindyck, Robert, "Climate Change Policy: What Do the Models Tell Us?," *Journal of Economic Literature*, 51 (2013), 860–872.
Reinhart, Carmen M., and Kenneth Rogoff, "The Aftermath of Financial Crises," *American Economic Review*, 99 (2009), 466–472.
Rosen, Sherwin, "Hedonic Prices and Implicit Markets: Product Differentiation in Pure Competition," *Journal of Political Economy*, 82 (1974), 34–55.

Shiller, Robert J., “Long-Term Perspectives on the Current Boom in Home Prices,” *Economists’ Voice*, 3, no. 4 (2006), 1–11.
Stern, Nicholas, *The Economics of Climate Change: The Stern Review* (Cambridge: Cambridge University Press, 2007).
Tirole, Jean, “On the Possibility of Speculation under Rational Expectations,” *Econometrica*, 50 (1982), 1163–1181.
Tyvimaa, Tanja, Karen M. Gibler, and Velma Zahirovic-Herbert, “The Effect of Ground Leases on House Prices in Helsinki,” *Journal of Housing and the Built Environment* (2014), 1–20. doi: 10.1007/s10901-014-9424-3.
Weitzman, Martin L., “Why the Far-Distant Future Should Be Discounted at Its Lowest Possible Rate,” *Journal of Environmental Economics and Management*, 36 (1998), 201–208.
———, “Gamma Discounting,” *American Economic Review*, 91 (2001), 260–271.
———, “A Review of the Stern Review on the Economics of Climate Change,” *Journal of Economic Literature*, 45 (2007), 703–724.
———, “Tail-Hedge Discounting and the Social Cost of Carbon,” *Journal of Economic Literature*, 51 (2013), 873–882.
Westminster City Council, “Homing in on the Issues: Perspectives on the Housing Crisis,” 2013.
Wong, Siu Kei, K. W. Chau, C. Y. Yiu, and M. K. W. Yu, “Intergenerational Discounting: A Case from Hong Kong,” *Habitat International*, 32 (2008), 283–292.

This page intentionally left blank