Received 25 April 2021
Revised 30 May 2021
Accepted 1 June 2021

# The effects of tourism on housing prices: applying a difference-in-differences methodology to the Portuguese market

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## Abstract

**Purpose** – This paper aims to explore the effects of a surge in tourism short-term rentals (STR) on housing prices in municipalities within Portugal's two largest Metropolitan Statistical Areas.

**Design/methodology/approach** – This study applies the difference-in-differences (DiD) methodology by using a feasible generalized least squares (FGLS) estimator in a seemingly unrelated regression (SUR) equation model.

**Findings** – The results show that the liberalization of STR had a significant impact on housing prices in municipalities where a higher percentage of housing was transferred to tourism. This transfer led to a leftward shift in the housing supply and a consequent increase in housing prices. These price increases are much higher than those found in previous studies on the same subject. The authors also found that municipalities with more STR had low housing elasticities, which indicates that adjustments to the transfer of real estate from housing to tourism were made by increasing house prices, and not by increasing supply quantities.

**Practical implications** – The study suggests that an unforeseen consequence of allowing property owners to transfer the use of real estate from housing to other services (namely, tourism) was extreme housing price increases due to inelastic housing supply.

**Originality/value** – This is the first time that the DiD methodology has been applied in real estate markets using FGLS in a SUR equation model and the authors show that it produces more precise estimates than the baseline OLS FE. The authors also find evidence of a supply shock provoked by STR.

**Keywords** Elasticities, Housing, Tourism, Difference-in-differences, Real estate prices, Short-term rentals

**Paper type** Research paper

## 1. Introduction

The importance of real estate prices extends beyond the role of financial supervisors, as it directly affects the living conditions of the population. Real estate is an asset class with several uses: either housing or services (offices, shops, tourism facilities, etc.). Although there is consensus that aggregate income is a real estate price determinant (Quigley, 1999) and that tourism positively affects economic growth (Balaguer and Cantavella-Jordá, 2002), the debate about the direct impact of tourism on real estate prices remains open. While some

International Journal of Housing Markets and Analysis
Vol. 15 No. 4, 2022
pp. 762-779
Emerald Publishing Limited
1753-8270
DOI 10.1108/IJHMA-04-2021-0047

authors find that cities specializing in the tourism industry often suffer from Dutch disease, leading to an increase in real estate prices (Sheng, 2011), other authors have shown that, in small coastal cities, tourism results in a decrease in real estate prices (Biagi, 2016). There are also new factors that affect the relationship between both variables.

Starting in the late 2000s, what is known as the “sharing economy” meant that real estate with housing permits began to be used for tourism, mainly through owner-occupiers swapping houses for short periods. This new industry created new problems, in particular, housing price increases (Jefferson-Jones, 2015) and a decrease in the supply of long-term rental units (Barron *et al.*, 2021). A study of the short-term rental activity in Berlin (Schäfer and Braun, 2016) found that rental growth was higher in markets where the short-term rental share was higher. In Portland, a study about relaxing building permit requirements (Brotman, 2021) in 2014 found that the increased listing of properties on short-term rental websites was not dependent on increased economic activity. In 2019, the city’s rental laws were tightened to increase the availability of rental housing available to residents at affordable prices. In several European cities such as Venice (Coldwell, 2017), Barcelona (Diaz, 2017) and Amsterdam (Stone, 2018), local governments began to limit the short-term rental of apartments for tourism. It is important to research the significance of the problems arising from short-term rentals (STR) and housing prices and the causality between the two variables. This will help politicians, the resident population and market players understand the consequences of liberalizing the licensing of short-term rental establishments and determine whether the remedies recently implemented are effective in addressing those problems.

For instance, in Portugal, the European Union country with the highest percentage increase in tourist arrivals (UNTWO, 2019) between 2010 and 2018 (a 234% increase), it is likely that house prices may be affected not only by economic and demographic factors but also by each city’s tourism activity and the transfer of apartments and houses to tourism activities. A Portuguese law of 2012 allowed housing to be transferred to tourism activities without requiring a change in the housing use permit of the real estate asset. Nevertheless, the licensing of STR remained in the hands of the municipalities, which allowed some discretionary political decisions. However, a law of August 2014 (Decree-Law 128/2014) fully liberalized the sector making the licensing automatic. In the first half of 2019, 34% of the houses sold in Lisbon, Portugal’s capital city, were purchased by foreigners from 70 different countries (Schiffmann, 2019). As there was no equivalent immigration of people during the period and the registration of apartments and houses in local establishments for tourism purposes (STR) was increasing exponentially, the logical conclusion was that foreigners were buying residential apartments and converting them into tourism facilities. Between 2018 and 2019, the two largest Portuguese municipalities (Lisbon and Porto) reversed the liberalization and implemented containment measures by suspending new tourism STR in certain zones of the cities. In 2020, the central government decided to suspend the issuance of Golden Visas for foreigners purchasing real estate in Lisbon and Porto Metropolitan Statistical Areas (MSA) (Tiago and Fernandes, 2020) to prevent real estate prices in these two MSA from rising even further.

Figure 1 shows that, since 2015, house prices in Portugal have increased at a faster pace than the European Union average.

