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Hot and cold seasons in the housing markets

Working paper

Hot and Cold Seasons in the Housing Markets*

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Abstract

In the U.K., every year during the second and third quarters (the "hot season"), regional housing markets experience sharp above-trend increases in (quality-adjusted) prices and in the number of transactions. During the fourth and first quarters (the "cold season"), housing prices and the number of transactions fall below trend. A similar seasonal cycle for transactions is observed in other developed countries. Housing prices, however, do not necessarily follow a seasonal pattern in all of them; in particular, in the U.S., while transactions are highly seasonal, prices display no seasonality. We discuss why the traditional asset-pricing approach to the housing market fails at explaining seasonal booms and busts and present a search model that can quantitatively mimic the seasonal fluctuations in transactions and prices in both the U.K. and the U.S. The model features a "thick-market" externality that can generate substantial differences in the number of transactions across seasons. The existence and extent of seasonality in prices depend on the distribution of market power between buyers and sellers. As a by-product, the model sheds new light on the mechanisms governing fluctuations in housing markets and it can be adapted to study lower-frequency movements.

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1 Introduction

A rich empirical and theoretical literature has been motivated by dramatic boom-to-bust episodes in regional and national housing markets.[1] Booms are typically defined as times when prices rise and there is intense trading activity, whereas busts are times when prices and trading activity fall below trend.

While the boom-to-bust episodes motivating the extant work are relatively infrequent and of unpredictable timing, this paper shows that in several housing markets, booms and busts are just as frequent and predictable as the seasons. In particular, in all regions of the U.K., as well as other continental European countries, every year a housing boom of considerable magnitude takes place in the second and third quarters of the calendar year (the "hot season"), followed by a bust in the fourth and first quarters (the "cold season"). In other countries, including the U.S., transactions display a strong seasonal pattern, while prices display no seasonality. The first contribution of this paper is to document the existence, quantitative importance, and cross-country variation of these seasonal booms and busts.

The surprising size and predictability of seasonal fluctuations in housing prices in some countries poses a challenge to standard models of durable-good markets. In those models, anticipated changes in prices cannot be large: If prices are expected to be much higher in May than in December, then buyers will shift their purchases to the end of the year, narrowing down the seasonal price differential. More formally, in the absence of risk, the asset-market equilibrium condition states that the one-period rental value of a house plus its appreciation should equal the one-period gross cost of housing services.[2] Calling p_t and d_t the real price of housing and rental services, respectively, and assuming that the gross real service cost is a (potentially changing) proportion c_t of the property price, the equilibrium asset-market condition is:

$$d_{t+1} + (p_{t+1} - p_t) = c_t \cdot p_t \tag{1}$$

where c_t is the sum of the (potentially time-varying) depreciation rate, maintenance and repair expenditure rate, property tax rate, and the tax-adjusted interest rate.[3] The arbitrage condition

[1] See for example Stein (1995), Krainer (2001), Ortalo-Magne and Rady (2005) and the contributions cited therein.

[2] For an early asset-market approach to the housing market, see Poterba (1984).

[3] The effective interest rate is a weighted average between mortgage interest rate plus the opportunity cost of housing equity, where the weights are given by the loan-to-value ratio.

thus states that the seasonals in real prices must be accompanied by seasonals in the cost of housing services c_t or in the rental service flow d_t. Rents, however, display no seasonality, implying a substantial and, as we shall argue, unrealistic degree of seasonality in service costs c_t. For example, the price seasonality observed in the U.K. implies that service costs should be roughly 300 percent higher in the cold season than in the hot season. This seems unlikely, particularly because interest rates and tax rates, two major components of c_t, display no seasonality.[4]

We investigate a number of possible explanations for the seasonal booms and busts. The seasonal in housing markets does not seem to be driven by seasonal differences in liquidity related to overall income. Income is typically high in the last quarter, a period in which housing prices and the volume of transactions tend to fall below trend.[5] At any rate, all these variables are predictable, and in an informationally efficient market, their effect should be incorporated in prices so that future price changes are unforecastable. Indeed, the predictable nature of housing prices fluctuations is confirmed by U.K. estate agents, who in conversations with the authors observed that during winter months there is less activity and owners tend to sell at a discount. And, perhaps more compelling, publishers of house price indexes go to great lengths to produce seasonally adjusted versions of their indexes, usually the index that is published in the media. As stated by the publishers:

"House prices are higher at certain times of the year irrespective of the overall trend. This tends to be in spring and summer, when more buyers are in the market and hence sellers do not need to discount prices so heavily, in order to achieve a sale." and *"...we seasonally adjust our prices because the time of year has some influence. Winter months tend to see weaker price rises and spring/summer see higher increases all other things being equal."* (From Nationwide House Price Index Methodology.)

"Houses prices are seasonal with prices varying during the course of the year irrespective of the underlying trend in price movements. For example, prices tend to be higher in the spring and summer months when more people are looking to buy." (From Halifax Price Index Methodology.)

The seasonal behavior of housing markets and the failure of *a priori* appealing explanations,

[4]The implied seasonality of service costs is even higher, and hence even more implausible in other countries.

[5]Beaulieu and Miron (1992) and Beaulieu, Miron, and MacKie-Mason (1992) show that in most countries, including the U.K., income peaks in the fourth quarter of the calendar year. There is also a seasonal peak in output in the second quarter, and seasonal recessions in the first and third quarters. Housing price seasonality is not in line with income seasonality: prices are above trend in the second and third quarters.

thus poses a new puzzle to the standard asset-pricing approach. This paper tries to resolve the puzzle by resorting to a search-theoretic approach.

Specifically, we develop a search model in which housing prices and the volume of transactions in each season (or semester) are derived from the maximizing behavior of both buyers and sellers. At the beginning of each season a house can be in one of two states: "matched," when it delivers a positive housing service flow to its owner, or "on sale," when it does no longer yield housing services to its owner. Each match has a probability of breaking, in which case the house goes "on sale." Agents who own houses where the match is broken seek to sell ("sellers") and agents who are *not* matched to a house seek to buy ("buyers"). Buyers and sellers are randomly matched. Upon visiting a house, the buyer draws an idiosyncratic matching quality reflecting the utility services the house will yield while matched; this match quality is only observed by the buyer. The potential buyer searches until she finds a house whose utility services, net of price, are above the option value of keeping searching.

In the model, a slightly lower *ex-ante* probability of moving houses in a given season (e.g., because of changes in school, household size, jobs, etc.) can trigger a "thick-market" externality that makes it appealing for a larger number of agents to buy and sell during that season. This is because in the model, buyers are more likely to find a better-quality match (and hence their willingness to pay increases) when there are more houses on sale. Hence, in a thick market (the hot season), the volume of transactions goes up. Whether or not prices also go up depends on the distribution of market power between buyers and sellers. Because the quality of matches and hence buyers' willingness to pay increase in a thick market, when sellers "set prices" (that is, sellers have monopoly power in the transaction) they can extract all the surplus of the transaction (buyers' higher willingness to pay) by charging higher prices in hot seasons. When instead buyers set prices, they get all the surplus of the transaction and prices are therefore insensitive to the season. The extent of price seasonality hence depends on the degree of market power of sellers and buyers.

We show that the calibrated model can account for most of the seasonal fluctuations in transactions in the U.K. and the U.S., and at the same time match the seasonality in prices in the U.K. together with the lack of seasonality in prices in the U.S.[6] The crucial distinction between the two economies in the calibrated model is that in the U.K. sellers have more power to set prices than

[6] Our focus on these two countries is largely driven by the reliability and quality of the data.

in the U.S. We argue that this distinction can be justified on the grounds that land per capita is substantially scarcer and building regulations more stringent in the U.K., two features that are in turn reflected in relatively higher housing price-to-income ratios in the U.K.

To summarize, the contribution of this paper is twofold. First, it documents seasonal booms and busts in housing markets and shows that the predictability and high extent of seasonality in prices observed in some of them cannot be quantitatively reconciled with the standard asset-pricing equilibrium condition embedded in most models of housing markets (or consumer durables, more generally). Second, it develops a search model that can quantitatively account for the empirical puzzle and shed new light on the mechanisms governing fluctuations in housing markets. As a by-product, the model can potentially be adapted to study lower frequency fluctuations.

The paper is organized as follows. Section 2 presents the empirical evidence and discusses different potential (though ultimately unsuccessful) explanations. Section 3 argues why, given the evidence, we need to deviate from the standard asset-pricing approach to housing markets. Section 4 presents the new model. Section 5 presents a quantitative analysis of the model and confronts it with the empirical evidence. Section 6 presents extensions of the baseline model; in particular, it studies the quantitative potential of transaction costs as alternative drivers of seasonality in housing markets. Section 7 presents concluding remarks.

2 Hot and Cold Seasons

This section documents the behavior of housing prices across what we refer to as the two main seasonal terms: the summer term (second and third quarters of the calendar year) and the winter term (first and fourth quarters) in different countries and regions within a country.

2.1 Data

In the analysis we shall pay particular attention to the housing-market records of the U.K. and the U.S., the countries for which the data are of highest quality. Below is a brief description of the data on prices and transactions in these two countries. A description of the data sets and sources for other countries studied in this Section is available in the Data Appendix.

U.K.

In the U.K. there are two main data sets providing quality-adjusted non-seasonally adjusted prices: the Halifax House Price Index, derived from the data collected by Halifax, one of the country's largest mortgage lenders, and the price index produced by the Office of the Deputy Prime Minister (ODPM).[7]

Halifax reports regional indexes on a quarterly basis for the 12 standard planning regions of the U.K., as well as for the U.K. as a whole. The indexes calculated are 'standardized' and represent the price of a typically transacted house. The standardization is based on hedonic regressions that control for a number of characteristics, including location, type of property (house, sub-classified according to whether it is detached, semi-detached or terraced, bungalow, flat), age of the property, tenure (freehold, leasehold, feudal), number of rooms (habitable rooms, bedrooms, living-rooms, bathrooms), number of separate toilets, central heating (none, full, partial), number of garages and garage spaces, garden, land area, and road charge liability. Accounting for these characteristics allows to control for the possibility of seasonal changes in the composition of the set of properties (for example, shifts in the location or sizes of properties). The index reports transaction prices based on mortgages to finance house purchase at the time the mortgage is approved; re-mortgages and further advances are excluded.

The ODPM index is based on the same method as is the Halifax index, and differs only in two respects. First, it collects information from a large sample of all mortgage lenders in the country.[8] And second, it reports the price at the time of completion, rather than approval. Completion might take on average three to four weeks after the agreement, due generally to paper-work delays. The ODPM index goes back to 1963, though only after 1993 does it include all mortgage lenders (before that time, prices are based on Building Societies reports).

To compute real price indexes, we later deflate the housing price indexes using the non-seasonally adjusted retail price index (RPI) including "All items except housing" provided by the U.K. Office for National Statistics.

As an indicator of the number of transactions, we use the number of mortgages advanced for home purchases; the data are collected by the ODPM through the Survey of Mortgage Lenders and are disaggregated by region.

[7] Other price indices, like Nationwide, report quality adjusted data but they are already seasonally adjusted. The Land Registry data reports average prices, without adjusting for quality.

[8] The Halifax index uses all the data from Halifax mortgages.

U.S.

The non-seasonally adjusted price index for the U.S. comes from the Office of Federal Housing Enterprise Oversight (OFHEO), which in turn builds its index from data provided by Fannie Mae and Freddie Mac, the biggest mortgage lenders; this is a repeat-sale index (and hence, barring depreciation, quality is kept constant). The index is calculated for the whole of the U.S. and also disaggregated by state (the 50 states and the District of Columbia) and by the 379 metropolitan statistical areas defined by OFHEO.

To compute real price indexes, we use the non-seasonally adjusted consumer price index (CPI) including "All items less shelter" provided by the U.S. Bureau of Labor Statistics.[9]

Data on the number of transactions come from the National Association of Realtors, and correspond to the number of sales of existing single-family homes. The data are disaggregated into the four major Census regions.

2.2 The Cross-Country Evidence

This Section briefly summarizes the cross-country evidence on seasonal fluctuations in housing prices and transactions. The extent of seasonality is summarized by means of country-by-country OLS regressions of the type:

$$g_{\tau_t} = \alpha_\tau + \beta_\tau S_t + \varepsilon_t \text{ and} \quad (2)$$

$$g_{p_t} = \alpha_p + \beta_p S_t + \varepsilon_t, \quad (3)$$

where g_{τ_t} is the annualized growth rate of the quarterly number of transactions, g_{p_t} is the annualized growth rate of the quarterly (quality-adjusted) house price index (expressed both in nominal and real terms), and S_t is a dummy variable that takes the value 1 if t corresponds to the second or third quarter of the calendar year, and 0 otherwise. The constant term α_τ (α_p) measures the average growth rate in the number of transactions (housing prices) during the period and β_τ (β_p) measures the average difference in growth rates between summers and winters. A statistically significant value for the β's rejects the null of no difference in growth rates across seasons. Table 1 and Table 2 report the slope coefficients and standard errors of the regressions for transactions (Table 1) and both nominal and real prices (Table 2).

[9] As it turns out, there is little seasonality in the U.S. CPI index, a finding first documented by Barsky and Miron (1989), and hence the seasonal patterns (or lack thereof) in nominal and real housing prices coincide.

Table 1 suggests a strong and positive "summer" effect in all countries for which non-seasonally adjusted data on housing transactions are available. Table 2 displays a uniform pattern of signs for housing prices, with countries in the northern hemisphere displaying a positive second-and-third quarter effect and countries in the southern hemisphere displaying a negative effect (note that the austral summer takes place in the fourth and first quarters and hence the negative signs in the southern hemisphere). However, the statistical and economic significance varies across countries. Belgium, France, and the U.K. display strongly significant summer effects in housing prices; Ireland, Sweden, and South Africa exhibit a less marked, though still significant summer effect; and finally, Denmark, Norway, the U.S., Australia, and New Zealand show no statistically significant summer effect.[10]

Table 1: Average Difference in the Annualized Growth Rate of the Number of Transactions between Second-Third Quarters and Fourth-First Quarters, by Country

Country	Coef.	Std. Error	Observations
Belgium	61.675**	(15.008)	51
Ireland	47.834**	(17.936)	120
Sweden	194.489**	(35.106)	75
United Kingdom	130.277**	(20.738)	124
United States	162.354**	(19.369)	149

Note: The Table shows the coefficients (and standard deviations) on the dummy variable S_t (second-third quarters) in the regressions $g_t=a+b\times S_t+e_t$, where g_t is the annualized growth rate of the number of transactions; a is a constant, omitted. The equations use quarterly data (see Appendix). Robust standard errors in parentheses. +Significant at the 10%; *significant at the 5%; **significant at 1%.

[10] While the time span differs across countries, a sensitivity analysis performed by the authors shows that the period covered does not significantly affect the extent of seasonality. Still, results should be read with the caveat that not all countries perform quality adjustments, as discussed in the data Appendix. This is why the paper focuses attention on the U.S. and the U.K., for which the data are quality-adjusted.

Table 2: Average Difference in Annualized Housing Price Growth (nominal and real) between Second-Third Quarters and Fourth-First Quarters, by Country

	Nominal price inflation		Real price inflation		
Northern Hemisphere	Coef.	Std. Error	Coef.	Std. Error	Observations
Belgium	14.447**	(1.507)	13.695**	(1.740)	95
Denmark	1.085	(2.074)	1.029	(2.072)	51
France	12.459**	(1.200)	12.198**	(1.220)	34
Ireland	6.076*	(2.934)	4.456	(2.999)	35
Netherlands	2.723	(1.537)	3.234	(1.701)	48
Norway	3.072	(3.333)	4.628	(3.224)	52
Sweden	4.504	(2.270)	5.484*	(2.187)	76
United Kingdom	8.233**	(2.325)	6.105*	(2.354)	91
United States	0.272	(0.772)	-0.692	(0.892)	120
Southern Hemisphere					
Australia	-1.163	(2.389)	-0.796	(2.415)	73
New Zealand	-1.516	(1.775)	-2.148	(1.808)	146
South Africa	-5.816*	(2.618)	-6.112	(3.129)	120

Note: The Table shows the coefficients (and standard deviations) on the dummy variable S_t (second-third quarters) in the regressions $g_t=a+b\times S_t+e_t$, where g_t is the annualized nominal or real house price growth, as indicated; a is a constant, omitted. The equations use quarterly data (see Appendix). Robust standard errors in parentheses. +Significant at the 10%; *significant at the 5%; **significant at 1%.