In this paper, we study the effects of political intervention on real estate prices in the two largest Portuguese MSA: Lisbon (2.8 million resident population and 18 municipalities) and Porto (1.7 million resident population and 17 municipalities). Both MSA belongs to the upper decile of the European Union MSA (also known as NUTS 3) in terms of population and both are classified by the Globalization and World Cities Research Network as world cities

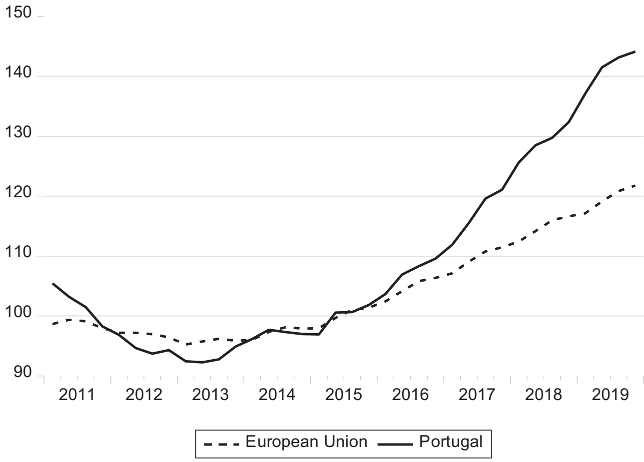



**Figure 1.**
House price indices

**Source:** Eurostat. Quarterly data. 2015 = 100

attracting business and foreign residents from around the world. Lisbon is classified as an Alpha-world city and Porto as a Gamma+ world city.

We apply a difference-in-differences (DID) methodology to study the impact of the 2014 liberalization of the registration of STR on housing prices in the municipalities of Lisbon and Porto MSA most impacted by the liberalization. We followed the recent econometric literature, which suggests the advantages of the feasible generalized least squares (FGLS) estimator over the ordinary least squares (OLS) estimator. To the best of our knowledge, there has been little or no application of FGLS in DiD studies about house prices. Moreover, most of the published research about the impact of tourism on house prices is based on the demand side of the price equilibrium equation, with studies of Airbnb density as an instrumental variable of tourism demand (Barron *et al.*, 2021; Franco and Santos, 2021). We wanted to explore the housing supply side, namely, the effects of short-term rental on the housing supply and also the housing supply elasticities.

We found that, by allowing real estate assets with residential use permits to be freely transferred to the tourism market, the liberalization effectively impacted housing prices in the municipalities most exposed to STR. Following the liberalization, for each one percentage point increase in the share of STR as a percentage of the housing stock, housing prices increased 27.4% and 16.1% in the Lisbon and Porto MSA municipalities most exposed to STR, respectively. These results represent a much higher impact than that estimated in previous studies (Franco and Santos, 2021). We estimated the housing supply elasticities of the groups of municipalities examined for each MSA and found that the groups of municipalities which were more exposed to STR in both MSA indicate much lower elasticities than the other groups, with the Lisbon MSA exposed group showing an inelastic housing supply. Moreover, we show that, in the exposed groups, STR provoked a supply shock, with the housing supply curve shifting leftward. We conclude that allowing for the transfer of real estate use from housing to tourism (or other services) caused a supply shock, with the shock absorbed mostly by price adjustments rather than by quantities.

The remainder of the paper is organized as follows. Section 2 reviews the literature on the relationship between real estate prices and tourism. Section 3 describes the data collected and the methodology for the model. Section 4 presents the results of the regression models applied to the sets of data, Section 5 shows the robustness test results and Section 6 presents the conclusions.

## 2. Literature review

The basic approach to real estate prices is based on David Ricardo's (1817) Rent Theory, a model where spatial differentiation (location of economic activities) and productivity are the most important factors affecting real estate prices. He further explains that the land (real estate) rents (prices) are a consequence of the productivity (income) of the economic activity carried out in the location and not the other way round. From an asset pricing perspective, this means that the price of a property equals the present value of its future rents. Plazzi *et al.* (2010) showed that a one percentage point increase in the property yield (rent on price) leads to an increase of up to 4% in property prices, but there are fundamental differences depending on the economic use of the property (housing, commercial or other). Recently, international house prices have become synchronized within an elite set of major cities (Duca, 2020), often in warmer or coastal areas, where the supply elasticity of housing is often limited. These markets have become more sensitive to shifts in the international demand for property by real estate investors, with the effects being amplified by the tendency for property owners to use extrapolative expectations of future house prices.

Regarding the impact of tourism on house prices, Schäfer and Hirsch (2017) concluded that Berlin rents are affected by urban tourism. Barron *et al.* (2021) also reached the same conclusions in a study about the impact of the Airbnb platform on USA neighborhood house prices and rents. They further found that, although the total supply of housing is not affected by the entry of Airbnb, its listings increase the supply of short-term rental units and decrease the supply of long-term rental units.

In terms of methodology, the hedonic pricing method, for which Rosen (1974) provides the theoretical foundation, is widely used in housing research and appraisals. Hedonic pricing method research into the relationship between tourism and property prices has focused on tourist accommodation such as cottages (Fleischer and Tchetchik, 2005) or holiday homes (Conroy and Milosch, 2011). As the hedonic pricing method is not suitable for estimating long-run equilibrium, there is a different approach for identifying the determinants of real estate prices by modeling equilibrium prices using equations representing inverted demand or supply of real estate. These models have been applied extensively, although, given the difficulty of finding data on the supply side of the market (such as, for instance, planning regulations and land use) and given the slow response of the housing supply and prices in producing any changes in the market, most of the applied research focuses on the demand side (Mankiw and Weil, 1989). Glaeser *et al.* (2014) developed a dynamic linear rational equilibrium model to explain house price variations between MSA. They found that most variations in housing price changes are local, not national and that price changes are predictable, except for the fact that each market has a specific elastic supply of homes that generate different housing market dynamics. Moreover, there is also the question of what happens to neighboring cities. In a study of 363 US MSA, Cohen *et al.* (2016) found that there are significant spatial diffusion patterns in the growth rates of real estate prices. This underscores the existence of spatial spillover effects between cities. Recent studies have started to address the impact of tourism on urban growth (Ma *et al.*, 2015). They found that tourism growth contributes to positive spatial correlations that lead to spillovers between neighboring municipalities within the same MSA. The debate on the importance of tourism for real estate prices is still ongoing. Biagi (2016) found that increased tourism activity leads to increased housing prices in certain Italian cities, but in other cities, increased tourism activity decreases housing prices. The negative effect of tourism on housing prices is found primarily in small cities where marine tourism predominates.

Finally, there is a flow of new research studying the impact on real estate of specific events or special information conditions using the DiD methodology. Dubé *et al.* (2017) developed a spatial-temporal DiD estimator to measure the effect of urban externalities (such as transport infrastructures) on real estate prices. Badarinza and Ramadorai (2018) found that event-triggered foreign capital inflows have an impact on London real estate prices. Somerville *et al.* (2020) studied the effects of imposing restrictions on non-owner-occupant purchases of real estate and concluded that these restrictions reduce activity levels by approximately 40%, but that the policies do not lead to any subsequent relative price changes. Franco and Santos (2021) quantified the impact of Airbnb STR on housing affordability in Portugal and found that, on average, a one percentage point increase in a municipality's Airbnb share results in a 3.7% increase in house prices.

This growing body of research using the DiD methodology is based on the OLS estimator, with fixed effects (FE). Although this method makes it possible to circumvent endogeneity problems, Bertrand *et al.* (2004) argue that these estimations severely understate the standard deviations of the estimator due to a possibly severe serial correlation problem. Hansen (2007) suggested using FGLS to correct for serial correlation, arguing that it will result in a more efficient estimator and more powerful tests than OLS. Hausman and Kuersteiner (2008) demonstrated that, in the presence of serial correlation, both the DiD methodology and the FE do not use all the time-series variation in the data. They also found that the FGLS estimator has more power than the OLS, with FE in the presence of serial correlation when the period sample is long. Brewer *et al.* (2017) also showed that in a DiD, as periods increase, the precision of the OLS estimator falls due to serial correlation. They recommend using FGLS with cluster-robust techniques in DiD applications.

In sum, the OLS estimator with FE that is widely used in DiD has some limitations in terms of precision. We will innovate by applying the FGLS to a seemingly unrelated regression (SUR) equation model (Zellner, 1962) in a DiD study of a sample of municipalities in Portugal's two largest MSA. This continues the work of Franco and Santos (2021), but with different data, different variables and a different estimator. We will also innovate by estimating the elasticities of the sample group municipalities to compare them to the DiD results.

## 3. Data and methodology

Measuring the impact on house prices of an increase in STR as a percentage of the housing stock could lead to an endogeneity problem if other variables are also affecting housing prices. Following the work of Ashenfelter and Card (1985), it is possible to control for those omitted variables by applying a DiD methodology. This approach investigates whether an intervention influences an outcome over time by comparing observed differences in a case sample that receives the intervention with observed differences in a control sample that does not. In other words, whether there is a difference in an outcome due to a hypothesized treatment over a difference in time. This makes it possible to isolate the treatment effect above and beyond any difference that would have been expected regardless of the treatment. Because it relies on direct before-and-after measures at the same site rather than relying on modeling differences between sites, the DiD approach provides a more direct measure of the added value attributable to short-term rental licensing than that provided by more frequently used hedonic modeling techniques. However, to apply the methodology, we need to have the date (or several dates) of the intervention to specify the event study. Fortunately, we have such an event in Portugal's two largest MSA.

*3.1 Data*

In Portugal, short-term rental is a relatively recent phenomenon. Before 2012, the activity was subject to tourism licensing and the real estate asset would have needed a use permit for services. After 2012, it became possible to use real estate with housing use permits for STR, but under municipal licensing. In August 2014, the activity was fully liberalized, with automatic licensing in the registry under Decree-Law N° 128/2014 (legal scheme for the exploitation of local accommodation establishments). From almost zero registries in 2011, licensed STR had increased to more than 90,000 nationwide by December 2019. It became possible to transfer houses and apartments from housing to tourism without any political or administrative opposition. This allows us to analyze the evolution of short-term rental licensing in the different municipalities without the bias of an omitted variable (efficiency of city's municipal services or lack of political will to authorize the licensing).

The dependent variable (Price) is the monthly median price per square meter of bank valuation of housing in the municipalities within both MSA. This was obtained from the Statistics Portugal website database – which is affiliated to Eurostat – for all municipalities of the Lisbon MSA and the Porto MSA from January 2011 to December 2019. Using the price per square meter allows us to avoid controlling for hedonic characteristics such as the number of rooms and size. Unfortunately, these two positive characteristics of the data set (median and price per square meter) limit the number of municipalities with complete information. In the Lisbon MSA, we have data on 17 out of 18 municipalities and in the Porto MSA, we have data on 12 out of 17 municipalities. We have 108 price observations for each of the 29 municipalities, which gives a total of 3,132 observations.

Regarding the independent variables, we used data from the short-term rental national registry – a special registry within the Tourism Department Registry – to create a new variable STR representing the ratio of the accumulated number of licensed apartment and house STR as a percentage of the housing stock of each municipality (data from the Statistics Portugal website database). This allows us to control for the increase in the housing supply (which could affect the results of the empirical study), as we will compare the municipalities according to the percentage of housing transferred to STR at the end of the period from January 2011 to December 2019. Registering an apartment or house in the national registry has a tax regime that discourages owners from renting those apartments or houses in the long-term residential rental market (to do so, they have to pay capital gains tax). With this limitation, we know that most registered STR was excluded from the long-term residential rental market and are exclusively in the short-term rental market. The ranking of municipalities is reported in Table 1.