2.3 The Within-Country Regional Evidence

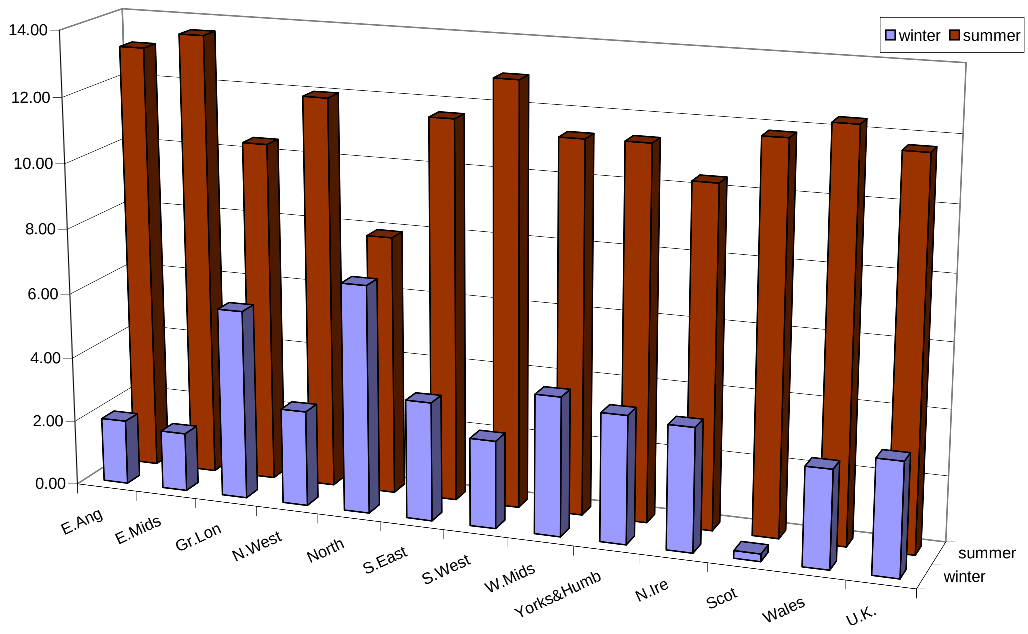
The size of countries (and hence the number of potential regional housing markets) varies substantially in the sample studied before. In particular, for large countries, it is in principle inappropriate to talk about a single national housing market. The finding of no seasonal patterns in prices at the aggregate level, for example, might mask different seasonal behaviors at more disaggregated levels. Conversely, the existence of a seasonal pattern in the aggregate might reflect some aggregation anomalies. It is hence of importance to study the behavior of prices (and transactions) at a more disaggregated level. we do so in this Section, starting with the U.K. and the U.S., the countries with highest-quality data; we also document the behavior of rentals and interest rates for these two countries. Finally, we describe the seasonal patterns for prices in different regions of Belgium and France.

Housing Market Seasonality in the U.K.

Nominal Housing Price Changes

Figure 1 reports the average annualized price growth rates in the summer term (second and third quarters) and the winter term (fourth and first quarters) over the period 1983 through 2005 using the Halifax index. As shown in the graph, the differences in price growth rates across seasons are generally very large and economically significant. During the period analyzed, the average price increases in the winter term were below 4 percent in all regions except for West Midlands (4.8 percent), Greater London (5.4 percent) and the North region (6.6 percent). In the summer term, the average growth rates were above 11 percent in all regions, except for the North (9 percent).

Figure 1: Average annualized housing price growth in summers and winters. Halifax Index 1983-2005.

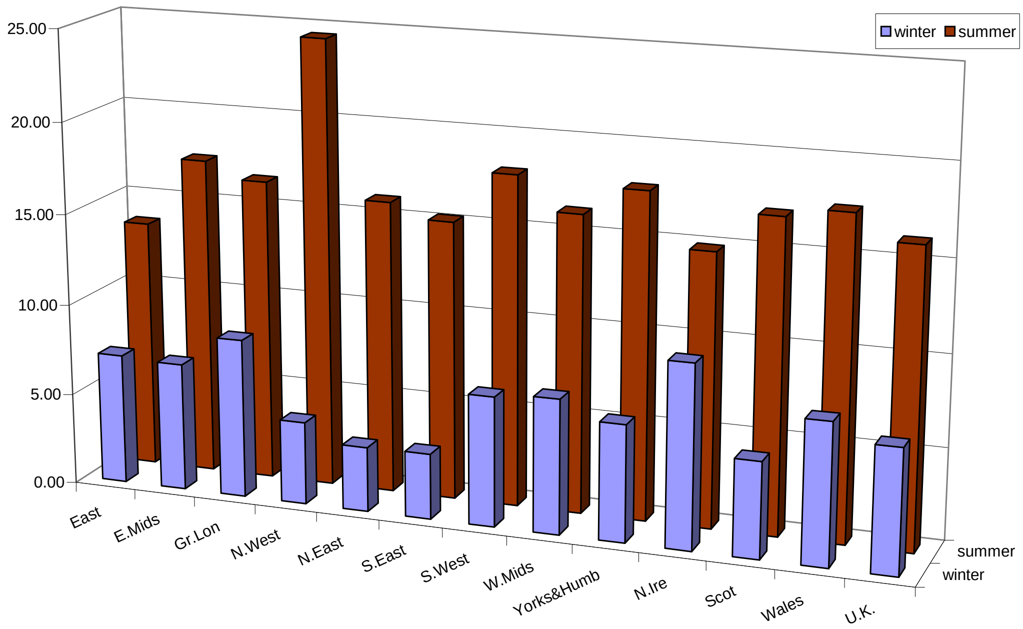


Note: Annualized (quality-adjusted) price growth rates in summers (second and third quarters) and winters (fourth and first quarters) in the U.K. and its regions. Halifax, 1983-2005.

Figure 2 shows the results from the ODPM index, starting in 1983 (for comparability with the Halifax Index). The annualized average price growth during the summer term is above 12 percent in all cases, whereas the increase during the winter term is systematically below 6 percent, except for Greater London and Northern Ireland. The qualitative patterns are hence similar to those obtained from Halifax; the relatively small quantitative differences between the two indexes might be explained by the lag between approval and completion.

As mentioned, the ODPM index goes back to 1968 for most regions. The average difference in growth rates between summers and winters during the longer period (not shown for the sake of brevity), are of the same order of magnitude, roughly above 8 percent.

Figure 2: Average annualized housing price growth in summers and winters. ODPM Index 1983-2005.



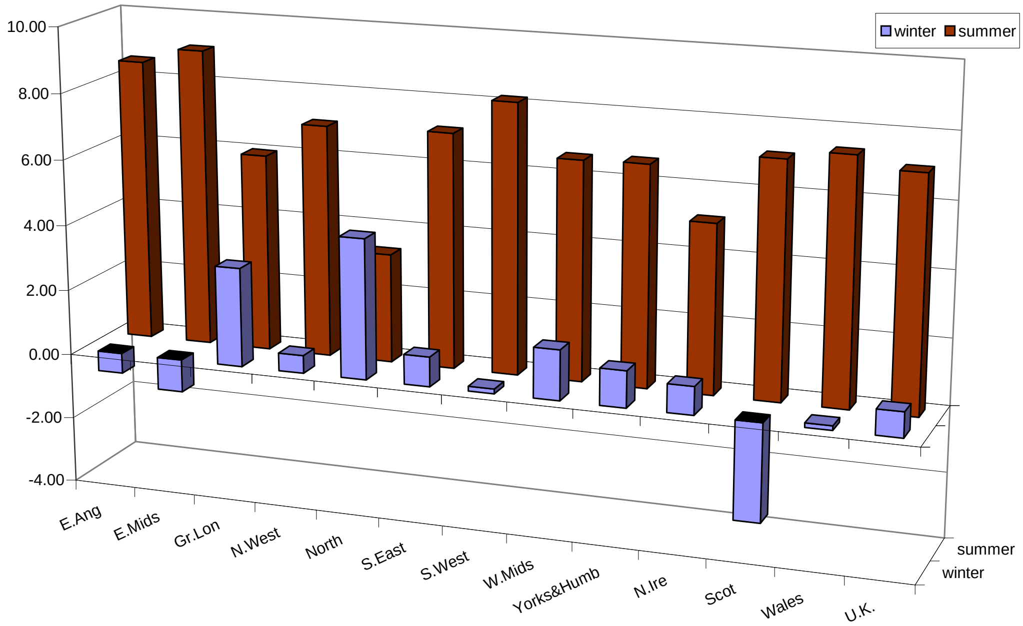
Note: Annualized (quality-adjusted) price growth rates in summers (second and third quarters) and winters (fourth and first quarters) in the U.K. and its regions. ODPM index 1983-2005.

Real Housing Price Changes

The previous Figures showed the seasonal pattern in nominal housing price inflation. The seasonal pattern of real housing prices (that is, housing prices relative to the overall non-seasonally-adjusted price index) depends of course on the seasonality of overall inflation. In the U.K. overall price inflation displays a slightly seasonal pattern. In particular, over the period 1983 through 2005, the average annualized non-seasonally-adjusted inflation rate in the summer term has been 4.7 percent, whereas the corresponding figure in the winter term has been 2.8 percent. The difference of 2 percent can hardly "undo" the differences of over 8 percent in nominal housing price inflation, implying a significant seasonal in real housing prices. This is illustrated in Figure 3. The graph is based on the Halifax index, but the results are similar for the ODPM index, not shown in the interest of space. Netting out the effect of overall inflation reduces the differences in growth rates between winters and summers to a country-wide average just above 6 percent.

We should note in addition that non-seasonally adjusted indexes of inflation are rarely used in practice (indeed it is even hard to find them), so they are unlikely to serve in contracts as financial means to "hedge" part of the seasonal nominal housing price fluctuations.

Figure 3: Average annualized real housing price growth in summers and winters. Halifax Index 1983-2005.

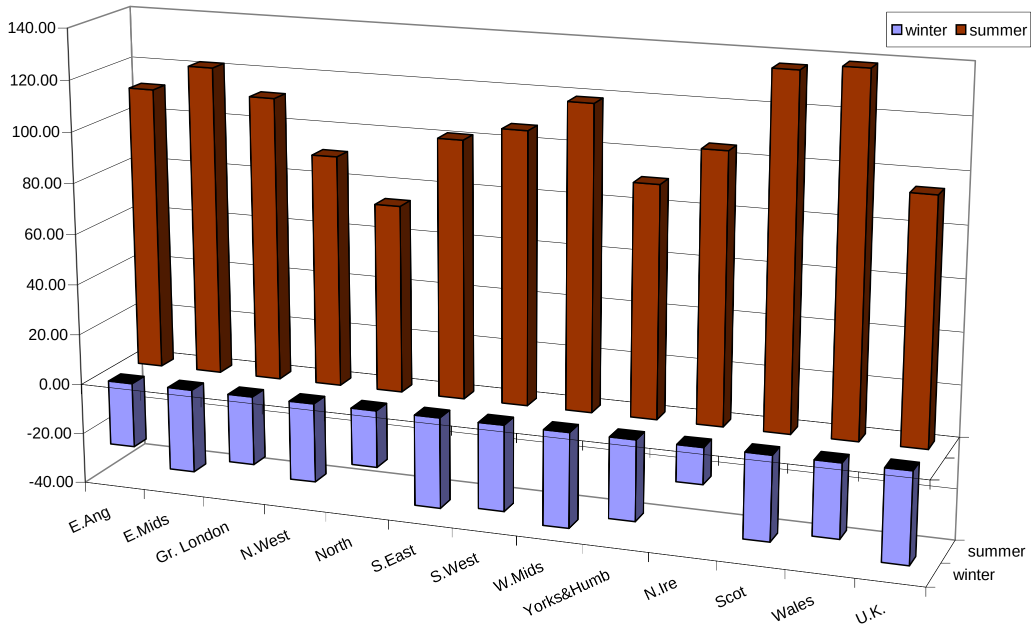


Note: Annualized (quality-adjusted) real price growth rates in summers (second and third quarters) and winters (fourth and first quarters) in the U.K. and its regions. Halifax, 1983-2005.

Number of Transactions

The seasonal differences in housing prices are mirrored by the patterns exhibited by the number of loans for housing purchases, which are a good proxy for the number of transactions. The data are collected by the National Survey of Mortgage Lenders and go back to 1974. For comparability with the price sample, Figure 4 shows the growth rate in the number of loans for mortgage completions in the U.K. from 1983 to 2005. (The 1974-2005 pattern is qualitatively and quantitatively similar to the one depicted in the Figure.) As Figure 4 shows, the number of transactions increases sharply during the summer term and declines in the winter term. Similar results are obtained by detrending the data using a linear trend (not shown).

Figure 4: Annualized growth rate of the number of loans in summers and winters.



Note: Annualized growth rate of the number of loan transactions in summers (second and third quarters) and winters (fourth and first quarters) in the U.K. and its regions, 1983-2005.

Statistical Significance of the Differences between Summers and Winters

This Section reports on the statistical significance of the results displayed in the previous Figures, as well as the characteristics of the houses and buyers involved in the transactions, by way of region-by-region OLS regressions similar to those used for the countries as a whole, as described in equations (2) and (3). The regressions are based on the Halifax data series, although similar results are obtained from the ODPM data (results are available on request). Table 3 summarizes the results. The first two columns show the coefficients and standard errors for the regressions based on prices for all houses and buyers. They show that the differences in housing price inflation are statistically significant at standard levels in all regions, except the North.

The following four columns show the corresponding figures for the prices of existing houses and new houses. The figures indicate that seasonal differences are mainly driven by the prices of existing houses, though new houses also display a fair amount of seasonality in some regions. In particular, new houses' inflation rates display a strong seasonal pattern in Greater London, Scotland, Northern Ireland and West Midlands. Note that, while economically sizeable, however, the seasonal differences are in many cases not statistically significant; one consideration that might explain the lower precision in the seasonal effects in new houses is that new houses represent a very small share of the market (due mostly to stringent construction restrictions), and hence the test on mean differences across seasons unavoidably displays lower significance levels. Another explanation for the difference might be differences in repair and maintenance costs across the two seasons. To the extent that repair costs are smaller in the summer (because good weather and the time of the owners are important inputs in construction), sellers will take this into account and post accordingly higher prices in the market. If differences in seasonal repair costs are behind the differences in prices, then, insofar as new houses need less repair and the potential buyers can ask the developers to tailor the final touches of the house to their needs, we should observe less seasonality in the prices of new houses than in those of existing houses. Though qualitatively possible, yet, the question remains as whether plausible differences in repair costs alone can quantitatively match the seasonal variation in the data, a point to which we come back later.

Table 3: Average Difference in Annualized Housing Price Inflation Between Summer and Winters, by Region and Type of House or Buyer

	All Houses (All buyers)		Existing houses (All buyers)		New houses (All buyers)		Former owner occupiers (All houses)		First-time buyer (All houses)	
	Coef.	Std. Error	Coef.	Std. Error	Coef.	Std. Error	Coef.	Std. Error	Coef.	Std. Error
E. Anglia	10.770**	(3.509)	10.028**	(3.727)	5.513	(6.878)	12.201**	(3.453)	5.663	(4.385)
E.Midlands	12.125**	(3.607)	12.905**	(3.651)	1.849	(5.814)	13.637**	(3.847)	9.496*	(3.699)
Gr. London	6.291*	(2.865)	6.624*	(2.898)	18.970*	(9.316)	5.357*	(2.658)	6.355*	(3.086)
N. West	8.629**	(2.813)	9.915**	(2.871)	-1.164	(7.051)	10.168**	(3.026)	5.675+	(2.950)
North	1.864	(3.224)	2.319	(3.333)	1.559	(5.606)	0.742	(3.295)	3.294	(3.897)
S. East	7.675**	(2.908)	8.061**	(2.889)	3.112	(4.066)	8.775**	(2.900)	4.301	(2.952)
S. West	10.961**	(3.439)	11.202**	(3.556)	8.004	(4.945)	11.895**	(3.549)	6.530+	(3.907)
W. Midlands	7.380+	(3.766)	7.126+	(3.799)	14.721+	(8.072)	8.160*	(3.965)	6.257+	(3.606)
Yorkshire&Humb	7.477*	(3.137)	8.249*	(3.194)	2.561	(6.449)	8.203*	(3.121)	7.340*	(3.506)
N. Ireland	9.253**	(3.425)	11.172**	(4.055)	10.977+	(6.082)	7.319	(4.524)	10.237*	(5.014)
Scotland	11.028**	(2.604)	13.627**	(2.895)	15.305*	(7.130)	12.591**	(2.673)	6.257*	(3.046)
Wales	9.332*	(3.721)	9.255*	(3.726)	1.146	(7.924)	9.943**	(3.729)	6.902+	(3.938)
U.K.	8.233**	(2.325)	8.896**	(2.364)	5.674*	(2.484)	9.114**	(2.348)	5.809**	(2.196)

Note: The Table shows the coefficients (and standard errors) on the dummy variable S_t (second and third quarters) in the regression g_t=a+b×$Summer_t$+e_t, where g_t is the annualized rate of *nominal* housing price inflation; a is a constant (omitted). The equations use quarterly data from 1983 to 2005. Robust standard errors in parentheses. +Significant at the 10%; *significant at the 5%; **significant at 1%.

The last four columns of Table 3 show the coefficients and standard deviations corresponding to the regressions based on prices paid by former-owner occupiers and first-time buyers. The distinction between former-owner occupiers and first-time buyers is interesting as some might a priori hypothesize that repeated buyers have more information on the seasonal patterns of the housing market and will hence be able to time their purchases to get better prices. On the other hand, first-time buyers might be less dependent on chains (that is, they do not need to sell a house before buying) and can thus better arbitrage across seasons. The regressions tend to point to slightly stronger seasonality in prices paid by former-owner occupiers, favouring the second hypothesis, though as before, the results can also be driven by the natural loss of precision caused by the relatively small number of first-time buyers in the market.

Table 4 shows the corresponding numbers for average differences in *real* housing price growth. Since the average difference in overall inflation rates across summers and winters is around 2 percent, the average difference in real housing price growth is roughly equivalent to the difference in nominal housing price inflation minus 2 percent.