With this new variable STR, we rank each municipality into two groups and according to quantiles (quartile in this section, median in the robustness tests section). This creates a new independent variable T (Treated), of which the upper quartile group is the treated (higher share of STR in the housing stock) and assumes the value of one (dummy variable). The lower quartile group is the control (lower share of STR in the housing stock) and assumes the value of zero. For the Lisbon MSA, we include in the upper quartile (numbers 1, 2, 3 and 4) the municipalities of Lisbon, Sesimbra, Mafra and Cascais and in the lower quartile (numbers 14, 15, 16 and 17), the municipalities of Odivelas, Moita, Barreiro and Vila F. X. For the Porto MSA, we include in the upper quartile (numbers 1, 2 and 3) the municipalities of Porto, Póvoa and Vila N.G. and in the lower quartile (numbers 10, 11 and 12), the municipalities of Santo T., Valongo and Oliveira A. In the robustness test, we run the estimation for the municipalities in the above-median and below-median groups.

There is one important observation to be made from the rank identification of the municipalities in the upper and lower quartiles: those within the same group are geographically non-contiguous. Lisbon, Sesimbra, Mafra and Cascais (upper quartile of Lisbon MSA) are all separated by other municipalities. The same happens with Porto, Póvoa and Vila N.G., which either have other municipalities in between or a large river separating them. This is also the case of the municipalities in the lower quartile of the MSA: they are not contiguous and are far apart. This suggests that there could be price spatial spillovers from the treated group municipalities to the control group municipalities (which are spatially close). According to Cox (1958), to make an inference we need the stable unit treatment value assumption to hold. This means that the house prices of a municipality in the control group should not be affected by the house prices of contiguous municipalities in the treated group. Violating this assumption would lead to cross-section dependence. Nevertheless, in our study, a violation of this assumption would generate a smaller post-treatment effect and so we consider our estimates to be valid, as the estimated effect will be smaller than the real effect.

We have an independent dummy variable (L, Liberalization) representing the pre- and post-liberalization of the activity, with the value of zero from January 2011 to August 2014 (the month when the registration was fully liberalized according to Law-Decree 128/2014) and the value of one from September 2014 to December 2019, applied to all municipalities.

We included two independent time-varying covariates to control for supply and demand confounders affecting housing prices in each municipality: density (D) and salaries (S). Density (number of residents per square kilometer) is a measure of the price elasticity of housing supply (Green *et al.*, 2005) or supply constraints. Salaries (gross mean salary in euros per month) is a measure of the income elasticity of house demand (Green and Hendershott, 1996) or demand shocks. Data was collected from the Statistics Portugal website database.

### *3.2 Descriptive statistics*

Table 2 shows univariate statistics of the samples of the two groups (treated and control) for each MSA.

Within the same MSA, the mean of the variables in the treated group is higher than in the control group, except for Density in the Lisbon MSA. There are municipalities in the control groups with zero STR, as we expected. Between MSA, Lisbon shows higher means than Porto, except for Density in the treated group.

**Table 1.** Accumulated share of tourism short-term rentals as a percentage of the housing stock, in December 2019, by the municipality and metropolitan statistical area

| | | | |
|---|---|---|---|
| *Panel A – Lisbon MSA municipalities* | | | |
| 1 Lisbon (5.787%) | 2 Sesimbra (2.349%) | 3 Mafra (2.234%) | 4 Cascais (1.589%) |
| 5 Almada (0.881%) | 6 Setúbal (0.792%) | 7 Sintra (0.530%) | 8 Palmela (0.411%) |
| 9 Oeiras (0.316%) | 10 Seixal (0.226%) | 11 Montijo (0.174%) | 12 Loures (0.106%) |
| 13 Amadora (0.079%) | 14 Odivelas (0.075%) | 15 Moita (0.066%) | 16 Barreiro (0.045%) |
| 17 Vila F. X. (0.044%) | | | |
| *Panel B – Porto MSA municipalities* | | | |
| 1 Porto (5.475%) | 2 Póvoa V. (0.698%) | 3 Vila N. G. (0.671%) | 4 Vila Conde (0.528%) |
| 5 Matosinhos (0.383%) | 6 Maia (0.121%) | 7 Gondomar (0.104%) | 8 Paredes (0.076%) |
| 9 Santa M. F. (0.071%) | 10 Santo T. (0.063%) | 11 Valongo (0.054%) | 12 Oliveira A. (0.036%) |

### 3.3 Housing supply elasticities

Following the findings of Green *et al.* (2005), we were interested in finding out the housing supply elasticities of both groups and comparing them to the DiD results. We have the housing stock annual data from 2011 to 2019 and the median prices per square meter of both samples for the treated and control groups. We formulated the housing supply function in equation (1) and proceeded with the logarithmic transformation of the variables to account for the non-linearity of distributions. The log-log specification allowed us to interpret the coefficient as housing supply elasticity.

$$\mathrm{Log}\left(\mathrm{Q}_{gt}\right) = \theta_0 + \theta_1 \cdot \mathrm{Log}\left(\mathrm{P}_{gt}\right) \tag{1}$$

where $\mathrm{Q}_{gt}$ is the aggregate quantity of housing stock of the group $g$ in year $t$, $\mathrm{P}_{gt}$ is the aggregate median price per square meter of group $g$ in year $t$, $\theta_1$ is the coefficient to estimate.

We also investigated how the registration of tourism STR impacts the housing supply of each group, by subtracting the annual number of registered STR from the annual housing stock and calculating the net growth of housing supply. As mentioned in Section 3.1, the change from STR back to residential housing had several tax disadvantages that discouraged property owners from transferring the use of the real estate (Warren and Almeida, 2020).

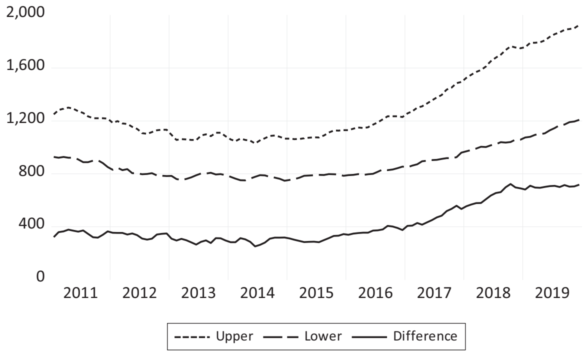
### 3.4 Parallel trends assumption

For the delta between the trends to be a valid estimate of the causal effect, the DiD methodology relies on the assumption that prices of the two groups (treated and control) followed a parallel trend before the treatment date. If not, the effect of time on price would be different for the treated group, even if it would not get the treatment, and thus we would have an endogeneity problem because the error term would correlate with the independent variables. Verifying the parallel trend assumption is done by graphic visualization. We created a new series, named Difference, composed of the difference between the mean price of the two groups (upper quartile and lower quartile). We expected that series to remain flat until August 2014 and then start to drift upwards. Figure 2 shows the graphic representation of the Lisbon MSA series and Figure 3 shows the Porto MSA series.

Both graphs indicate that the series "difference" shows no trend until just after the middle of 2014 and that it starts to drift upwards after the liberalization (the treatment date).

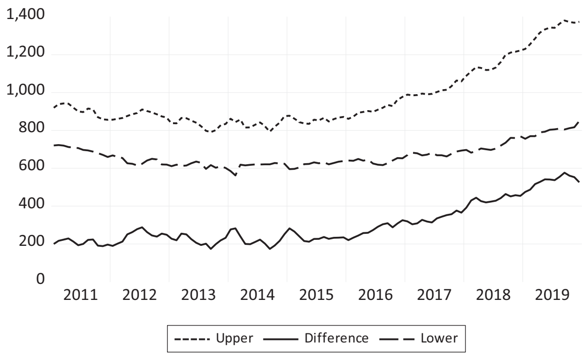
| | Treated | | | | | Control | | | | |
|---|---|---|---|---|---|---|---|---|---|---|
| Variables | Obs. | Mean | S.D. | Min. | Max. | Obs. | Mean | S.D. | Min. | Max. |
| *Panel A – Lisbon MSA* | | | | | | | | | | |
| *Price* | 432 | 1,299 | 526 | 718 | 3,021 | 432 | 880 | 214 | 563 | 1,720 |
| *STR* | 432 | 1.008 | 1.338 | 0.011 | 5.786 | 432 | 0.022 | 0.035 | 0 | 0.174 |
| *Density* | 432 | 2,010 | 2,086 | 257 | 6,324 | 432 | 2,386 | 2,070 | 424 | 6,095 |
| *Salaries* | 432 | 1,145 | 271 | 889 | 1,616 | 432 | 1,027 | 101 | 894 | 1,216 |
| *Panel B – Porto MSA* | | | | | | | | | | |
| *Price* | 324 | 968 | 266 | 670 | 1,936 | 324 | 668 | 76 | 493 | 970 |
| *STR* | 324 | 0.704 | 1.340 | 0.003 | 5.475 | 324 | 0.016 | 0.019 | 0 | 0.063 |
| *Density* | 324 | 2,618 | 1,982 | 758 | 5,627 | 324 | 733 | 391 | 410 | 1,297 |
| *Salaries* | 324 | 1,127 | 151 | 946 | 1,375 | 324 | 958 | 71 | 829 | 1,097 |

**Table 2.** Lisbon MSA and Porto MSA treated and control groups' descriptive statistics




**Note:** "Upper" is the mean price of the upper quartile municipalities, "lower" is the mean price of the lower quartile municipalities, "difference" is the difference between upper and lower for each monthly observation

**Figure 2.**
Lisbon MSA parallel trend assumption




**Note:** "Upper" is the mean price of the upper quartile municipalities, "lower" is the mean price of the lower quartile municipalities, "difference" is the difference between upper and lower for each monthly observation

**Figure 3.**
Porto MSA parallel trend assumption

### *3.5 Estimation method*

We started by modeling the impact of STR on price. In this equation (2), we included the treated and control groups in the same sample (we did not run the DiD in this specification), as a proxy for measuring the impact on housing prices of the transfer of housing to a short-term rental. We proceeded with the logarithmic transformation of the variables to account for the non-linearity of the variables' distributions. Before the logarithmic transformation, we added the number one to all observations of variable STR, as this variable has several zero value observations. We use White robust standard errors clustered in the cross-sections:

$$\mathrm{Log}(\mathrm{P}_{ijt}) = \beta_{0j} + \beta_1 \cdot \mathrm{Log}(\mathrm{STR}_{ijt}), \tag{2}$$

where Log($P_{ijt}$) is the log-price of residential real estate in municipality $i$ from MSA $j$ in month $t$, Log($STR_{ijt}$) is the log-share of tourism STR as a percentage of the housing stock of municipality $i$ from MSA $j$ in year $t$ and $\beta$ terms are the coefficients to be estimated by the model. We tested for cointegration (Johansen Fisher Panel test t-statistics 43.14 and 27.98 for the Lisbon and Porto MSA, respectively, $p$-values smaller than 1%). We did not include the covariates because they are not cointegrated. We tested for cross-section and period effects (Redundant FE t-statistics 142.862 and 47.816 for Lisbon and Porto MSA, respectively, p-values smaller than 1%). We used the OLS estimator with FE in the cross-section and period.