Table 4: Average Difference in Annualized Real Housing Price Growth Between Summer and Winters, by Region and Type of House or Buyer

	All Houses (All buyers)		Existing houses (All buyers)		New houses (All buyers)		Former owner occupiers (All houses)		First-time buyer (All houses)	
	Coef.	Std. Error	Coef.	Std. Error	Coef.	Std. Error	Coef.	Std. Error	Coef.	Std. Error
E. Anglia	8.597*	(3.589)	7.787*	(3.780)	3.114	(6.815)	10.160**	(3.531)	3.444	(4.483)
E.Midlands	10.148**	(3.675)	10.854**	(3.716)	-0.027	(5.989)	11.766**	(3.951)	7.495+	(3.772)
Gr. London	4.161	(3.006)	4.435	(3.034)	15.296	(9.526)	3.585	(2.803)	4.115	(3.275)
N. West	6.224*	(2.784)	7.620**	(2.847)	-4.022	(7.140)	7.456*	(3.012)	3.764	(2.905)
North	-0.224	(3.238)	0.284	(3.356)	-0.637	(5.747)	-1.315	(3.327)	1.446	(3.910)
S. East	5.677+	(3.015)	6.084*	(2.990)	0.756	(4.211)	6.854*	(3.001)	2.259	(3.109)
S. West	8.569*	(3.579)	8.863*	(3.701)	4.188	(4.997)	9.567*	(3.687)	3.869	(4.012)
W. Midlands	5.291	(3.800)	4.983	(3.823)	14.448+	(8.201)	6.02	(4.004)	4.285	(3.656)
Yorkshire&Humb	5.468+	(3.113)	6.195+	(3.169)	0.53	(6.536)	6.155+	(3.132)	5.521	(3.467)
N. Ireland	7.422*	(3.580)	9.976*	(4.186)	11.885*	(5.813)	4.701	(4.544)	8.936+	(5.216)
Scotland	9.305**	(2.462)	12.317**	(2.695)	12.163+	(7.260)	11.010**	(2.544)	4.476	(3.021)
Wales	6.895+	(3.723)	6.818+	(3.749)	-1.32	(8.084)	7.659*	(3.743)	5.021	(3.957)
U.K.	6.105*	(2.354)	6.788**	(2.393)	3.444	(2.579)	7.016**	(2.387)	3.760+	(2.255)

Note: The Table shows the coefficients (and standard errors) on the dummy variable S_t (second and third quarters) in the regression g_t=a+b×Summer$_t$+e$_t$, where g_t is the annualized rate of *real* housing price inflation; a is a constant (omitted). The equations use quarterly data from 1983 to 2005. Robust standard errors in parentheses. +Significant at the 10%; *significant at the 5%; **significant at 1%.

The behavior of prices is mimicked by that of the number of transactions. Table 5 shows the average differences in growth rates in the number of transactions between summers and winters. The Table reports the slope coefficients and standard errors of the summer-dummy regression (2) corresponding to each region. The annualized difference in growth rates is roughly 120 percent. Northern Ireland and the North region show the smallest average difference, which is roughly 100 percent. As the Table shows, the difference is stronger for former-owner occupiers than for first-time buyers, consistent with the price patterns observed before. (Unfortunately, the data are not disaggregated by type of house).

Table 5: Average Difference in Annualized Growth Rates in the Number of Transactions Between Summer and Winters, by Region and Type of Buyer

	All Houses (All buyers)		Former owner occupiers (All houses)		First-time buyer (All houses)	
	Coef.	Std. Error	Coef.	Std. Error	Coef.	Std. Error
E. Anglia	137.066**	(22.313)	214.294**	(38.983)	136.538**	(29.901)
E.Midlands	154.761**	(44.188)	215.595**	(58.098)	204.546*	(89.538)
Gr. London	138.723**	(40.132)	204.390**	(71.944)	112.855**	(28.587)
N. West	121.901**	(17.117)	155.872**	(19.788)	105.037**	(21.158)
North	95.811**	(16.419)	183.704**	(35.753)	82.895*	(37.257)
S. East	136.708**	(16.753)	164.647**	(18.295)	102.878**	(15.453)
S. West	140.322**	(24.109)	182.283**	(27.215)	109.224**	(21.898)
W. Midlands	155.984**	(29.471)	207.046**	(37.535)	112.131**	(24.538)
Yorkshire&Humb	121.736**	(20.539)	171.579**	(31.494)	106.622**	(22.217)
N. Ireland	118.920**	(38.895)	172.178*	(74.599)	119.912**	(41.468)
Scotland	169.156**	(42.906)	320.131**	(67.460)	84.948**	(25.485)
Wales	167.241**	(39.668)	184.066**	(38.418)	158.468**	(40.656)
U.K.	130.277**	(20.738)	168.636**	(22.563)	102.730**	(19.682)

Note: The Table shows the coefficients (and standard errors) on the dummy variable S_t (Summer) in the regression x_t=a+b×Summert+et, where x_t is the annualized growth rate of the number of transactions; a is a constant (omitted). The equations use quarterly data from 1983 to 2005. Robust standard errors in parentheses. + Significant at 10%; * Significant at the 5%; ** significant at 1%.

Put together, the data point to a strong seasonal cycle, with a large increase in transactions and prices during the summer relative to the winter term. Also, the seasonal patterns are similar across regions, except for the North region, which tends to display less seasonality in prices.

Rents

Data on rents are not documented in as much detail as the data on prices. The series available corresponds to the aggregate of the U.K. and comes from the ODPM; the data are not disaggregated by region. We run regressions using as dependent variables both the rent levels and the log of rents on the summer-term dummy. We also include, where indicated, a trend term. The results are summarized in Table 6, which shows that there is virtually no seasonality in rents for the U.K. as a whole. This is in line with anecdotal evidence suggesting that rents are fairly sticky.

Table 6: Summer Differentials in Rents in the U.K.

	Rents		log(Rent)	
Summer-dummy S_t	-47.90833	12.53771	-0.01406	0.00743
	(255.798)	(29.529)	(0.091)	(0.010)
Trend		61.67964**		0.02194**
		(1.276)		(0.000)

Note: The Table shows the coefficients (and standard deviations) on the dummy variable S_t (second-third quarters) in the regressions $x_t=a+b\times S_t+e_t$, where x_t is either the rent level or the log of the rent; a is a constant (omitted); a trend term is included where indicated. Data are quarterly, from 1989-2005. Robust standard errors in parentheses. + Significant at 10%; * Significant at the 5%; ** significant at 1%.

Mortgage Rates

Interest rates in the U.K. do not seem to exhibit a seasonal pattern. The evidence is summarized in Table 7, which shows the summer dummy coefficients for different interest rate series provided by the Bank of England. The first column shows the results for the quarterly average of the repo (base) rate; the second column shows the corresponding results for the average interest rate charged by 4 U.K. major banks (Barclays Bank, Lloyds Bank, HSBC, and National Westminster Bank); and the third column shows the results for the weighted average standard variable mortgage rate from Banks and Building Societies. The first two series cover the period 1978 through 2005, whereas the third goes from 1994 through 2005.

As the Table shows, none of the interest rate measures appears to be different, on average, during the summer term.

Table 7: Summer Differentials in Interest Rates in the U.K.

	Repo rate	Bank-4 Rate	Mortgage Rate
Summer-dummy S_t	-0.163	-0.144	0.018
	(0.701)	(0.696)	(0.310)

Note: The Table shows the slope coefficients (and standard deviations) on the dummy variable S_t (second-third quarters) in the regressions x_t=a+b×S_t+e_t, where x_t is the Repo rate, the average of the 4 largest banks, or the mortgage interest rate, correspondingly; a is a constant (omitted). The equations use quarterly data from 1978 to 2005, except for the mortgage rate series, which starts in 1994. Robust standard errors in parentheses. + Significant at 10%; * Significant at the 5%; ** significant at 1%.

Housing Market Seasonality in the U.S.

Housing Price Changes

As noted before, the U.S. aggregate price index displays no seasonal patterns. The question is whether this result masks different seasonal patterns at a more disaggregated level. As it turns out, this not the case. In the interest of space, and given the lack of seasonality in the data, we omit the graphs and summarize the results in Table 8, displaying the summer-effects coefficients and standard deviations using state-level data.[11] As shown in the Table, only in one state (Kentucky) there is a statistically significant summer effect on prices. (Similar results are found when using the metropolitan- statistical-area-level data from the same source.[12]) The summer effect is also

[11] The data correspond to the 50 states and the district of Columbia.

[12] This is based on the 379 metropolitan areas defined by OFHEO.

insignificant from an economic point of view; the only states with sizeable (though not statistically significant) effects are Hawaii (exhibiting a negative summer effect), and Massachusetts, South Dakota, Delaware, and West Virginia (exhibiting a positive summer effect).

Real prices display a similar pattern, as there is no significant differences in overall inflation rates across seasons in the US (results not shown).[13]

Table 8: Average Difference in Annualized Housing Price Growth between Second-Third Quarters and Fourth-First Quarters, by US State.

State	Coef.	Std. Error	State	Coef.	Std. Error
Alabama	0.008	(3.591)	Montana	4.453	(3.426)
Alaska	-0.692	(1.655)	North Carolina	0.406	(0.904)
Arizona	0.789	(2.473)	North Dakota	-6.223	(5.320)
Arkansas	0.588	(1.930)	Nebraska	-2.829	(1.819)
California	2.094	(1.757)	New Hampshire	3.030	(3.231)
Colorado	1.606	(1.387)	New Jersey	-0.237	(1.685)
Connecticut	2.890	(2.004)	New Mexico	0.207	(1.719)
District of Columbia	1.967	(6.001)	Nevada	-0.688	(2.128)
Delaware	10.757	(6.212)	New York	-0.903	(2.262)
Florida	0.133	(2.524)	Ohio	0.853	(0.797)
Georgia	-0.119	(1.202)	Oklahoma	-0.538	(1.595)
Hawaii	-31.388	(30.242)	Oregon	0.406	(1.798)
Idaho	-0.131	(2.086)	Pennsylvania	1.008	(1.495)
Illinois	0.292	(3.326)	Rhode Island	2.345	(2.276)
Indiana	0.066	(1.287)	South Carolina	-0.671	(1.723)
Iowa	-0.776	(1.238)	South Dakota	16.066	(10.273)
Kansas	1.031	(0.935)	Tennessee	-1.878	(1.705)
Kentucky	-1.883*	(0.941)	Texas	0.104	(1.449)
Louisiana	-2.335	(1.615)	Utah	-1.363	(1.590)
Maine	2.000	(1.878)	Virginia	-0.427	(1.278)
Maryland	0.570	(1.258)	Vermont	-3.082	(7.957)
Massachusetts	16.131	(10.186)	Washington	1.084	(1.623)
Michigan	1.616	(1.468)	Wisconsin	2.594	(1.886)
Minnesota	-0.025	(1.373)	West Virginia	9.703	(10.014)
Missouri	2.726	(2.351)	Wyoming	-0.572	(2.663)
Mississippi	-0.004	(3.132)			

Note: The Table shows the coefficients (and standard errors) on the dummy variable S_t (second and third quarters) in the regression g_t=a+b×$Summer_t$+e_t, where g_t is the annualized rate of *nominal* housing price inflation; a is a constant (omitted). The equations use quarterly data from 1975 to 2005. Robust standard errors in parentheses. +Significant at the 10%; *significant at the 5%; **significant at 1%.

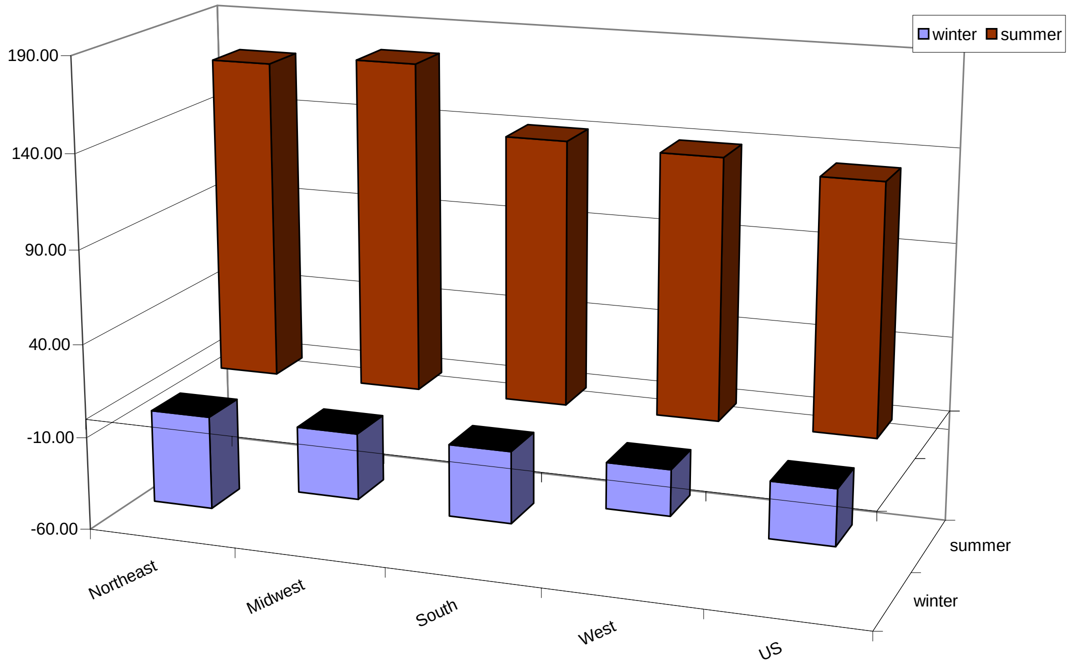
Number of Transactions

As already observed, the U.S. as a whole displays a strong seasonality in the number of transactions. This remains true across all four major regions of the U.S. (state-level data are not available). The growth rates in the number of transactions in summers and winters are plotted in Figure 5.

[13] On the lack of seasonality of overall inflation in the U.S., see Barsky and Miron (1989).

The average difference across seasons, together with the standard errors are summarized in Table 9. In sum, the data for the U.S. point to a strong seasonal pattern in the number of transactions, with no discernible seasonal pattern in housing prices.

Figure 5: Annualized growth rate of the number of transactions in summers and winters in the U.S. and its regions



Note: Annualized growth rates of the number of transactions in summers (second and third quarters) and winters (fourth and first quarters) in the U.S. and its regions, 1975-2005. (Data for the U.S. as a whole corresponds to 1968-2005.)

Table 9: Average Difference in Annualized Growth Rates in the Number of Transactions Between Summer and Winters, by Regions in the U.S.

Region	Coef.	Std. Error
Northeast	220.718**	(19.762)
Midwest	210.968**	(27.558)
South	179.038**	(21.219)
West	162.818**	(25.816)
United States	162.354**	(19.369)

Note: The Table shows the coefficients (and standard errors) on the dummy variable S_t (Summer) in the regression x_t=a+b×$Summer_t$+e_t, where x_t is the annualized growth rate of the number of transactions; a is a constant (omitted). The equations use quarterly data from 1975 to 2005 for the regions and 1968-2005 for the U.S. as a whole. Robust standard errors in parentheses. + Significant at 10%; * Significant at the 5%; ** significant at 1%.

Rents

Data on rents for the U.S. come from the Bureau of Labor Statistics (BLS); as a measure of rents we use the non-seasonally adjusted series of owner's equivalent rent and the non-seasonally adjusted rent of primary residence; both series are produced for the construction of the CPI and correspond to averages over all cities. For each series, we run regressions using as dependent variables both the rent levels and the log of rents on the summer-term dummy. we also include, where indicated, a trend term. The results are summarized in Tables 10 (owner's equivalent rent) and 11 (rent of primary residence). Both Tables show that there is no evidence of seasonality in rents for the U.S. as a whole.

Table 10: Summer Differential in Rents in the U.S.: Owner's Equivalent Rent

	Rents		log(Rent)	
Summer-dummy S_t	-0.19638	-0.19638	-0.00102	-0.00102
	(8.133)	(0.269)	(0.051)	(0.006)
Trend		1.45183**		0.00905**
		(0.005)		(0.000)

Note: The Table shows the coefficients (and standard deviations) on the dummy variable S_t (second-third quarters) in the regressions $x_t=a+b\times S_t+e_t$, where x_t is either the rent level or the log of the rent; a is a constant (omitted); a trend term is included where indicated. Data are quarterly, from 1983-2005. Robust standard errors in parentheses. + Significant at 10%; * Significant at the 5%; ** significant at 1%. (BLS, owner's equivalent rent.)

Table 11: Summer Differential in Rents in the U.S.: Rent of Primary Residence

	Rents		log(Rent)	
Summer-dummy S_t	-0.16594	-0.16594	-0.00098	-0.00098
	(7.120)	(0.638)	(0.047)	(0.005)
Trend		1.26671**		0.00827**
		(0.012)		(0.000)

Note: The Table shows the coefficients (and standard deviations) on the dummy variable S_t (second-third quarters) in the regressions $x_t=a+b\times S_t+e_t$, where x_t is either the rent level or the log of the rent; a is a constant (omitted); a trend term is included where indicated. Data are quarterly, from 1983-2005. Robust standard errors in parentheses. + Significant at 10%; * Significant at the 5%; ** significant at 1%. (BLS, rent of primary residence.)

Mortgage Rates

Interest rates in the U.S. do not exhibit a seasonal pattern (Barsky and Miron, 1989). Since housing service costs are of particular interest here, we summarize In Table 12 the summer effect (or lack thereof) in mortgage rates. The data come from the Board of Governors of the Federal Reserve and correspond to contract interest rates on commitments for fixed-rate first mortgages;

the data are quarterly averages beginning in 1972; the original data are collected by Freddie Mac. As the Table shows, mortgage rates do not appear to be higher on average during the summer term, consistent with the findings in Barsky and Miron (1989).

Table 12: Summer Differential in Mortgage Rates in the U.S.

	Mortgage Rate
Summer-dummy S_t	0.104
	(0.477)

Note: The Table shows the slope coefficient (and standard deviation) on the dummy variable S_t (second-third quarters) in the regressions $x_t=a+b\times S_t+e_t$, where x_t is the average mortgage interest rate; a is a constant (omitted). The equations use quarterly data from 1972 through 2005. Robust standard errors in parentheses. + Significant at 10%; * Significant at the 5%; ** significant at 1%.

Housing Market Seasonality in Belgium and France

Tables 13 and 14 show the housing price regressions for Belgium and France, disaggregated by regions with available data. As the Tables show, in both countries all regions display a strong seasonal pattern, comparable to that reported for the country as a whole. Data on transactions at the regional level are not available. As noted in the Data Appendix, the housing price indexes for these countries are not quality adjusted and hence seasonal variation in prices might mask variation in the quality of the houses on the market; this is why we emphasize throughout the paper the results from the U.K. and the U.S.

Table 13: Average Difference in Annualized Housing Price Growth between Second-Third Quarters and Fourth-First Quarters in Belgium, by Region.

Region	Coef.	Std. Error
Great Brussels	13.242**	(3.039)
Flanders	10.753**	(1.746)
Wallonia	19.329**	(1.903)

Note: The Table shows the coefficients (and standard errors) on the dummy variable S_t (second and third quarters) in the regression $g_t=a+b\times Summer_t+e_t$, where g_t is the annualized rate of *nominal* housing price inflation; a is a constant (omitted). The equations use quarterly data from 1981 to 2005. Robust standard errors in parentheses. +Significant at the 10%; *significant at the 5%; **significant at 1%.

Table 14: Average Difference in Annualized Housing Price Growth between Second-Third Quarters and Fourth-First Quarters in France, by Region.

Region	Coef.	Std. Error
Ile-de-France	9.275**	(2.294)
Province (All regions except Ile-de-France)	17.347**	(1.906)
Provence-Alpes-Côte d'Azur	10.915**	(2.624)
Rhône-Alpes	11.977**	(2.648)

Note: The Table shows the coefficients (and standard errors) on the dummy variable S_t (second and third quarters) in the regression g_t=a+b×$Summer_t$+e_t, where g_t is the annualized rate of *nominal* housing price inflation; a is a constant (omitted). The equations use quarterly data from 1994 to 2005. Robust standard errors in parentheses. +Significant at the 10%; *significant at the 5%; **significant at 1%.

3 Quantifying the Price Puzzle

We carry out a back-of-envelope calculation using the findings for the U.K., given that the data are of better quality than those in other continental-European countries that also feature seasonality in prices. The U.S., as seen, displays no seasonality in prices.

We argued before that the predictability and size of the seasonal variation in housing prices in some countries pose a puzzle to models of the housing market relying on standard asset-market equilibrium conditions. In particular, the equilibrium condition embedded in most dynamic general-equilibrium models states that the marginal benefit of housing services should equal the marginal cost. Following Poterba (1984) the asset-market equilibrium conditions for any seasons $j = s$ (summer), w (winter) at time t is:[14]

$$d_{t+1,j'} + (p_{t+1,j'} - p_{t,j}) = c_{t,j} \cdot p_{t,j} \tag{4}$$

where j' is the corresponding season at time $t+1$, $p_{t,j}$ and $d_{t,j}$ are the real asset price and rental price of housing services, respectively; $c_{t,j} \cdot p_{t,j}$ is the real *gross* (gross of capital gains) $t-$period cost of housing services of a house with real price $p_{t,j}$; and $c_{t,j}$ is the sum of after-tax depreciation, repair costs, property taxes, mortgage interest payments, and the opportunity cost of housing equity. Note that the formula assumes away risk (and hence no expectation terms are included); this is appropriate in this context because we are focusing on a "predictable" variation of prices.[15] As in

[14] See also Mankiw and Weil (1989) and Muellbauer and Murphy (1997), among others.

[15] Note that Poterba's formula also implicitly assumes linear preferences and hence perfect intertemporal substitution. This is a good assumption in the context of seasonality, given that substitution across semesters (or relatively

Poterba (1984), we make the following simplifying assumptions so that service-cost rates are a fixed proportion of the property price, though still potentially different across seasons ($c_{t,j} = c_{t+2,j} = c_j$, $j = s, w$): *i*) Depreciation takes place at rate δ_j, $j = s, w$, constant for a given season, and the house requires maintenance and repair expenditures equal to a fraction κ_j, $j = s, w$, also constant for a given season. *ii*) The income-tax-adjusted real interest rate and the marginal property tax rates (for given real property prices) are constant over time, though also potentially different across seasons; they are denoted, respectively as r_j and τ_j, $j = s, w$ (in the data, as seen, they are actually constant across seasons; we come back to this point below).[16] This yields $c_j = \delta_j + \kappa_j + r_j + \tau_j$, for $j = s, w$.

Subtracting (4) from the corresponding expression in the following season and using the condition that there is no seasonality in rents ($d_w \approx d_s$), we obtain:

$$\frac{p_{t+1,s} - p_{t,w}}{p_{t,w}} - \frac{p_{t,w} - p_{t-1,s}}{p_{t-1,s}} \frac{p_{t-1,s}}{p_{t,w}} = c_w - c_s \cdot \frac{p_{t-1,s}}{p_{t,w}} \tag{5}$$

Considering the *real* differences in house price growth rates documented for the whole of the U.K., $\frac{p_s - p_w}{p_w} = 7.04\%$, $\frac{p_w - p_s}{p_s} = 0.75\%$, the left-hand side of (5) equals $6.3\% \approx 7.04\% - 0.75\% \cdot \frac{1}{1.0075}$. Therefore, $\frac{c_w}{c_s} = \frac{0.063}{c_s} + \frac{1}{1.0075}$. The value of c_s can be pinned-down from equation (4) with $j = s$, depending on the actual rent-to-price ratios (d/p) in the economy. In Table 15, we summarize the extent of seasonality in service costs $\frac{c_w}{c_s}$ implied by the asset-market equilibrium conditions, for different values of d/p (and hence different values of $c_s = \frac{d_w}{p_s} + \frac{p_w - p_s}{p_s} = \frac{d_w}{p_s} + 0.75\%$).

Table 15: Ratio of Winter-To-Summer Cost Rates

(annualized) d/p Ratio	Relative winter cost rates $\frac{c_w}{c_s}$
1.0%	459%
2.0%	328%
3.0%	267%
4.0%	232%
5.0%	209%
6.0%	193%

As the Table illustrates, a remarkable amount of seasonality in service costs is needed to explain the differences in housing price inflation across seasons. Specifically, assuming annualized rent-to-

short periods of time) should in principle be quite high.

[16] We implicitly assume the property-price brackets for given marginal rates are adjusted by inflation rate, though strictly this is not the case (Poterba, 1984): inflation can effectively reduce the cost of homeownership. This, however, should not alter the conclusions concerning seasonal patterns emphasized here. As in Poterba (1984) we also assume that the opportunity cost of funds equals the cost of borrowing.

price ratios in the range of 2 through 5 percent, total costs in the winter should be between 328 and 209 percent of those in the summer. Depreciation and repair costs $(\delta_j + \kappa_j)$ might be seasonal, being potentially lower during the summer.[17] But income-tax-adjusted interest rates and property taxes $(r_j + \tau_j)$, two major components of service costs are not seasonal. Since depreciation and repair costs are only part of the total costs, given the seasonality in other components, the implied seasonality in depreciation and repair costs across seasons in the U.K. is even larger. Assuming, quite conservatively, that the a-seasonal component $(r_j + \tau_j = r + \tau)$ accounts for only 50 percent of the service costs in the summer $(r + \tau = 0.5c_s)$, then, the formula for relative costs $\frac{c_w}{c_s} = \frac{\delta_w + \kappa_w + 0.5c_s}{\delta_s + \kappa_s + 0.5c_s}$ implies that the ratio of depreciation and repair costs between summers and winters is $\frac{\delta_w + \kappa_w}{\delta_s + \kappa_s} = 2\frac{c_w}{c_s} - 1$.[18] For rent-to-price ratios in the range of 2 through 5 percent, depreciation and maintenance costs in the winter should be between 557 and 318 percent of those in the summer. (If the a-seasonal component $(r + \tau)$ accounts for 80 percent of the service costs $(r + \tau = 0.8c_s)$, the corresponding values are 1542 and 944 percent). By any metric, these figures seem extremely large.

Let us now for the sake of the argument concede that these figures are indeed as large as implied by the asset-pricing equilibrium condition, the question is then: Why is it the case that depreciation and repair costs are so seasonal in the U.K. (and potentially higher in other continental European countries that exhibit larger seasonality in prices) while in other countries, such as the U.S., they are a-seasonal? Deviations from the standard asset-pricing equilibrium condition are needed to match both the U.K. and the U.S. data.

The need to deviate from the asset-market approach has been acknowledged, in a different context, among others, by Stein (1995). While static in nature, Stein's model is capable of generating unexpected booms and busts in prices (and transactions) in a rational-expectation setting. In a dynamic setting with forward-looking agents, however, predictably large changes in prices cannot be sustained: Expected price increases in the next season will actually be priced in the current season (or, in other words, sellers will refuse to sell at lower prices today given the perspective of higher prices in the next season); similarly, prospective buyers will benefit from waiting (at most a

[17] Good weather can help with external repairs and owners' vacation might reduce the opportunity cost of time—though it is key here that leisure is not too valuable for the owners.

[18] Call λ the asesonal component as a fraction of the summer service cost rate: $r + \tau = \lambda c_s$, $\lambda \in (0, 1)$ (and hence $\delta_s + \kappa_s = (1 - \lambda)c_s$). Then: $\frac{c_w}{c_s} = \frac{\delta_w + \kappa_w + \lambda c_s}{\delta_s + \kappa_s + \lambda c_s} = \frac{\delta_w + \kappa_w + \lambda c_s}{c_s}$. Or $c_w = \delta_w + \kappa_w + \lambda c_s$. Hence: $\frac{c_w - \lambda v_s}{(1-\lambda)c_s} = \frac{\delta_w + \kappa_w}{(1-\lambda)c_s}$; that is $\frac{\delta_w + \kappa_w}{\delta_s + \kappa_s} = \frac{c_w}{(1-\lambda)c_s} - \frac{\lambda}{1-\lambda}$, which is increasing in λ for $\frac{c_w}{c_s} > 1$.

few months) and paying a significantly lower price. Even when agents are both sellers and buyers, if they are aware of the differences in prices (or price growth rates), in a dynamic setting they will seek to sell in the summer and to buy in the winter; the excess supply in the summer will then push prices down, while the excess demand in the winter will push them up.

In the next Section, we develop a search model for the housing market that can generate significant differences in the number of transactions across seasons. The model can also deliver seasonality in prices, comparable to that observed in U.K. data, as well as no seasonality, as in U.S. data.

4 A Search Model for the Housing Markets

The basic setup of the model builds on previous contributions by Krainer (2001), Wheaton (1990), and Williams (1995), which in turn borrow from the labor search literature (see, for example, Pissarides (1990)).

The model economy is populated by a unit measure of infinitely lived agents who have linear preferences over a non-durable consumption good and a housing good. Each period agents receive a fixed endowment of the consumption good which they can use to buy houses with. The housing good is indivisible and agents can only live in one house at a time (though they can potentially own more than one). The housing stock is constant and there are as many houses as agents. Each house starts a period in one of two "states:" It can be either "matched," when it delivers positive housing services flows to its owner, or "on sale," when it does no longer yield any services to its owner. As long as a house is "matched," it yields idiosyncratic housing services ε to its owner, which we assume to be constant over time. The "quality of the match" ε is only observed by the potential buyer, but not by the seller.

There are two seasons, $j = s, w$ (for summer and winter); each model period is a season, and seasons alternate. At the beginning of a period, each match has a probability $(1-\phi^j)$ of breaking, and the house goes "on sale." The parameter ϕ^j is the only (*ex ante*) difference between any two seasons.

Agents who are *not* matched to a house seek to buy one ("buyers") and agents who own houses where the match is broken seek to sell them ("sellers"). Note that an agent may be only "buyer," only "seller," and both "buyer" and "seller." Also, sellers may have multiple houses to sell. Buyers and sellers are randomly matched. Each period a buyer visits only one house, and each house is visited by only one buyer.

We call v^j the stock of vacant houses and b^j be the number of agents without a house in season $j = s, w$, all of which are determined in equilibrium. Since when a match is destroyed a homeowner becomes both a buyer and a seller simultaneously, it is always the case that $v^j = b^j$, that is, the number of vacant houses equals the number of potential buyers.

The sequence of events is as follows. At the beginning of period t, an existing match between a homeowner and his house breaks with probability $1 - \phi^j$, adding to the stock of vacant houses and potential buyers. Every seller meets with a buyer randomly. The potential buyer observes the utility services ε (not observable to the seller) generated by the match and decides whether or not to buy. If the transaction goes through, the buyer pays p^j to the seller, and starts enjoying the utility flow from that period. If the transaction does not go through, the house lies empty and the buyer does not receive any flow utility from housing. We discuss two price-setting scenarios. In the first and benchmark, which we call "the seller's market," the seller posts a price and the buyer decides whether the match quality is high enough to pay the price. In the second, which we call "the buyer's market," the buyer sets a price and makes a take-it-or-leave-it offer to the seller.

4.1 Sellers' Market

4.1.1 Utility services

The model embeds the intuitively appealing notion that in a market with many houses on sale a buyer can find a house closer to her ideal and hence her willingness to pay increases. We model this idea by assuming that the (idiosyncratic) quality of a match, ε, in season j follows a distribution $F^j(\varepsilon)$ with positive support and finite mean such that:

$$F^j(\varepsilon) \leqslant F^{j'}(\varepsilon); \forall \varepsilon \Leftrightarrow v^j \geqslant v^{j'} \text{ for } j, j' = s, w \tag{6}$$

That is, $F^j(.)$ stochastically dominates $F^{j'}(.)$ if and only if $v^j > v^{j'}$, where v^j $(v^{j'})$, the stock of houses in season j, is endogenously determined. In other words, when the stock of houses v^j is bigger, the draw ε is likely to be higher.[19]

[19]One way to interpret this assumption is the following. Suppose there are v (discrete) units of houses. The buyer can sample all v houses (e.g. by searching online or through newspapers). Let $(X_1, X_2, ..X_v)$ denote an *iid* random sample of idiosyncratic utility flows from the continuous distribution $G(.)$. Let ε be the maximum X_i; then

Note that ε captures the quality of a match between a house and the potential buyer. In other words, for any vacant house, the potential utility services are idiosyncratic to the match between the house and the buyer. Hence, ε is not the type of the house (or of the seller who owns a particular house); indeed, there is only one representative house in our model, with the utility derived from living in the house being idiosyncratic. This is consistent with the data we look at, which are adjusted for houses' characteristics, such as size and location, but not for the (unobserved) quality of a match.[20]

To study pricing and transaction decisions, we first need to derive the value of living in a house if a transaction goes through. The value function for a homeowner who lives in a house with quality ε in season s is given by:

$$H^s(\varepsilon) = \varepsilon + \beta\phi^w H^w(\varepsilon) + \beta(1-\phi^w)[V^w + B^w]$$

With probability $(1-\phi^w)$ he receives a moving shock and becomes both a seller and a buyer, with continuation value $(V^w + B^w)$, where V^j is the lifetime utility of being a seller and B^j is the lifetime utility of being a buyer in season j, defined below. With probability ϕ^w he keeps receiving utility services ε and stays in the house. (Notice that the formula for $H^w(\varepsilon)$ is perfectly isomorphic to $H^s(\varepsilon)$; in the interest of space we omit here and throughout the paper the corresponding expressions for season w.) The value of being a homeowner can be therefore re-written as:

$$H^s(\varepsilon) = \frac{1+\beta\phi^w}{1-\beta^2\phi^w\phi^s}\varepsilon + \frac{\beta(1-\phi^w)(V^w+B^w) + \beta^2\phi^w(1-\phi^s)(V^s+B^s)}{1-\beta^2\phi^w\phi^s}. \tag{7}$$

4.1.2 The Buyer

Upon visiting a house, the buyer draws a match quality ε from the distribution $F^s(\varepsilon) \equiv F(\varepsilon \mid v^s)$ in season s. (And, correspondingly, from $F^w(\varepsilon) \equiv F(\varepsilon \mid v^w)$ in season w). Since the match quality is idiosyncratic to a house and buyer, it is natural to assume that the seller does not observe ε. Thus, in a seller's market, the seller posts a price p^s independent of the level of ε. The buyer's value function in season s is:

$$B^s = E^s_\varepsilon \max\{H^s(\varepsilon) - p^s, \beta B^w\}, \tag{8}$$

the distribution of ε is $F^v(\varepsilon) = [G(X)]^v$. It follows that $F^v(\varepsilon) \leq F^{v'}(\varepsilon)$ for $v > v'$. Intuitively, as the sample size v increases, the maximum becomes "stochastically larger". Assume that in each period buyers only visit the house that is ranked first. The price is then determined by the take-it-or-leave-it offer made by the corresponding seller in a seller's market, or by the buyer in a buyer's market.

[20] Neither repeat-sale indices nor hedonic price indices can control for the quality of a match.

where $E^s[.]$ indicates the expectation taken with respect to the distribution $F^s(.)$.

As said, buyers consume no housing services until they find a successful match. This can be the case, for example, if buyers searching for a house pay a rent equal to the utility they derive from the rented property; what is key is that the rental property is not owned by the same potential seller with whom the buyer meets.

Since $H^s(\varepsilon)$ is increasing in ε, a "reservation policy," whereby the buyer accepts the posted price if ε exceeds a cutoff level, is optimal. The transaction is hence carried out if $\varepsilon > \varepsilon^s$, where the cutoff ε^s is given by:

$$H^s(\varepsilon^s) - p^s = \beta B^w. \tag{9}$$

$1 - F^s(\varepsilon^s)$ is thus the probability that a transaction is carried out. From (7), the response of the reservation quality ε^s to a change in price is given by:

$$\frac{\partial \varepsilon^s}{\partial p^s} = \frac{1 - \beta^2 \phi^w \phi^s}{1 + \beta \phi^w}. \tag{10}$$

4.1.3 The Seller

Taking the optimal decision rules of the buyer as given, the seller chooses a price to maximize the expected surplus value of a sale. The seller's value function is hence

$$V^s = \beta V^w + u + \max_p \{[1 - F^s(\varepsilon^s(p))](p - \beta V^w - u)\}, \tag{11}$$

where u is the utility flow from being a seller; this could be interpreted, for example, as a net rental income received by the seller while the house in on the market; to be consistent with the data, we assume that the rental income u does not vary across seasons. Again, what is key is that the tenant is not the same potential buyer who visits the house.