We formulated equation (3) to specify the DiD model. We were particularly interested in the sign and significance of the interaction term T.L. We expected that housing prices in municipalities in the control group would not be affected by the time change from pre- to post-liberalization above the expected price growth trend. We expected that municipalities in the treated group would exhibit a significant change in housing prices from pre- to post-liberalization. Thus, a significant and positive T.L coefficient would indicate that the liberalization of STR increased housing prices. The specification for estimating the interaction is the following:

$$\mathrm{Log}(\mathrm{P}_{ijt}) = \beta_{0j} + \beta_1 \cdot \mathrm{T}_{ij} \cdot \mathrm{L}_t \tag{3}$$

where Log($P_{ijt}$) is the log-price of residential real estate in municipality $i$ from MSA $j$ in month $t$, $T_{ij}.L_t$ is a dummy variable set to one if the group of municipalities $i$ is in the upper quartile of the MSA $j$ and the month $t$ is post-liberalization, or zero otherwise; $\beta$ terms are the coefficients to be estimated by the model. The estimator is the OLS FE.

Following the recommendation of Hansen (2007) for the DiD methodology, we also used an FGLS estimator, applied to a SUR equation model (Zellner, 1962), thus correcting for heteroscedasticity and serial and contemporaneous correlation in the residuals (Parks, 1967), clustering standard errors by the municipality. We applied this estimator to equation (3).

Then finally, we included the logged covariates in the DiD model for control, formulating equation (4). We tested the relevance of the inclusion of the control variables (density and salaries), through for the omitted variable test (likelihood ratio) and we rejected the hypothesis that these two variables do not make a relevant contribution (density: $t$-statistics 5.715 and 3.951, for the Lisbon and Porto MSA, respectively, with $p$-values smaller than 1%; salaries: $t$-statistics 4.946 and 5.134, for the Lisbon and Porto MSA, respectively, with $p$-values smaller than 1%).

$$\mathrm{Log}(\mathrm{P}_{ijt}) = \beta_{0j} + \beta_1 \cdot \mathrm{T}_{ij} \cdot \mathrm{L}_t + \beta_2 \cdot \mathrm{Log}(\mathrm{D}_{ijt}) + \beta_3 \cdot \mathrm{Log}(\mathrm{S}_{ijt}), \tag{4}$$

where Log($D_{ijt}$) is the number of resident persons per square kilometer of the municipalities $i$ from MSA $j$ in month $t$ (annual data repeated monthly); Log($S_{ijt}$) is the mean salary of municipality $i$ form MSA $j$ in month $t$ (annual data repeated monthly); the remaining variables are the same as the previous specifications. The estimator is the OLS FE.

Again, we also estimate the specification with logged covariates using FGLS SUR instead of OLS FE, applying this estimator to equation (4).

## 4. Empirical results

Table 3 shows the results of the estimates of housing supply elasticities of the treated and control groups.

Within the same MSA, there are differences in housing supply elasticities between the group of municipalities in the upper quartile of Short-Term Rental as a percentage of housing stock (treated group) and the group of municipalities in the lower quartile (control group). The Lisbon MSA treated group shows a 0.678 elasticity. Elasticities lower than 1 are evidence of an inelastic supply. Although the Porto MSA treated group shows a housing supply elasticity greater than 1 (1.885), it is much smaller than the control group (2.367). These numbers may be related to the fact that treated group municipalities are geographically constrained due to water bodies (sea and river) and high-density levels, as explained by Paciorek (2013).

In Table 4 we report how the short-term rental growth impacted the housing supply of both groups.

The results show that, in both panels, an increase in STR in the treated groups was accompanied by a decrease in the housing stock available for residential purposes. We conclude that, in the treated groups, the construction of new housing was not enough to offset the transfer of housing to STR. The effect of this phenomenon was a negative supply shock, with the supply curve shifting leftwards.

In Table 5 we report the estimated coefficients $\beta$ for the equations (2) to (4).

Equation (2) shows that a 1% increase in short-term rental leads to a 14% and 21.3% average increase in the housing prices of Panel A and Panel B, respectively. This serves as a proxy for the mean impact on housing prices of an aggregate decrease in housing supply in the sample municipalities of the respective MSA, although not controlling for other factors, which would render the coefficient smaller. As empirical evidence of this impact, from January 2011 to December 2019, the accumulated transfer of housing to STR in the Lisbon and Porto municipalities was 5.787% and 5.475%, respectively (Table 1). In the same period, the increase in housing prices was 76.9% and 79% in Lisbon and Porto, respectively.

**Table 3.** Housing supply elasticities of the treated and control groups

| | Panel A – Lisbon MSA | | Panel B – Porto MSA | |
|---|---|---|---|---|
| Group of municipalities | Treated group | Control group | Treated group | Control group |
| Elasticity | 0.678*** | 1.465** | 1.885*** | 2.367*** |

**Notes:** The table reports the estimated $\theta_1$ coefficient of equation (1). The dependent variable is the log-quantity of housing, the independent variable is log-price. The estimated coefficient is the housing supply price elasticity. The estimator is the OLS with Hubert-White heteroscedasticity-consistent standard errors and covariance. The estimated coefficients are multiplied by 100, for easier interpretation. ***, **, indicate significance at the 1% and 5% levels, respectively

**Table 4.** Housing supply of the treated and control groups

| | Panel A – Lisbon MSA | | Panel B – Porto MSA | |
|---|---|---|---|---|
| Group of municipalities | Treated group (%) | Control group (%) | Treated group (%) | Control group (%) |
| Gross growth | 0.700 | 1.426 | 1.316 | 1.253 |
| Net growth | −3.641 | 1.370 | −1.448 | 1.203 |

**Notes:** Gross growth is the percentual accumulated growth of the housing stock of each group, from 2011 to 2019. Net growth is the percentual accumulated growth of the housing stock net of the accumulated growth of short-term rentals of each group, from 2011 to 2019