The optimal price p^s solves

$$[1 - F^s(\varepsilon^s)] - [p - \beta V^w - u] f^s(\varepsilon^s) \frac{\partial \varepsilon^s}{\partial p^s} = 0. \tag{12}$$

Rearranging terms we obtain:

$$\underbrace{\frac{p^s - \beta V^w - u}{p^s}}_{\text{mark-up}} = \underbrace{\left[\frac{p^s f^s(\varepsilon^s) \frac{\partial \varepsilon^s}{\partial p^s}}{1 - F^s(\varepsilon^s)}\right]^{-1}}_{\text{inverse-elasticity}},$$

which makes clear that the price-setting problem of the seller is similar to that of a monopolist who sets a markup equal to the inverse of the elasticity of demand (where demand in this case is given by the probability of a sale, $1 - F^s(\varepsilon^s)$.

4.1.4 Market equilibrium

Prices Let $S_v^s \equiv p^s - \beta V^w - u$ be the surplus to a seller from a housing transaction. The optimal decisions of the buyer (10) and the seller (12) together imply:

$$S_v^s = \frac{1 - F^s(\varepsilon^s)}{f^s(\varepsilon^s)} \frac{1 + \beta\phi^w}{1 - \beta^2\phi^w\phi^s}. \tag{13}$$

Equation (13) says that the net surplus to a seller generated by the transaction is higher when the distribution has a thicker tail, $\frac{1-F^s(\varepsilon^s)}{f^s(\varepsilon^s)}$. Together with the value function of the seller, the optimal price satisfies (see derivation in Appendix 8.1):

$$p^s = \frac{u}{1-\beta} + \frac{1 - \beta^2 F^s(\varepsilon^s)}{1 - \beta^2} S_v^s + \frac{\beta\left[1 - F^w(\varepsilon^w)\right]}{\left(1 - \beta^2\right)} S_v^w. \tag{14}$$

Reservation quality Let $S_b^s(\varepsilon) \equiv H^s(\varepsilon) - p^s - \beta B^w$ be the surplus to a buyer from buying a house with flow value ε. The reservation quality ε^s satisfies $S_b^s(\varepsilon^s) = 0$. Using (7), the surplus to a buyer is

$$S_b^s(\varepsilon) = H^s(\varepsilon) - H^s(\varepsilon^s) = \frac{1 + \beta\phi^w}{1 - \beta^2\phi^w\phi^s}(\varepsilon - \varepsilon^s). \tag{15}$$

Let $S^s(\varepsilon) \equiv S_b^s(\varepsilon) + S_v^s$ be the total surplus from a transaction with flow value ε. An equilibrium with positive number of transactions exists if the total surplus is positive. By definition, in a seller's market ε^s also satisfies $S^s(\varepsilon^s) = S_v^s$, that is

$$H^s(\varepsilon^s) = S_v^s + \beta B^w + \beta V^w + u, \tag{16}$$

which equates the housing value of a marginal owner in season s, $H^s(\varepsilon^s)$, to the sum of the surplus generated to the seller (S_v^s), plus the sum of outside options for the buyer (βB^w) and seller ($\beta V^w + u$). Using (7), ε^s solves:

$$\frac{1 + \beta\phi^w}{1 - \beta^2\phi^w\phi^s}\varepsilon^s = S_v^s + u + \frac{\beta\phi^w\left(1 - \beta^2\phi^s\right)}{1 - \beta^2\phi^w\phi^s}(V^w + B^w) - \frac{\beta^2\phi^w(1 - \phi^s)}{1 - \beta^2\phi^w\phi^s}(V^s + B^s). \tag{17}$$

The reservation quality ε^s depends on the sum of the outside options for buyers and sellers in both seasons, which in turn depend on the discounted values of becoming a seller and a buyer. The value of being a buyer, in turn, depends on the expected surplus value of homeownership, conditional on drawing a match quality $\varepsilon > \varepsilon^s$, that is, $E^s\left[S_b^s(\varepsilon) \mid \varepsilon > \varepsilon^s\right]$. The value of being a seller depends on the present value of the rental income u and the expected surplus from meeting other buyers. We can hence write (see Appendix 8.1)

$$\begin{aligned} & B^s + V^s \\ = \; & \frac{u}{1-\beta} + \frac{1 - F^s(\varepsilon^s)}{1 - \beta^2} E^s\left(S^s(\varepsilon) \mid \varepsilon > \varepsilon^s\right) + \frac{\beta\left[1 - F^w(\varepsilon^w)\right]}{1 - \beta^2} E^w\left(S^w(\varepsilon) \mid \varepsilon > \varepsilon^w\right). \end{aligned} \tag{18}$$

The thick-and-thin market equilibrium through the distribution F^j affects the equilibrium price and reservation quality (p^j, ε^j) in season $j = s, w$ through two channels: the tail of the distribution $\frac{1-F^j}{f^j}$ and the conditional mean $E^j [\varepsilon \mid \varepsilon > \varepsilon^j]$. As shown in (13), a thicker tail implies higher expected surplus to the seller S^j_v, which increases the equilibrium price p^j in (14). Similarly as shown in (15), higher a conditional mean implies a higher expected surplus to the buyer S^j_b. These two channels affect the total outside options of the buyer and the seller in (18), and as a result affect the reservation quality ε^j in (17).

Stock of vacant houses The law of motion for the stock of vacant houses (and hence for the stock of buyers) is

$$v^s = (1 - \phi^s) \left(v^w \left[1 - F^w \left(\varepsilon^w\right)\right] + 1 - v^w\right) + v^w F^w \left(\varepsilon^w\right)$$

where the first term includes houses that received a moving shock this season and the second term comprises vacant houses from last period that did not find a buyer. The expression simplifies to

$$v^s = v^w \phi^s F^w \left(\varepsilon^w\right) + 1 - \phi^s, \tag{19}$$

that is, in equilibrium v^s depends on the equilibrium matching values of ε and on the distribution $F(.)$.

An equilibrium is a vector $(p^s, p^w, B^s + V^s, B^w + V^w, \varepsilon^s, \varepsilon^w, v^s, v^w)$ that jointly satisfies equations (14),(17), (18) and (19), with the surpluses S^j_v and $S^j_b(\varepsilon)$ for $j = s, w$, derived as in (13), and (15).

4.2 Buyers' Market

The setup of the model is the same as before, except that now we assume that it is the buyer, rather than the seller, who makes a take-it-or-leave-it offer after visiting a house. The buyer extracts all the surplus from the seller by setting a price such that $S^j_v = 0$, $j = s, w$. Together with the value function of the seller, the equilibrium prices in seasons s and w become

$$p^s = p^w = \frac{u}{1 - \beta}, \tag{20}$$

which is the same as (14) after setting $S^s_v = S^w_v = 0$. The optimal strategy of the buyer still follows the reservation rule defined in (9). The equilibrium values of ε^s and $B^s + V^s$ are the same as in

(17) and (18) with $S_v^s = 0$. (The values for ε^w and $B^w + V^w$ are, as before, isomorphic to ε^s and $B^s + V^s$.) The equilibrium values of the stock of vacancies v^s and v^w follow the same law of motion as in (19).

5 Model-generated Seasonality of Prices and Transactions

5.1 Qualitative Results

We now derive the extent of seasonality in prices and transactions generated by the model, and show how they depend on whether sellers or buyers have the "power" to set prices, as well as on the level of rental income u. The driver for seasonality in the model is the probability of a moving shock, which we assume to be higher in the summer: $1 - \phi^s > 1 - \phi^w$. Using (19), the stock of vacant houses in season s is given by:

$$v^s = \frac{(1 - \phi^w)\, \phi^s F^w(\varepsilon^w) + 1 - \phi^s}{1 - \phi^w \phi^s F^s(\varepsilon^s)\, F^w(\varepsilon^w)}. \tag{21}$$

(The expression for v^w is correspondingly isomorphic). The *ex ante* higher probability of a shock in the summer $(1 - \phi^s > 1 - \phi^w)$ clearly has a direct positive effect on v^s. Because $F^s(\varepsilon)$ first-order stochastically dominates $F^w(\varepsilon)$ when $v^s > v^w$ (that is, $F^s(\varepsilon) \leqslant F^w(\varepsilon)\,; \forall\varepsilon$), this can amplify the seasonal shock to generate a *higher* seasonality in vacancies (as long as the indirect effect through higher ε^s is small). This amplification effect is what we call a thick-market externality. As shown in (14), since the rental income u is a-seasonal, housing prices are seasonal only if the surplus to the seller is seasonal. Two observations follow:

Remark 1 *In a buyer's market, there is no seasonality in prices.*

Remark 2 *In a seller's market, prices are seasonal. The extent of seasonality in prices is decreasing in the rental flow u.*

To see this, note that when the buyer sets prices, the surplus of the seller is zero; the equilibrium price is equal to the outside option of the seller, that is, his rental income u, which is a-seasonal. Hence, prices are a-seasonal in a buyer's market. When the seller sets prices his surplus is positive in both seasons; the equilibrium price is hence the sum of his outside option (u) plus a positive surplus from the sale. The surplus S_v^s, as shown in (13), is seasonal. Given $v^s > v^w$, the thick-market effect implies a thicker tail in quality in the hot season. In words, the quality of matches goes up in the summer and hence buyers's willingness to pay increases; sellers can then extract a higher surplus in the summer: thus, $S_v^s > S_v^w$. The extent of seasonality in prices decreases as the a-seasonal component—the outside option u—increases.

We next turn to the degree of seasonality in transactions. The number of transactions in equilibrium in season s is given by:

$$Q^s = v^s \left[1 - F^s\left(\varepsilon^s\right)\right]. \tag{22}$$

(An isomorphic expression holds for Q^w). A bigger stock of vacancies in the summer, $v^s > v^w$, tends to increase transactions in the summer. Whether buyers or sellers set prices also affects the degree of seasonality in transactions through the equilibrium value of ε^s. More specifically, a relatively higher reservation quality in the hot season, $\varepsilon^s > \varepsilon^w$, tends to decrease the degree of seasonality in transactions. As shown in (17), the equilibrium cutoff ε^s depends on the surplus to the seller (S_v^s) and on the sum of the seller's and buyer's outside options. We have already shown that $S_v^s > S_v^w$ in a seller's market because of the thick-market effect. Using (15), the thick market effect also implies that the expected surplus to the buyer is higher in the hot season, so the expected total surplus is also higher in the hot season: $E^s\left(S^s\left(\varepsilon\right) \mid \varepsilon > \varepsilon^s\right) > E^w\left(S^w\left(\varepsilon\right) \mid \varepsilon > \varepsilon^w\right)$. It follows from (18) that $\left(B^s + V^s\right) > \left(B^w + V^w\right)$. The seasonality of S_v^s implies a higher reservation value ε^s in the hot season s (the marginal house has to be of higher quality in order to generate a bigger surplus to the seller). The seasonality in sellers' and buyers' outside options, on the other hand, tends to reduce the cutoff ε^s in the hot season s. This is because the outside option in the hot season s is linked to the sum of values in the winter season: $B^w + V^w$. To see this negative effect more explicitly, rewrite (17) as

$$\begin{aligned} &\frac{1+\beta\phi^w}{1-\beta^2\phi^w\phi^s}\varepsilon^s \\ = \;& S_v^s + u + \frac{\beta\phi^w\left(1-\beta\right)\left(1+\beta\phi^s\right)}{1-\beta^2\phi^w\phi^s}\left(V^w + B^w\right) + \frac{\beta^2\phi^w\left(1-\phi^s\right)}{1-\beta^2\phi^w\phi^s}\left(V^w + B^w - V^s - B^s\right), \end{aligned} \tag{23}$$

which makes clear that $\left(B^s + V^s\right) > \left(B^w + V^w\right)$ has a negative effect on $\varepsilon^s/\varepsilon^w$. This gives rise to the following observations:

Remark 3 *In both a seller's and a buyer's market, transactions are seasonal. The seasonality of transactions is higher in a buyer's market.*

To see this, note that the outside option for both the buyer and the seller in the hot season is to wait and transact in the cold season. This makes both buyers and sellers less demanding in the hot season, yielding a larger number of transactions. In other words, the "counter-seasonality" in outside options increases the seasonality in transactions. On the other hand, when the seller sets prices, the surplus of the seller is higher in the hot season and hence sellers are more demanding and less willing to transact, which reduces the seasonality of transactions. Hence, in a seller's market, the seasonality of outside options and of the seller's surplus have opposite effects on the seasonality of reservation quality, causing a relatively lower degree of seasonality in transactions. In a buyer's market instead, only the seasonality of the outside options affects (positively) the degree of seasonality. Therefore, the seasonality of transactions is higher when the buyer sets prices. Finally, the effect of the rental flow u on the seasonality of transactions is as follows:

Remark 4 *In both the seller's and buyer's market, the extent of seasonality of transactions is decreasing in the rental flow u.*

The observation follows from the fact that the extent of the seasonality of outside options for buyers and sellers is decreasing in u (similar to the reasoning in Remark 2). Hence, as u increases, transactions become less seasonal.

We show in Appendix 8.4 that the existence of seasonality in transactions in the decentralized economy is an efficient outcome that is, the planner's solution also yields seasonality in transactions; this results naturally from the seasonality of moving probabilities. The actual extent of seasonality in both the seller's and buyer's markets, however, is inefficient, because the decentralized economy takes the stock of vacant houses as given and therefore ignores the thick market externality on the housing market. We clarify these points in Appendix 8.4.

5.2 Calibration of the model

5.2.1 Parameter values

We now calibrate the model to study its quantitative implications. We set the discount factor β so that the implied annual real interest rate is 5 percent.

We set the average probability of staying in the house $\phi = (\phi^s + \phi^w)/2$ to match survey data on the average duration of stay in a given house, which in the model is given by $\frac{1}{1-\phi}$). The median duration in the U.S. from 1993 through 2005, according to the American Housing Survey, was 18 semesters; the median duration in the U.K. during this period, according to the Survey of English Housing was 26 semesters. The implied (average) moving probabilities ϕ per semester are

0.056 and 0.039 for the US and the UK, respectively. These two surveys also report the main reasons for moving. Around $30-40$ percent of the respondents report that living closer to work or to their children's school and getting married are the main reasons for moving.[21] These factors are of course not entirely exogenous, but they can carry a considerably exogenous component; in particular, the school calendar is certainly exogenous to housing market movements (see Tucker, Long, and Marx (1995)'s study of seasonality in children's residential mobility.) In all, the survey evidence supports our working hypothesis that the *ex ante* probability to move is higher in the summer (or, equivalently the probability to stay is higher in the winter).

To illustrate the thick market effect, We assume that $F^j(.)$ follows a uniform distribution on the support $[0, v^j]$ (where v^j is endogenously determined) Intuitively, a hot season with a higher v^s is characterized by a housing market where the matching quality is better on average.

We calibrate the net rental flow received by the seller, u, to match the implied average (deseasonalized) rent-to-price ratio *received by the seller* in a seller's market.[22] In the UK, the average *gross* rent-to-price ratio is 5 percent per year, according to *Global Property Guide*.[23] The u/p ratio in our model corresponds to the *net* rental flow received by the seller after paying taxes and other relevant costs. It is accordingly lower than the *gross* rent-to-price ratio. As a benchmark, we choose u so that u/p is equal to 3 percent per year (equivalent to paying a 40 percent income tax on rent).[24] To do so, we use the equilibrium equations in the model without seasonality, that is, the model in which $\phi^s = \phi^w = \phi$. From (14), the equilibrium price in a seller's market in the

[21] Using monthly data on marriages from 1980 through 2003 for the U.K. and the U.S., we find that marriages are highly seasonal in both countries, with most marriages taking place between April and September. (The difference in annualized growth rates of marriages between the broadly defined "summer" and "winter" semesters are 200 percent in the U.S. and 400 percent in the U.K.). Results are available from the authors.

[22] Note that in a buyer's market, u/p is always constant, given by: $\frac{u}{p^s} = \frac{u}{p^w} = 1 - \beta = 0.024$ per semester (just below 5 percent per year).

[23] Data for the U.K. and other European countries can be found in http://www.globalpropertyguide.com/Europe/United-Kingdom/price-rent-ratio

[24] In principle, other costs can trim down the 3-percent u/p ratio, including maintenance costs, and inefficiencies in the rental market that lead to a higher wedge between what the tenant pays and what the landlord receives; also, it might not be possible to rent the house immediately, leading to lower average flows u. Note, however, that lower values of u/p lead to even higher seasonality in prices and transactions for any given level of seasonality in moving shocks. In that sense, lower u/p-ratios make it "easier" for our model to generate seasonality in prices.

absence of seasonality in moving probabilities is

$$p = \frac{u}{1-\beta} + \left(\frac{1-\beta F\left(\varepsilon^{d}\right)}{1-\beta}\right) S_v, \tag{24}$$

where the equilibrium reservation quality ε^d can be derived from (17) (see Appendix 8.2):

$$\frac{\varepsilon^{d}}{1-\beta\phi} = S_v + \frac{u + \frac{\beta\phi}{1-\beta\phi}\int_{\varepsilon^{d}}^{\bar{\varepsilon}} \varepsilon dF\left(\varepsilon\right)}{1-\beta\phi F\left(\varepsilon^{d}\right)}, \tag{25}$$

and, from (13), $S_v = \frac{1-F\left(\varepsilon^{d}\right)}{f\left(\varepsilon^{d}\right)(1-\beta\phi)}$. We substitute $u = 0.03 \cdot p$ and find the equilibrium value of p given the calibrated values for β and $F\left(.\right)$. The implied values for u are 0.13 and 0.11 when the average durations of stay are 9 years and 13 years, respectively.