**Table 5.** Upper and lower quartiles sample regressions results

| Specifications | Equation (2) | Equation (3) | Equation (3) | Equation (4) | Equation (4) |
|---|---|---|---|---|---|
| Estimator | OLS FE | OLS FE | FGLS SUR | OLS FE | FGLS SUR |
| *Panel A – Lisbon MSA* | | | | | |
| Intercept | 6.900*** (0.002) | 6.909*** (0.002) | 6.826*** (0.010) | 14.173*** (0.955) | 0.063 (0.081) |
| Short-term rental | 0.140*** (0.005) | – | – | – | – |
| Treated × liberalization | – | 0.065*** (0.008) | 0.272*** (0.010) | 0.056*** (0.009) | 0.274*** (0.011) |
| Density | – | – | – | −0.343*** (0.047) | 0.096*** (0.002) |
| Salaries | – | – | – | −0.691*** (0.141) | 0.872*** (0.011) |
| $R^2$ adjusted | 0.971 | 0.960 | 0.252 | 0.962 | 0.917 |
| F statistic | 236.934*** | 181.053*** | 291.310*** | 189.867*** | 3,196.124*** |
| Observations | 864 | 864 | 864 | 864 | 864 |
| *Panel B – Porto MSA* | | | | | |
| Intercept | 6.635*** (0.001) | 6.647*** (0.004) | 6.546*** (0.008) | 23.236*** (2.075) | 2.490*** (0.162) |
| Short-term rental | 0.213** (0.001) | – | – | – | – |
| Treated × liberalization | – | 0.098*** (0.012) | 0.268*** (0.009) | 0.0355*** (0.013) | 0.161*** (0.006) |
| Density | – | – | – | −1.157*** (0.171) | 0.113*** (0.005) |
| Salaries | – | – | – | −1.216*** (0.240) | 0.482*** (0.025) |
| $R^2$ adjusted | 0.969 | 0.936 | 0.436 | 0.944 | 0.826 |
| *F* statistic | 176.746*** | 82.396*** | 490.814*** | 485.537*** | 1,000.363*** |
| Observations | 635 | 635 | 635 | 635 | 635 |

**Notes:** The table reports the estimated coefficients from equations (2), (3) and (4), with the observations from the sample upper and lower quartiles. The dependent variable is the natural logarithm of housing prices. The independent variable short-term rental is the natural logarithm of the share of tourism short-term rentals as a percentage of housing stock with the value of one added. The interaction between tourism and liberalization is a dummy variable with the value of one if the municipalities are in the upper quartile of the variable STR at the end of 2019 and if the month of observation is after August 2014 and otherwise zero. Density is the natural logarithm of the municipality density and salaries is the natural logarithm of the municipality mean gross monthly salary. OLS FE is the ordinary least square estimator with fixed effects both at the cross-section and period. FGLS SUR is the feasible least squares estimator of the seemingly unrelated regressions. Robust standard errors clustered at the municipality level are reported in parentheses. ***indicates significance at the 1% level

Equation (3) shows that the liberalization had a positive and significant impact on housing prices in the treated group (municipalities in the upper quartile of short-term rental), meaning that the control group did not have an equivalent surge in housing prices following the liberalization. The impact with OLS FE is 6.5% and 9.8% in Panel A and Panel B housing prices, respectively, for each 1% increase in the share of STR as a percentage of the housing stock.

In equation (3) we also measure the same impact, but with the FGLS SUR estimator instead of the OLS FE, as suggested by Hansen (2007). We find much larger coefficients (27.2% for Panel A and 26.8% for Panel B) for similar size standard errors. We also find a much smaller adjusted R-squared (0.252 for Panel A and 0.436 for Panel B) than in the OLS FE estimate (0.960 for Panel A and 0.936 for Panel B). The F statistic is higher in equation (3) with FGLS SUR than with OLS FE, suggesting that the FGLS estimator produces a better fit than the intercept-only model when compared to the OLS FE estimator.

The treated and control groups have different socioeconomic levels that might affect the estimated coefficients. The time-variant covariates (density and salaries) are proxies for the socioeconomic differences, so in equation (4), we include them using the OLS FE estimator. Although the coefficients remain significant at the 1% level, the estimated coefficients of the interaction terms are smaller (5.6% in Panel A and 3.55% in Panel B), as expected. It is interesting to note that these estimates are like those found by Franco and Santos (2021) – 3.7%. They used the same estimator (OLS FE) but with different data and different variables (quarterly price indexes and Airbnb activity). The coefficients of the covariates have inverted signs, suggesting less precision in the estimates, as was found by Hausman and Kuersteiner (2008).

When equation (4) is estimated with the FGLS SUR estimator we arrive at much larger coefficients for the interaction terms (27.4% for Panel A and 16.1% for Panel B), with similar or smaller standard errors. The estimated coefficients of the covariates no longer have inverted signs and their standard errors are much smaller than with the OLS FE estimator. The F statistic is also much larger with FGLS SUR and confirms the findings of Brewer *et al.* (2017) regarding higher power tests when using FGLS SUR.

Since the liberalization of short-term rental registration in August 2014, housing prices have increased by 108% in the Lisbon municipality and 98% in the Porto municipality (as of December 2019). These municipalities have the largest shares of housing transferred to STR (over 5%) of the sample. We believe that the coefficients estimated in equation (4) with the FGLS SUR estimator reflect more precisely the impact on housing prices of a supply shock, especially in municipalities with low housing price elasticity such as Lisbon and Porto.

## 5. Robustness tests

We ran the specifications with all the municipalities of the MSA, splitting them into two quantiles: above-median (upper quantile) and below-median (bottom quantile). In the Lisbon MSA, we excluded the Oeiras municipality from the sample, as it is the odd number at the median. Table 6 shows the estimated coefficients $\beta$ for the specifications with the median quantiles.

All coefficients, although smaller than in the upper and lower quartile regressions, remain statistically significant at the 1% confidence level (except for Salaries, in equation (4) with OLS FE of Panel A). The results are robust to an enlarged sample with less extreme values.