Finally, we need to calibrate the extent of seasonality in the probability to move, $\left(1-\phi^{s}\right)/\left(1-\phi^{w}\right)$. Since there is no direct evidence on this—the available data can only help to pin down the average $\phi = \left(\phi^{s}+\phi^{w}\right)/2$—we report results for different values of $\left(1-\phi^{s}\right)/\left(1-\phi^{w}\right)$ ranging from 1.2 to 2 (which implies a range for the difference in moving probabilities, $\phi^{w}-\phi^{s}$, of 0.01 to 0.037 when the duration is 9 years and of 0.007 to 0.026 when the duration is 13 years).

5.2.2 Buyer's versus seller's market

Throughout the quantitative analysis that follows, we will maintain the working hypothesis that the U.S. behaves like a buyer's market, while the U.K. behaves like a seller's market. The question is of course whether this mapping is a good characterization of the U.S. and U.K. housing markets. There are at least five reasons why we think this is a reasonable approximation. First, population density in the U.K. (246 inhabitants per km^2) is 700 percent higher than in the U.S. (31 inhabitants per km^2), making land significantly scarcer relative to population in the U.K. and potentially conferring home owners more power in price setting.

Second, anecdotal evidence suggests that land use regulations are particularly stringent in the U.K. Indeed in its international comparison of housing markets, the OECD Economic Outlook 78 highlights the "complex and inefficient local zoning regulations and slow authorization process" in the U.K. economy, which the report cites as one of the reasons for the remarkable rigidity of housing supply.[25] Restrictions reinforce the market power of owners in practice by reducing the supply of houses.

[25] OECD Economic Outlook Number 78, chapter III (available at http://www.oecd.org/dataoecd/41/56/35756053.pdf)

Third, in the model, prices are higher in a seller's market than in a buyer's market for any given level of u, β, and ϕ^j, since the surplus to the seller is positive in (14) in a seller's market; also, because income is normalized to 1, this is equivalent to saying that price-to-income ratios are higher in a seller's market. In the data, housing prices relative to income are significantly higher in the U.K. According to the "*Demographia* International Housing Affordability Survey 2008," in the U.K. it takes more than 14 years of median household income to pay for the median priced house, while in the U.S. as a whole it takes only 8 years on average. The same survey finds that in 25 out of the 28 housing markets identified for the U.K., houses are severely unaffordable, with the "median multiple" (median house price divided by median household income) above 5.1, while in the U.S. only 30 out of the 129 housing markets identified by the survey are above 5.[26]

Fourth, the housing stock in the U.K. is significantly older than in the U.S. Over time, differential degrees of maintenance efforts across owners, architectural styles, and construction technologies, can lead to more substantial heterogeneity in unobservables in the U.K. housing stock. To the extent that this differentiation reduces the degree of substitutability across houses, it can potentially generate higher monopoly power for the seller.

The fifth and final reason for our working hypothesis is simply the failure of alternative channels to generate *price* seasonality in one country and not in the other, while at the same time generating seasonality in transactions in both countries. We have discussed most potential channels in the empirical Section and argued why they fail to explain the cross-country patterns. We examine further the role of seasonality in transaction costs in Section 6 and conclude that they fail to generate quantitatively meaningful levels of seasonality.

5.3 Quantitative Results

We present the baseline results in Tables 16 and 17. Table 16 shows the model-generated seasonality in prices and transactions assuming an average stay in the house of 9 years—the average stay in the U.S. Column (1) shows the ratio of moving probabilities, $(1-\phi^s)/(1-\phi^w)$, and Column (2) shows the implied difference in probabilities, $\phi^w - \phi^s$, across the two seasons; Columnns (3) through (6) show correspondingly the differences across seasons in annualized growth rates for

[26] Of course, international comparisons are not free of problems; for example, it may well be that the U.K. housing markets provide more amenities, which are then reflected in higher housing prices; we accordingly read this evidence as suggestive, rather than conclusive.

prices and transactions in both a seller's and a buyer's market, which can be directly compared to the statistics reported in the empirical Section. Table 17 shows the corresponding figures when the average stay is 13 years—the average in the U.K.

The Tables illustrate the analytical results derived earlier. First, for all values of $(1-\phi^s)/(1-\phi^w)$, housing prices display seasonality in a seller's market but no seasonality in a buyer's market. Second, for all values of $(1-\phi^s)/(1-\phi^w)$, seasonality in transactions is significantly larger in a buyer's market than in a seller's market. Indeed, small differences in moving probability across seasons are substantially amplified in a buyer's market.

The question we address in this Section is whether the model can quantitatively mimic the empirical patterns described before. The answer is positive, under our conjecture that the U.S. housing market behaves more like the buyer's market in our model and the U.K. housing market behaves like the seller's market. (See our justification for this conjecture in Section 5.2.2.) Under this characterization, it is clear from Table 16 (where the average stay corresponds to that in the U.S.), that a buyer's market can yield a-seasonal prices (column 5), while at the same time yielding considerably high levels of seasonality in transactions (column 6), comparable to those found in U.S. data (see Table 9), when the ratio of moving probabilities is 1.4, or, equivalently, when the difference in moving probabilities across seasons is 0.019. Furthermore, as it is apparent from Table 17 (where the average stay matches that in the U.K.), a seller's market can generate seasonality in prices comparable to that observed in the U.K. (column 3), when the ratio of moving probabilites is 2. This number might seem at first large but it simply requires that the probability of moving in the summer be higher than the probability of moving by 0.026; in other words, this implies that there are 2.6 percent more houses-homeowners matches subject to a moving shock in the summer than there are in the winter; in this range also, the model generates differences in annual growth rates of transactions of around 104 percent, just below the numbers observed for the U.K. economy (see Table 5).

Table 16. Model-Generated Seasonality for an Average Stay of 9 years (U.S.)

		Average moving probability: 0.0556			
		Seller's Market		*Buyer's Market*	
		Differences in Annualized Growth Rates		*Differences in Annualized Growth Rates*	
ratio of moving probabilities across seasons (1)	**difference in moving probabilities across seasons (2)**	**Prices (3)**	**Transactions (4)**	**Prices (5)**	**Transactions (6)**
1.2	0.010	1.6%	27.7%	0.0%	90.3%
1.4	0.019	2.9%	51.2%	0.0%	179.3%
1.6	0.026	4.0%	71.6%	0.0%	278.0%
1.8	0.032	4.9%	89.7%	0.0%	389.7%
2.0	0.037	5.7%	106.0%	0.0%	518.4%

Table 17. Model-Generated Seasonality for an Average Stay of 13 years (U.K.).

		Average moving probability: 0.0385			
		Seller's Market		*Buyer's Market*	
		Differences in Annualized Growth Rates		*Differences in Annualized Growth Rates*	
ratio of moving probabilities across seasons (1)	**difference in moving probabilities across seasons (2)**	**Prices (3)**	**Transactions (4)**	**Prices (5)**	**Transactions (6)**
1.2	0.007	1.7%	27.3%	0.0%	82.8%
1.4	0.013	3.1%	50.5%	0.0%	176.0%
1.6	0.018	4.2%	70.6%	0.0%	271.0%
1.8	0.022	5.3%	88.5%	0.0%	378.2%
2.0	0.026	6.1%	104.5%	0.0%	500.9%

As we argued in Section 3, matching the seasonal features of housing markets is not straightforward in models with standard asset-pricing conditions. Recall that, in that framework, we needed percent differences in depreciation rates and repair costs across seasons of the order of 400 percent or higher to match U.K. price seasonality. And even if one were willing to (heroically) assume that those costs were indeed so different across seasons, there would still remain the puzzle of why these differences were not present in the U.S.

Though stylised, our model can mimic the empirical seasonal patterns in both the U.S. and the U.K. for reasonable parameter values. In particular, it can yield considerable amplification in the seasonality of transactions in a buyer's market, without leading to seasonality in prices, a feature consistent with U.S. data, and at the same time generate seasonality in both transactions and prices comparable to those observed in U.K. data.

6 Transaction Costs

It is interesting to ask whether and how the introduction of transaction costs would alter the patterns of seasonality in prices and transactions obtained in the baseline model. We hence extend our model to allow for transaction costs associated to the purchase (or sale) of a house for both the buyers and sellers in seasons $j = s, w$:

$$\begin{aligned} T_b^j\left(p^j\right) &= \bar{\tau}_b^j + \tau_b p^j; \\ T_v^j\left(p^j\right) &= \bar{\tau}_v^j + \tau_v p^j, \end{aligned}$$

where $T_b^j\left(p^j\right)$ is the transaction cost paid by the buyer in season j and $T_v^j\left(p^j\right)$ is the corresponding cost paid by the seller. We allow the fixed-cost components, $\bar{\tau}_b^j$ and $\bar{\tau}_v^j$, such as moving costs and repairing costs, to be seasonal.[27] The proportional components, τ_b and τ_v, such as estate agents' fees or taxes, are (realistically) assumed to be a-seasonal.

We show in Appendix 8.5 that the equilibrium price equation (14) stills holds by simply replacing p^s with $p^s - T_v\left(p^s\right)$, the net price received by the seller:

$$p^s - T_v^s\left(p^s\right) = \frac{u}{1-\beta} + \frac{1-\beta^2 F^s\left(\varepsilon^s\right)}{1-\beta^2} S_v^s + \frac{\beta\left[1-F^w\left(\varepsilon^w\right)\right]}{\left(1-\beta^2\right)} S_v^w, \tag{26}$$

where the surplus S_v^s in the seller's market (13) is now multiplied by $\frac{1-\tau_v}{1+\tau_b}$, which is analogous to the "tax wedge" applied to a match between a firm and a worker in the labour economics literature:

$$S_v^s = \left(\frac{1-\tau_v}{1+\tau_b}\right)\left(\frac{1-F^s\left(\varepsilon^s\right)}{f^s\left(\varepsilon^s\right)}\right)\frac{1+\beta\phi^w}{1-\beta^2\phi^w\phi^s}. \tag{27}$$

The reservation-quality equation (17) also holds by including the total transaction costs $T^s\left(p^s\right) = T_b^s\left(p^s\right) + T_v^s\left(p^s\right)$ on the right-hand side:

$$\frac{1+\beta\phi^w}{1-\beta^2\phi^w\phi^s}\varepsilon^s = S_v^s + u + T^s\left(p^s\right) + \frac{\beta\phi^w\left(1-\beta^2\phi^s\right)}{1-\beta^2\phi^w\phi^s}\left(V^w + B^w\right) - \frac{\beta^2\phi^w\left(1-\phi^s\right)}{1-\beta^2\phi^w\phi^s}\left(V^s + B^s\right). \tag{28}$$

Finally, $S_b^s\left(\varepsilon\right)$ and $\left(B^s + V^s\right)$ remain exactly as in (15) and (18) in the baseline model.

There are two interesting observations. First, for small enough proportional costs, $\left(1-\tau_v\right)/\left(1+\tau_b\right) \simeq 1-\left(\tau_v + \tau_b\right)$, the modified surplus to the seller, S_v^s, depends (to a first approximation) on the sum

[27] Repair costs (both for the seller who's trying to make the house more attractive and for the buyer who wants to adapt it before moving in) may be smaller in the summer because good weather and the opportunity cost of time (assuming vacation is taken in the summer) are important inputs in construction). Moving costs, similarly, might be lower during vacation (both job and school holidays).

of proportional costs $(\tau_v + \tau_b)$ and on ε^s. From the modified equations (15) and (17), ε^s depends on total costs only. Hence, from (19), in equilibrium, the number of vacant houses, v^s, depends only on total costs. It follows that also the extent of seasonality in the number of transactions depends only on total costs.

Second, the modified price equation (26) shows that the extent of seasonality in prices depends not only on total costs but also on how these costs are distributed between buyers and sellers. Specifically, seasonality in prices in the seller's market (barring seasonality in the fixed cost itself) increases when the buyer bears most of the fixed cost of the transaction.

From a quantitative standpoint, however, transaction costs do not significantly alter the extent of seasonality documented in the baseline model, where seasonality is measured (as before) as the difference in annualized growth rates of prices or transactions across seasons. We study the quantitative implications of transaction costs in various steps. Table 18 explores the role of *proportional* transaction costs for different degrees of seasonality in moving shocks. Using the same calibrated parameters as before, for an average stay of 9 years (or average moving probability $\phi = 0.0556$), we assume that both the buyer and the seller pay a transaction cost of 3 percent of the property price, which is consistent with the averages reported in various guides for home buying and selling in both the U.S. and the U.K. That is, we set $\tau_v = \tau_b$ equal to 3 percent of the property price. (Note that the only difference with Table 16 is that in there we set $\tau_v = \tau_b = 0$). Though proportional transaction costs increase seasonality, quantitatively, the increase is only slight and the results are very close to those in Table 16. The results when the average stay is 13 years (not reported for the sake of brevity) are equally insensitive (from a quantitative standpoint) to the introduction of proportional costs of this size.

Table 18. Model-Generated Seasonality with Proportional Costs

	Average moving probability: 0.0556			
ratio of moving probabilities across seasons (1)	***Seller's Market***		***Buyer's Market***	
	Differences in Annualized Growth Rates		***Differences in Annualized Growth Rates***	
	Prices (2)	**Transactions (3)**	**Prices (4)**	**Transactions (5)**
1.2	1.7%	28.9%	0.0%	92.1%
1.4	3.0%	53.5%	0.0%	184.1%
1.6	4.2%	75.0%	0.0%	285.9%
1.8	5.2%	94.0%	0.0%	402.9%
2.0	6.0%	111.6%	0.0%	539.2%

Tables 19 through 21 explore the role of fixed transaction costs. As said, both the sum of the fixed costs paid by the seller and the buyer, as well as its distribution between the two parties matter for seasonality. To make this distinction clear, we let $\bar{\tau}_b = \theta \cdot \bar{\tau}$ and $\bar{\tau}_v = (1-\theta) \cdot \bar{\tau}$. (Note that we are for the moment ignoring any seasonality in fixed costs, a point to which we come back later.) Given that there is no systematic empirical evidence on the size of fixed costs, we explore a relatively large range for $\bar{\tau}$, equivalent to 0 through to 10 percent of the ex-post value of the house. We then study the degree of seasonality generated by the model for different fixed costs and different ratios of moving probabilities across seasons. Table 19 assumes the fixed cost is equal for both sellers and buyers, that is, $\theta = 0.50$. All other parameters are calibrated as in Table 16. As Table 19 shows, increasing the fixed cost from 0 to 10 percent, increases price and transaction seasonality in a seller's market only slightly; seasonality in transactions in the buyer's market appears to be more sensitive, but only when the moving shock is highly seasonal (in the Table, when the ratio of moving probabilities is 2).

Table 19. Model-Generated Seasonality with Fixed Costs and $\theta = 0.50$

Average moving probability: 0.0556				
	Seller's Market		***Buyer's Market***	
	Differences in Annualized Growth Rates		***Differences in Annualized Growth Rates***	
Fixed-cost-to-price ratio (1)	**Prices (2)**	**Transactions (3)**	**Prices (4)**	**Transactions (5)**
ratio of moving probabilities across seasons = 1.4				
0%	2.9%	51.2%	0.0%	179.9%
5%	3.0%	53.2%	0.0%	184.7%
10%	3.1%	55.3%	0.0%	190.2%
ratio of moving probabilities across seasons = 1.8				
0%	4.9%	89.7%	0.0%	389.7%
5%	5.1%	93.4%	0.0%	405.0%
10%	5.3%	97.2%	0.0%	422.5%
ratio of moving probabilities across seasons =2.0				
0%	5.7%	106.0%	0.0%	518.4%
5%	6.0%	110.5%	0.0%	542.7%
10%	6.2%	115.0%	0.0%	570.7%

Table 20 reports the results from the same exercise carried out in Table 19, with the only difference that we now assume that only the buyer pays a fixed cost, that is, $\theta = 1$. As observed while discussing the analytical results, the extent of seasonality in transactions is virtually unaffected by the distribution of the total fixed cost between the two parties. Hence, the figures for seasonality in transactions are almost identical to the corresponding ones in Table 19.[28] Seasonality in prices, though, increases when the buyer pays the fixed cost, with the increase being more substantial when the moving shock is more seasonal (that is, for higher values of the ratio of moving probabilities).

Table 20. Model-Generated Seasonality with Fixed Costs and $\theta = 1$

Average moving probability: 0.0556				
Fixed-cost-to-price ratio (1)	***Seller's Market*** *Differences in Annualized Growth Rates*		***Buyer's Market*** *Differences in Annualized Growth Rates*	
	Prices (2)	**Transactions (3)**	**Prices (4)**	**Transactions (5)**
ratio of moving probabilities across seasons = 1.4				
0%	2.9%	51.2%	0.0%	179.9%
5%	3.2%	53.2%	0.0%	185.3%
10%	3.4%	55.0%	0.0%	190.5%
ratio of moving probabilities across seasons = 1.8				
0%	4.9%	89.7%	0.0%	389.7%
5%	5.4%	93.4%	0.0%	406.7%
10%	5.8%	96.8%	0.0%	423.3%
ratio of moving probabilities across seasons =2.0				
0%	5.7%	106.0%	0.0%	518.4%
5%	6.3%	110.4%	0.0%	545.2%
10%	6.8%	114.5%	0.0%	571.8%

Table 21 reports the corresponding results when only the seller pays the fixed cost of the transaction, that is, $\theta = 0$. As discussed earlier, seasonality in transactions is unaltered (compared the corresponding figures in Tables 19 and 20), while seasonality in prices decreases *vis-à-vis* the cases in which the buyer pays some or all of the fixed cost. Note also that seasonality in prices in a seller's market decreases (though slightly) with the size of the fixed cost. This is because the fixed cost (paid by the sellers) is an a-seasonal component in the price equation (26), which reduces the extent of seasonality in prices generated by the surplus terms. The conclusions drawn from Tables 19 through 21 are unchanged when the average stay is 13 rather than 9 years. (The results, not reported for the sake of space, are available from the authors).

[28] The analytical result shows that the extent of seasonality in transactions depends only on the level of total costs. Note that Tables $19 - 21$ report the results when the fixed-cost-to-price ratio are equal (varying from 5 to 10 percent). Therefore, the figures in column (5) of Tables $19 - 21$ are not exactly identical because the level of prices vary under the alternative settings of $\theta = 0; 0.5; 1.$ and hence so do the implied calibrated levels of the total cost ($\bar{\tau}$)

Table 21. Model-Generated Seasonality with Fixed Costs and $\theta = 0$

Average moving probability: 0.0556				
Fixed-cost-to-price ratio (1)	***Seller's Market*** *Differences in Annualized Growth Rates*		***Buyer's Market*** *Differences in Annualized Growth Rates*	
	Prices (2)	**Transactions (3)**	**Prices (4)**	**Transactions (5)**
ratio of moving probabilities across seasons = 1.4				
0%	2.9%	51.2%	0.0%	179.9%
5%	2.8%	53.2%	0.0%	184.2%
10%	2.8%	55.6%	0.0%	190.1%
ratio of moving probabilities across seasons = 1.8				
0%	4.9%	89.7%	0.0%	389.7%
5%	4.9%	93.5%	0.0%	403.4%
10%	4.8%	97.8%	0.0%	422.4%
ratio of moving probabilities across seasons =2.0				
0%	5.7%	106.0%	0.0%	518.4%
5%	5.7%	110.5%	0.0%	540.2%
10%	5.6%	115.8%	0.0%	570.7%

Finally, it is of interest to explore the quantitative effects of seasonality in the fixed cost itself and whether this seasonality might be the underlying driver of seasonality in housing markets. As expressed before, it might be natural to think that certain costs, such as repair or moving costs are lower in the summer (due to good weather and vacation time). How much lower, it is hard to point out, given the paucity of empirical evidence on the subject; for this reason, we study a range of possibilities, with relative total fixed costs in winters and summers, $\frac{\bar{\tau}^w}{\bar{\tau}^s}$, ranging from 1.2 through 2. We fix the average fixed cost (over summer and winter, $\frac{\bar{\tau}^w+\bar{\tau}^s}{2}$) at 5 percent and assume that this cost is shared equally between the buyer and seller (that is, $\theta = 0.5$).

To focus only on the effect of seasonality in fixed costs, we calibrate the moving probabilities to be equal across seasons $\phi^s = \phi^w$ and assume as before an average stay of 9 years. All other parameters are calibrated as in Table 16. The results from this exercise, displayed in Table 22 show that in a seller's market, seasonality in costs can generate very little seasonality in prices, while generating significant seasonality in transactions. Moreover, and perhaps surprisingly, in a buyer's market, seasonality in prices is reverted, while seasonality in transactions becomes extremely large. This result for prices follows immediately from the fact that the only seasonal component in the price equation (26) for a buyer's market is the cost term; hence, a higher cost in the winter implies a higher price, which in turns implies an even lower number of transactions in the winter. The results then suggest that it is very unlikely that seasonality in moving or repair costs could be the triggering force for the empirical findings documented in this paper.

Table 22. Model-Generated Seasonality with Seasonal Fixed Costs, $\phi^s = \phi^w$, and $\theta = 0.5$

Average moving probability: 0.0556				
ratio of winter to summer fixed costs (1)	Seller's Market — Differences in Annualized Growth Rates		Buyer's Market — Differences in Annualized Growth Rates	
	Prices (2)	Transactions (3)	Prices (4)	Transactions (5)
1.2	0.6%	46.0%	-2.8%	137.0%
1.4	1.1%	86.2%	-5.2%	301.5%
1.6	1.4%	123.1%	-7.1%	534.6%
1.8	1.8%	157.7%	-8.8%	885.9%
2.0	2.1%	190.7%	-10.3%	1433.6%

7 Concluding Remarks

This paper documents seasonal booms and busts in housing markets and argues that the predictability and high extent of seasonality in prices observed in some of them cannot be quantitatively reconciled with standard asset-pricing equilibrium conditions.

To explain the empirical patterns, the paper presents a search model that can quantitatively account for most of the empirical puzzle. As a by product, the model sheds new light on interesting mechanisms governing fluctuations in housing markets that can potentially be useful in a study of lower-frequency movements. In particular, the model highlights the roles of thick-market externalities and the distribution of power between buyers and sellers as the key determinants of housing markets' behavior.

In future work, the authors plan to adapt the model presented in the paper to study lower frequency movements in the housing markets.

8 Appendix

8.1 Derivation for the model with seasons

Deriving the optimal price (14):

Using the definition of $S_v^s \equiv p^s - \beta V^w - u$, rewrite the value function of the seller (11) as

$$V^s = \beta V^w + u + [1 - F^s(\varepsilon^s)] S_v^s.$$

Solving out V^s explicitly:

$$V^s = \frac{1 - F^s(\varepsilon^s)}{1-\beta^2} S_v^s + \frac{\beta\left[1 - F^w(\varepsilon^w)\right]}{1-\beta^2} S_v^w + \frac{u}{1-\beta}, \tag{29}$$

where S_v^s is solved in (13). Substituting (29) into the definition of S_v^s, we obtain (14).

Deriving $B+V$ in (18):

By definition of optimal price p^s and reservation quality ε^s, the value function (8) becomes:

$$B^s = \int_{\varepsilon^s}^{\bar{\varepsilon}^s} \left(H^s(\varepsilon) - p^s\right) dF^s(\varepsilon) + F^s(\varepsilon^s)\beta B^w$$

which can be rewritten as

$$B^s = \beta B^w + \left(1 - F^s(\varepsilon^s)\right) E^s\left[S_b^s(\varepsilon) \mid \varepsilon > \varepsilon^s\right],$$

where S_b^s is solved in (15). Solving out B^s explicitly:

$$B^s = \frac{1 - F^s(\varepsilon^s)}{1-\beta^2} E^s\left[S_b^s(\varepsilon) \mid \varepsilon > \varepsilon^s\right] + \frac{\beta\left[1 - F^w(\varepsilon^w)\right]}{1-\beta^2} E^w\left[S_b^w(\varepsilon) \mid \varepsilon > \varepsilon^w\right] \tag{30}$$

Combining (30) with (29), we derive (18).

8.2 The model without seasons

The value functions for the model without seasonality are identical to those in the model with seasonality without the superscripts s and w. It can be shown that the equilibrium equations are also identical by simply setting $\phi^s = \phi^w$. Using (17),

$$\frac{\varepsilon^d}{1-\beta\phi} = S_v + u + \frac{\beta\phi}{1-\beta\phi}(1-\beta)(V+B)$$

where S_v follows from (13),

$$S_v = \frac{1 - F\left(\varepsilon^d\right)}{f\left(\varepsilon^d\right)(1-\beta\phi)}.$$

and $B+V$ from (18),

$$\begin{aligned} B+V &= \frac{u}{1-\beta} + \frac{1 - F\left(\varepsilon^d\right)}{1-\beta}\left\{\frac{\left[E\left(\varepsilon - \varepsilon^d \mid \varepsilon > \varepsilon^d\right)\right]}{1-\beta\phi} + S_v\right\} \\ &= \frac{1}{1-\beta}\left[u + \left[1 - F\left(\varepsilon^d\right)\right] S_v + \int_{\varepsilon^d}^{\bar{\varepsilon}} \left(\frac{\varepsilon - \varepsilon^d}{1-\beta\phi}\right) dF(\varepsilon)\right]. \end{aligned}$$

substitute into the reservation equation,

$$\frac{\varepsilon^d}{1-\beta\phi} = S_v + u + \frac{\beta\phi}{1-\beta\phi}\left[u + \left[1 - F\left(\varepsilon^d\right)\right] S_v + \int_{\varepsilon^d}^{\bar{\varepsilon}} \left(\frac{\varepsilon - \varepsilon^d}{1-\beta\phi}\right) dF(\varepsilon)\right]$$

which simplifies to (25).

8.3 Efficiency

This Section discusses the efficiency of equilibrium in the decentralized economy under both a seller's and a buyer's market scenarios. The planner observes the match quality ε and is subject to the same exogenous moving shocks that hit the decentralized economy. The interesting comparison is the level of reservation quality achieved by the planner with the corresponding levels in a seller's and a buyer's market.

To spell out the planner's problem, we follow Pissarides (2000) and assume that in any period t the planner takes as given the expected value of the housing utility service per person in period t (before he optimizes), which we denote by h_{t-1}, as well as the beginning of period's stock of vacant houses, v_t. Thus, taking as given the initial levels h_{-1} and v_0, and the sequence $\{\phi_t\}_{t=0...}$, which alternates between ϕ^j and $\phi^{j'}$ for seasons $j, j' = s, w$, the planner's problem is to choose a sequence of $\{\varepsilon_t\}_{t=0,..}$ to maximize

$$U\left(\{\varepsilon_t, h_t, v_t\}_{t=0...}\right) \equiv \sum_{t=0}^{\infty} \beta^t \left[h_t + uv_t F\left(\varepsilon_t; v_t\right)\right] \tag{31}$$

subject to the law of motion for h_t :

$$h_t = \phi_t h_{t-1} + v_t \int_{\varepsilon_t}^{\bar{\varepsilon}(v_t)} x dF\left(x; v_t\right), \tag{32}$$

the law of motion for v_t (which is similar to the one in the decentralized economy):

$$v_{t+1} = v_t \phi_{t+1} F\left(\varepsilon_t; v_t\right) + 1 - \phi_{t+1}, \tag{33}$$

and the inequality constraint:

$$0 \leqslant \varepsilon_t \leqslant \bar{\varepsilon}\left(v_t\right). \tag{34}$$

Intuitively, the planner faces two types of trade-offs when deciding the optimal reservation quality ε_t : a static one and a dynamic one. The static trade-off stems from the comparison of utility values generated by occupied houses and vacant houses in period t in the objective function (31). The utility per person generated from vacant houses is the rental income per person, captured by $uv_t F\left(\varepsilon_t\right)$. The utility generated by occupied houses in period t is captured by h_t, the expected housing utility service per person conditional on the reservation value ε_t set by the planner in period t. The utility h_t, which follows the law of motion (32), is the sum of the pre-existing expected housing utility h_{t-1} that survives the moving shocks and the expected housing utility from the new matches. By increasing ε_t, the expected housing value h_t decreases, while the utility generated by vacant houses increases (since $F(\varepsilon_t)$ increases). The dynamic trade-off operates through the law of motion for the stock of vacant houses in (33). By increasing ε_t (which in turn decreases h_t), the number of transactions in the current period decreases; this leads to more vancant houses in the following period, v_{t+1}, and consequently to a thicker market in the next period.

Assuming the inequality constraints are not binding, i.e. markets are open in both the cold and hot seasons, the optimal reservation quality, ε^j, $j = s, w$, in the periordic steady state is (see Appendix 8.4):

$$\varepsilon^j \left(\frac{1+\beta\phi^{j'}}{1-\beta^2\phi^j\phi^{j'}} \right) = \frac{\left(1+\beta\phi^{j'}A^{j'}\right)u + \beta^2\phi^j\phi^{j'}A^{j'}D^j + \beta\phi^{j'}D^{j'}}{1-\beta^2\phi^j\phi^{j'}A^jA^{j'}}, \tag{35}$$

where

$$A^j \equiv F^j\left(\varepsilon^j\right) - v^jT_1^j; \qquad D^j \equiv \frac{1+\beta\phi^{j'}}{1-\beta^2\phi^j\phi^{j'}} \left(\int_{\varepsilon^j}^{\bar{\varepsilon}^j} \varepsilon dF^j(\varepsilon) + v^jT_2^j \right), \tag{36}$$

and the stock of vacant houses, v^j, $j = s, w$, satisfies (19) as in the decentralized economy.

The thick-market effect enters through two terms: $T_1^j \equiv \frac{\partial}{\partial v^j}\left[1 - F^j\left(\varepsilon^j\right)\right] > 0$ and $T_2^j \equiv \frac{\partial}{\partial v^j}\int_{\varepsilon^j}^{\bar{\varepsilon}^j} \varepsilon dF^j(\varepsilon) > 0$. The first term, T_1^j, indicates that the thick-market effect shifts up the acceptance schedule $\left[1 - F^j(\varepsilon)\right]$. The second term, T_2^j, indicates that the thick-market effect increases the conditional quality of transactions. The interior solution (35) is an implicit function of ε^j that depends on $\varepsilon^{j'}$, v^j, and $v^{j'}$. It is not straightforward to derive an explicit condition for $\varepsilon^j < v^j$, $j = s, w$. However, when there are no seasons, $\phi^s = \phi^w$, it follows immediately from (19) that the solution is interior, $\varepsilon < v$. On the other hand, when the exogenous difference in moving propensities across seasons is large enough, the Planner might find it optimal to close down the market in the cold season. Before we turn to such situation, it is helpful to understand the sources of inefficiency in the decentralized economy when there are no seasons.

Abstracting from seasonality for the moment, there are two sources of inefficiency in the decentralized economy. First, the match quality ε is private information: Only buyers observe it. This implies that the number of transactions in a seller's market is inefficiently low. Second, the optimal decision rules of buyers and sellers take the stock of houses in each period as given, thereby ignoring the effects of their decision rules on the stock of vacant houses in the following periods. The thick-market effect generates a negative externality that makes the number of transactions in the decentralized economy (both in a seller's and a buyer's market) inefficiently high for any given stock of vacant houses. More specifically, setting $\phi^s = \phi^w = \phi$ in (35) implies the planner's optimal reservation quality ε^p satisfies:

$$\frac{\varepsilon^p}{1-\beta\phi} = \frac{u + \frac{\beta\phi}{1-\beta\phi}\left(\int_{\varepsilon^p}^{\bar{\varepsilon}} \varepsilon dF(\varepsilon) + vT_2\right)}{1 - \beta\phi F(\varepsilon^p) + \beta\phi v T_1}. \tag{37}$$

The two sources of inefficiency can now be seen explicitly by comparing (37) with (25). The positive term S_v affecting ε^d in the decentralized seller's market increases the reservation quality and hence lowers the number of transactions with respect to the efficient (Planner's) outcome. This source of inefficiency disappears in a buyer's market, since $S_v = 0$. The second source of inefficiency, which operates through the thick-market externality, is present in both sellers' and buyers' markets. The thick-market effect, captured by $T_1$and T_2, generates two opposite forces. The term T_1 decreases ε^p, while the term T_2 increases ε^p in the planner's solution. Thus, the positive thick-market effect on the acceptance rate T_1 implies that the number of transactions is too low in the decentralized economy, while the positive effect on quality T_2 implies that the number of transactions is too high. Since $1 - \beta\phi$ is close to zero, however, the term T_2 dominates. Therefore, the overall effect of the thick-market externality is to increase the number of transactions in the decentralized economy relative to the efficient outcome.[29] Hence, the number of transactions in a buyer's market is too high compared to the planner's solution, while in a seller's market it can be too low or too high, depending ultimately on the shape of the distribution $F(.)$.

We now return to the planner's problem in the case in which it is optimal to close down the market during the cold season. In this case, the solution implies setting $\varepsilon_t^w = \bar{\varepsilon}_t^w$ in the planner's problem. The optimal reservation quality, ε^s, in the periodic steady state is (see Appendix 8.4):

$$\frac{\varepsilon^s}{1-\beta^2\phi^w\phi^s} = \frac{u + \frac{\beta^2\phi^w\phi^s}{1-\beta^2\phi^w\phi^s}\left(\int_{\varepsilon^s}^{\bar{\varepsilon}^s} \varepsilon dF^s(\varepsilon) + v^s T_2^s\right)}{1 - \beta^2\phi^s\phi^w\left[F^s(\varepsilon^s) - v^s T_1^s\right]}, \tag{38}$$

which is similar to the Planner's solution with no seasons in (37) with $\beta^2\phi^w\phi^s$ replacing $\beta\phi$.

8.4 Analytical derivations of the planner's solution

The Planner's solution when the housing market is open in all seasons

Because the sequence $\{\phi_t\}_{t=0,\ldots}$alternates between ϕ^j and $\phi^{j'}$ for seasons $j, j' = s, w$, the planner's problem can be written recursively. Taking (h_{t-1}, v_t), and $\{\phi_t\}_{t=0,..}$ as given, and provided that

[29] This result is similar to that in the stochastic job matching model of Pissarides (2000, chapter 8), where the reservation productivity is too low compared to the efficient outcome in the presence of search externalities.

the solution is interior, that is, $\varepsilon_t < v_t$, the Bellman equation for the planner is given by:

$$\begin{aligned} W\left(h_{t-1}, v_t, \phi_t\right) &= \max_{\varepsilon_t} \left[h_t + u v_t F\left(\varepsilon_t; v_t\right) + \beta W\left(h_t, v_{t+1}, \phi_{t+1}\right)\right] \\ & s.t. \\ h_t &= \phi_t h_{t-1} + v_t \int_{\varepsilon_t}^{\bar{\varepsilon}(v_t)} x dF\left(x; v_t\right), \\ v_{t+1} &= v_t \phi_{t+1} F\left(\varepsilon_t; v_t\right) + 1 - \phi_{t+1}. \end{aligned} \tag{39}$$

The first-order condition implies

$$\left(1 + \beta \frac{\partial W\left(h_t, v_{t+1}, \phi_{t+1}\right)}{\partial h_t}\right) v_t \left(-\varepsilon_t f\left(\varepsilon_t; v_t\right)\right) + \left(\beta \phi_{t+1} \frac{\partial W\left(h_t, v_{t+1}, \phi_{t+1}\right)}{\partial v_{t+1}} + u\right) v_t f\left(\varepsilon_t; v_t\right) = 0,$$

which simplifies to

$$\varepsilon_t \left(1 + \beta \frac{\partial W\left(h_t, v_{t+1}, \phi_{t+1}\right)}{\partial h_t}\right) = u + \beta \phi_{t+1} \frac{\partial W\left(h_t, v_{t+1}, \phi_{t+1}\right)}{\partial v_{t+1}}. \tag{40}$$

Using the envelope-theorem conditions, we obtain:

$$\frac{\partial W\left(h_{t-1}, v_t, \phi_t\right)}{\partial h_{t-1}} = \phi_t \left(1 + \beta \frac{\partial W\left(h_t, v_{t+1}, \phi_{t+1}\right)}{\partial h_t}\right) \tag{41}$$

and

$$\begin{aligned} \frac{\partial W\left(h_{t-1}, v_t, \phi_t\right)}{\partial v_t} &= \left(u + \beta \phi_{t+1} \frac{\partial W\left(h_t, v_{t+1}, \phi_{t+1}\right)}{\partial v_{t+1}}\right) \left(F\left(\varepsilon_t; v_t\right) - v_t T_{1t}\right) \\ &+ \left(1 + \beta \frac{\partial W\left(h_t, v_{t+1}, \phi_{t+1}\right)}{\partial h_t}\right) \left(\int_{\varepsilon_t}^{\bar{\varepsilon}(v_t)} x dF\left(x; v_t\right) + v_t T_{2t}\right) \end{aligned} \tag{42}$$

where $T_{1t} \equiv \frac{\partial}{\partial v_t}\left[1 - F\left(\varepsilon_t; v_t\right)\right] > 0$ and $T_{2t} \equiv \frac{\partial}{\partial v_t} \int_{\varepsilon_t}^{\bar{\varepsilon}(v_t)} x dF\left(x; v_t\right) > 0$.

In the periodic steady state, the first order condition (40) becomes

$$\varepsilon^j \left(1 + \beta \frac{\partial W^{j'}\left(h^j, v^{j'}\right)}{\partial h^j}\right) = u + \beta \phi^{j'} \frac{\partial W^{j'}\left(h^j, v^{j'}\right)}{\partial v^{j'}} \tag{43}$$

The envelope condition (41) implies

$$\frac{\partial W^j\left(h^{j'}, v^j\right)}{\partial h^{j'}} = \phi^j \left[1 + \beta \left(\phi^{j'} + \beta \phi^{j'} \frac{\partial W^j\left(h^{j'}, v^j\right)}{\partial h^{j'}}\right)\right]$$

which yields:

$$\frac{\partial W^j\left(h^{j'}, v^j\right)}{\partial h^{j'}} = \frac{\phi^j \left(1 + \beta \phi^{j'}\right)}{1 - \beta^2 \phi^j \phi^{j'}} \tag{44}$$

Substituting this last expression into (42), we obtain:

$$\frac{\partial W^j\left(h^{j'}, v^j\right)}{\partial v^j} = \left(u + \beta\phi^{j'}\frac{\partial W^{j'}\left(h^j, v^{j'}\right)}{\partial v^{j'}}\right) A^j + D^j,$$

where

$$A^j \equiv F^j\left(\varepsilon^j\right) - v^j T_1^j; \qquad D^j \equiv \frac{1+\beta\phi^{j'}}{1-\beta^2\phi^j\phi^{j'}}\left(\int_{\varepsilon^j}^{\bar{\varepsilon}^j} \varepsilon dF^j\left(\varepsilon\right) + v^j T_2^j\right).$$

Hence, we have

$$\frac{\partial W^j\left(h^{j'}, v^j\right)}{\partial v^j} = \left\{u + \beta\phi^{j'}\left[\left(u + \beta\phi^j\frac{\partial W^j\left(h^{j'}, v^j\right)}{\partial v^j}\right) A^{j'} + D^{j'}\right]\right\} A^j + D^j,$$

which implies

$$\frac{\partial W^j\left(h^{j'}, v^j\right)}{\partial v^j} = \frac{uA^j\left(1+\beta\phi^{j'}A^{j'}\right) + \beta\phi^{j'}D^{j'}A^j + D^j}{1-\beta^2\phi^j\phi^{j'}A^jA^{j'}}. \tag{45}$$

Substituting (44) and (45) into the first-order condition (43), we get:

$$\varepsilon^j\left(1+\beta\frac{\phi^{j'}\left(1+\beta\phi^j\right)}{1-\beta^2\phi^j\phi^{j'}}\right) = u + \beta\phi^{j'}\frac{uA^{j'}\left(1+\beta\phi^jA^j\right) + \beta\phi^jD^jA^{j'} + D^{j'}}{1-\beta^2\phi^j\phi^{j'}A^jA^{j'}}$$

simplify to (35).

The Planner's solution when the housing market is closed in the cold season

Setting $\varepsilon_t^w = \bar{\varepsilon}_t^w$, the Bellman equation (39) can be rewritten as:

$$\begin{aligned} W^s\left(h_{t-1}^w, v_t^s\right) &= \max_{\varepsilon_t^s}\left[\begin{array}{c} \phi^s h_{t-1}^w + v_t^s\int_{\varepsilon_t^s}^{\bar{\varepsilon}_t^s}\varepsilon dF_t^s\left(\varepsilon\right) + uv_t^sF_t^s\left(\varepsilon_t^s\right) \\ +\beta\left(h_{t+1}^w + u\left[v_t^s\phi^wF_t^s\left(\varepsilon_t^s\right) + 1 - \phi^w\right]\right) \\ +\beta^2W^s\left(h_{t+1}^w, v_{t+2}^s\right) \end{array}\right] \\ &\quad s.t. \\ h_{t+1}^w &= \phi^w\left[\phi^s h_{t-1}^w + v_t^s\int_{\varepsilon_t^s}^{\bar{\varepsilon}_t^s}\varepsilon dF_t^s\left(\varepsilon\right)\right], \\ v_{t+2}^s &= \phi^s\left[v_t^s\phi^wF_t^s\left(\varepsilon_t^s\right) + 1 - \phi^w\right] + 1 - \phi^s. \end{aligned} \tag{46}$$

Intuitively, "a period" for the decision of ε_t^s is equal to $2t$. The state variables for the current period are given by the vector $\left(h_{t-1}^w, v_t^s\right)$, the state variables for next period are $\left(h_{t+1}^w, v_{t+2}^s\right)$, and the control variable is ε_t^s.

The first order condition:

$$\begin{aligned} 0 &= v_t^s\left(-\varepsilon_t^sf_t^s\left(\varepsilon_t^s\right)\right) + uv_t^sf_t^s\left(\varepsilon_t^s\right) \\ &\quad +\beta\left(\phi^wv_t^s\left(-\varepsilon_t^sf_t^s\left(\varepsilon_t^s\right)\right) + uv_t^s\phi^wf_t^s\left(\varepsilon_t^s\right)\right) \\ &\quad +\beta^2\left[\frac{\partial W^s}{\partial h_{t+1}^w}\left(\phi^wv_t^s\left(-\varepsilon_t^sf_t^s\left(\varepsilon_t^s\right)\right)\right) + \frac{\partial W^s}{\partial v_{t+2}^s}\left(\phi^sv_t^s\phi^wf_t^s\left(\varepsilon_t^s\right)\right)\right], \end{aligned}$$

which simplifies to:

$$\begin{aligned} 0 &= -\varepsilon_t^s + u + \beta\left(-\phi^w \varepsilon_t^s + u\phi^w\right) \\ &\quad +\beta^2 \left[\frac{\partial W^s\left(h_{t+1}^w, v_{t+2}^s\right)}{\partial h_{t+1}^w}\left(-\phi^w \varepsilon_t^s\right) + \frac{\partial W^s\left(h_{t+1}^w, v_{t+2}^s\right)}{\partial v_{t+2}^s}\phi^s\phi^w\right] \end{aligned}$$

and can be written as:

$$\varepsilon_t^s\left[1 + \beta\phi^w + \beta^2\phi^w \frac{\partial W^s\left(h_{t+1}^w, v_{t+2}^s\right)}{\partial h_{t+1}^w}\right] = \left(1+\beta\phi^w\right)u + \beta^2\phi^w\phi^s \frac{\partial W^s\left(h_{t+1}^w, v_{t+2}^s\right)}{\partial v_{t+2}^s} \tag{47}$$

Using the envelope-theorem conditions, we obtain:

$$\frac{\partial W^s\left(h_{t-1}^w, v_t^s\right)}{\partial h_{t-1}^w} = \phi^s + \beta\phi^w\phi^s + \beta^2\phi^w\phi^s \frac{\partial W^s\left(h_{t+1}^w, v_{t+2}^s\right)}{\partial h_{t+1}^w}, \tag{48}$$

and

$$\begin{aligned} &\frac{\partial W^s\left(h_{t-1}^w, v_t^s\right)}{\partial v_t^s} \\ =&\ \left(1+\beta\phi^w\right)\left(\int_{\varepsilon_t^s}^{\bar{\varepsilon}_t^s} \varepsilon dF_t^s\left(\varepsilon\right) + v_t^s T_{2t}^s\right) + \left(1+\beta\phi^w\right)u\left[F_t^s\left(\varepsilon_t^s\right) - v_t^s T_{1t}^s\right] \\ &+\beta^2 \frac{\partial W^s\left(h_{t+1}^w, v_{t+2}^s\right)}{\partial h_{t+1}^w}\phi^w \left(\int_{\varepsilon_t^s}^{\bar{\varepsilon}_t^s} \varepsilon dF_t^s\left(\varepsilon\right) + v_t^s T_{2t}^s\right) \\ &+\beta^2 \frac{\partial W^s\left(h_{t+1}^w, v_{t+2}^s\right)}{\partial v_{t+2}^s}\phi^s\phi^w \left[F_t^s\left(\varepsilon_t^s\right) - v_t^s T_{1t}^s\right], \end{aligned}$$

where $T_{1t}^s \equiv \frac{\partial}{\partial v_t^s}\left[1 - F_t^s\left(\varepsilon^s\right)\right] > 0$ and $T_{2t}^s \equiv \frac{\partial}{\partial v_t^s}\int_{\varepsilon_t^s}^{\bar{\varepsilon}_t^s} \varepsilon dF_t^s\left(\varepsilon\right) > 0$. This last expression can hence be written as:

$$\begin{aligned} &\frac{\partial W^s\left(h_{t-1}^w, v_t^s\right)}{\partial v_t^s} \\ =&\ \left(1 + \beta\phi^w + \beta^2\phi^w \frac{\partial W^s\left(h_{t+1}^w, v_{t+2}^s\right)}{\partial h_{t+1}^w}\right)\left(\int_{\varepsilon_t^s}^{\bar{\varepsilon}_t^s} \varepsilon dF_t^s\left(\varepsilon\right) + v_t^s T_{2t}^s\right) \\ &+\left(\left(1+\beta\phi^w\right)u + \beta^2\phi^s\phi^w \frac{\partial W^s\left(h_{t+1}^w, v_{t+2}^s\right)}{\partial v_{t+2}^s}\right)\left[F_t^s\left(\varepsilon_t^s\right) - v_t^s T_{1t}^s\right] \end{aligned} \tag{49}$$

In steady state, (48) and (49) become

$$\frac{\partial W^s\left(h^w, v^s\right)}{\partial h^w} = \frac{\phi^s\left(1+\beta\phi^w\right)}{1 - \beta^2\phi^w\phi^s}, \tag{50}$$

and

$$\begin{aligned} &\frac{\partial W^s\left(h^w, v^s\right)}{\partial v^s}\left(1 - \beta^2\phi^s\phi^w\left[F^s\left(\varepsilon^s\right) - v^s T_1^s\right]\right) \\ =&\ \left(1 + \beta\phi^w + \beta^2\phi^w \frac{\phi^s\left(1+\beta\phi^w\right)}{1-\beta^2\phi^w\phi^s}\right)\left(\int_{\varepsilon^s}^{\bar{\varepsilon}^s} \varepsilon dF^s\left(\varepsilon\right) + v^s T_2^s\right) \\ &+\left(1+\beta\phi^w\right)u\left[F^s\left(\varepsilon^s\right) - v^s T_1^s\right]. \end{aligned} \tag{51}$$

Substituting into the FOC (47),

$$\varepsilon^s \frac{1+\beta\phi^w}{1-\beta^2\phi^w\phi^s} = (1+\beta\phi^w)\, u + \beta^2\phi^w\phi^s \frac{(1+\beta\phi^w)\, u\left[F^s(\varepsilon^s) - \upsilon^s T_1^s\right] + \frac{1+\beta\phi^w}{1-\beta^2\phi^w\phi^s}\left(\int_{\varepsilon^s}^{\bar{\varepsilon}^s} \varepsilon dF^s(\varepsilon) + \upsilon^s T_2^s\right)}{1-\beta^2\phi^s\phi^w\left[F^s(\varepsilon^s) - \upsilon^s T_1^s\right]}$$

which simplifies to (38).

8.5 The model with Transaction costs

We now introduce transaction costs for buying and selling a house into the baseline model. The value function of the homeowner is the same as (7) in the baseline model. The buyer's value function is modified to:

$$B^s = E_\varepsilon^s \max\left\{H^s(\varepsilon) - p^s - T_b^s(p^s), \beta B^w\right\},$$

so the cutoff ε^s is given by:

$$H^s(\varepsilon^s) - p^s - T_b^s(p^s) = \beta B^w,$$

and

$$\frac{\partial \varepsilon^s}{\partial p^s} = \frac{1-\beta^2\phi^w\phi^s}{1+\beta\phi^w}(1+\tau_b).$$

The seller's value function is modified to:

$$V^s = \beta V^w + u + \max_p \left[1 - F^s(\varepsilon^s(p))\right](p - T_v^s(p) - \beta V^w - u),$$

where the optimal price p^s solves

$$\frac{p^s - T_v^s(p^s) - \beta V^w - u}{(1-\tau_v)\, p^s} = \left(\frac{p^s f^s(\varepsilon^s)\frac{\partial \varepsilon^s}{\partial p^s}}{1-F^s(\varepsilon^s)}\right)^{-1}.$$

Following similar simplifications as in Appendix 8.1, we obtain

$$S_v^s = \left(\frac{1-\tau_v}{1+\tau_b}\right)\left(\frac{1-F^s(\varepsilon^s)}{f^s(\varepsilon^s)}\right)\frac{1+\beta\phi^w}{1-\beta^2\phi^w\phi^s},$$

$$p^s - T_v^s(p^s) = \frac{u}{1-\beta} + \left(1 + \frac{\beta^2\left[1-F^s(\varepsilon^s)\right]}{1-\beta^2}\right) S_v^s + \frac{\beta\left[1-F^w(\varepsilon^w)\right]}{(1-\beta^2)} S_v^w,$$

$$\frac{1+\beta\phi^w}{1-\beta^2\phi^w\phi^s}\varepsilon^s = S_v^s + u + T^s(p^s) + \frac{\beta\phi^w\left(1-\beta^2\phi^s\right)}{1-\beta^2\phi^w\phi^s}(V^w + B^w) - \frac{\beta^2\phi^w(1-\phi^s)}{1-\beta^2\phi^w\phi^s}(V^s + B^s),$$

and $B^s + V^s$ as in (18).

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To be Completed

9 Data Sources

For U.K. and U.S. data, see text.

Australia The housing price index comes from the Australia Bureau of Statistics (ABS); it is a weighted average for eight capital cities, available from 1986; the series is based on prices at settlement and are based on data provided to the land titles office; it is not quality adjusted. The CPI (non seasonally adjusted, NSA) also comes from the ABS and is a national index, not available at a disaggregated level; in what follows, for all countries, the price index considered in the analysis corresponds to the national index.

Belgium The housing price index comes from STADIM (*Studies & advies Immobiliën*) and covers Belgium and its three main regions from 1981; the series is based on the average selling prices of small and average single-family houses; apartments are not included; the data come from registered sales, and are not quality adjusted. The CPI (NSA) comes from the National Institute for Statistics.

Denmark The housing price index comes from the Association of Danish Mortgage Banks and corresponds to existing single-family homes (including flats and weekend cottages). The data come from the Land Registry, where all housing transactions are registered; they are not adjusted by quality and start in 1992. The CPI (NSA) comes from *Danmarks Statistik*.

France The housing price index comes from INSEE (National Institute for Statistics and Economic Studies) and corresponds to existing single-family homes. The data are not quality adjusted and start in 1994. The index covers all regions, and comes also disaggregated into 4 regions. The CPI (NSA) comes from the same source.

Ireland The housing price index comes from *Permanent TSB*, which accounts for about 20 percent of residential mortgage loans in the country, starting in 1996; the index is adjusted by the size of the property, dwelling type (detached, semi-detached, terrace, or apartment), and heating system. The number of transactions (loans) comes from the same source. The CPI (NSA) comes from the Central Statistical Office in Ireland.

Netherlands The housing price index comes from the Dutch Land Registry; it is a repeat-sale index, starting in 1993. The CPI (NSA) comes from the CBS (Statistics Netherlands).

New Zealand The housing price index comes from the Reserve Bank of New Zealand, starts in 1968, and is not adjusted by quality; the CPI (NSA) comes from the same source.

Norway The housing price index comes from Statistics Norway, starting in 1992; the data are not adjusted by quality as meticulously as in the U.K., however, the properties considered need to satisfy a set of broadly defined) characteristics to be included in the index; the CPI (NSA) comes from the same source.

South Africa The housing price index comes from ABSA, a commercial bank that covers around 53 percent of the mortgage market in South Africa. The data are recorded at the application stage of the mortgage lending process and the series starts in 1975. There is no quality adjustment, although the properties considered need to satisfy a set of (broadly defined) characteristics to be included in the index. The CPI (NSA) comes from Statistics South Africa.

Sweden The housing price index comes from *Statistika Centralbyrån*; the data correspond to one and two-dwelling properties and are not quality-adjusted; the series starts in 1986; data on transactions and CPI (NSA) come from the same source.