## 6. Conclusions

Our work contributes to the important debate about the impact of STR on housing prices. We innovated by using the FGLS estimator in a DiD methodology with a SUR equation

**Table 6.** Upper and lower median quantiles sample regressions results

| Specifications<br>Estimator | Equation (2)<br>OLS FE | Equation (3)<br>OLS FE | Equation (3)<br>FGLS SUR | Equation (4)<br>OLS FE | Equation (4)<br>FGLS SUR |
|---|---|---|---|---|---|
| *Panel A – Lisbon MSA* | | | | | |
| Intercept | 6.852*** (0.001) | 6.868*** (0.019) | 6.820*** | 9.075*** (0.557) | 2.558*** (0.037) |
| Short-term rental | 0.118*** (0.005) | – | – | – | – |
| Treated × liberalization | – | 0.019*** (0.004) | 0.146*** | 0.0134*** (0.005) | 0.158*** (0.005) |
| Density | – | – | – | −0.276*** (0.053) | 0.098*** (0.001) |
| Salaries | – | – | – | −0.040 (0.059) | 0.501*** (0.005) |
| $R^2$ adjusted | 0.967 | 0.959 | 0.334 | 0.960 | 0.937 |
| F statistic | 404.025*** | 330.989*** | 867.858*** | 333.483*** | 8,473.980*** |
| Observations | 1,716 | 1,728 | 1,728 | 1,716 | 1,716 |
| *Panel B – Porto MSA* | | | | | |
| Intercept | 6.627*** | 6.633*** | 6.544*** | 20.324*** | 0.749*** |
| Short-term rental | 0.189*** | – | – | – | – |
| Treated × liberalization | – | 0.062*** | 0.265*** | 0.037*** | 0.116*** |
| Density | – | – | – | −0.912*** | 0.062*** |
| Salaries | – | – | – | −1.058*** | 0.783*** |
| $R^2$ adjusted | 0.960 | 0.942 | 0.404 | 0.948 | 0.840 |
| *F* statistic | 261.582*** | 176.039*** | 869.454*** | 194.492*** | 2,249.041*** |
| Observations | 1,283 | 1,283 | 1,283 | 1,283 | 1,283 |

**Notes:** The table reports the estimated coefficients from equations (2), (3) and (4), with the observations from the sample above-median and below-median quantiles. The dependent variable is the natural logarithm of housing prices. The independent variable short-term rental is the natural logarithm of the share of tourism short-term rentals as a percentage of the housing stock with the value of one added. The interaction between tourism and liberalization is a dummy variable with the value of one if the municipality is in the upper quantile (above-median) of the variable STR at the end of 2019 and if the month of observation is after August 2014 and zero otherwise. Density is the natural logarithm of the municipality density and salaries is the natural logarithm of the municipality mean gross monthly salary. OLS FE is the ordinary least square estimator with fixed effects both at the cross-section and period. FGLS SUR is the feasible least squares estimator of the seemingly unrelated regressions. Robust standard errors clustered at the municipality level are reported in parentheses. ***indicates significance at the 1% level model instead of the baseline OLS with FE. This produced more precise estimations and higher power tests, as suggested by the recent econometric research literature (Hausman and Kuersteiner, 2008; Brewer *et al.*, 2017).

We show that liberalizing the transfer of use of real estate assets from housing to tourism without any formal requirements (such as obtaining a commercial or services permit for the real estate asset) led to an increase in housing prices. We estimate that each one percentage point increase in the share of STR results in an increase of 27.4% and 16.1% in housing prices of the Lisbon and Porto MSA municipalities in the upper quartile of tourism STR, respectively. These results are much higher than those found in previous studies using a similar methodology. In a similar study of the same geographical locations, Franco and Santos (2021) found a 3.7% increase in housing prices. We also provide evidence that STR produced a negative shift in the supply curve, in contrast to what was found in previous studies (Barron *et al.*, 2021).

We conclude that the group of municipalities in the upper quartile of the short-term rental share of the housing stock (the treated group) has much lower housing supply elasticities than the group of municipalities in the lower quartile (the control group). In the Lisbon MSA, the treated group shows an inelastic supply (0.678) and in the Porto MSA, the treated group housing supply elasticity is also low (1.885). The lower elasticity of the Lisbon MSA treated group in comparison to the Porto MSA may be linked to the higher impact on prices of an increase in the transfer of real estate from housing to STR. The housing supply elasticity in high-density locations is remarkably low, with significant changes in housing prices producing small changes in the housing supply (Green *et al.*, 2005).

Our findings suggest that policy measures aimed solely at the demand side of the equation will not be effective, due to an almost vertical supply curve. Aastveit *et al.* (2020) conclude that housing supply elasticities in the US have been declining in the past decade, mainly due to zoning laws, with demand shocks absorbed mostly by price adjustments rather than by quantity adjustments. The remedies implemented in several cities for tackling increased prices due to the increase in STR (creation of contention zones where no new STR are licensed) do not seem to address the real problem because prices do not adjust quickly in low elasticity zones. According to Somerville *et al.* (2020), a consequence of the zoning restrictions is significantly reduced transactions in the restricted zone, but the housing prices do not fall accordingly. Moreover, in a study about short-term rental zoning restrictions implemented in New Orleans neighborhoods, Valentin (2021) concludes that short-term rental usage increased in the adjacent neighborhoods, transferring the problems to other areas of the city. Our results are consistent with these findings and contribute to the policy debate about the effects of STR on housing prices by showing that supply-side interventions that help to flatten the housing supply curve will be more efficient in tackling rising housing prices. If the housing supply becomes more elastic, housing transferred to STR can be quickly replaced by new construction, without a significant impact on housing prices.

According to research reports from consultancy firms (Deloitte, 2017; Athena Advisers, 2016), a one-bedroom apartment in the central Lisbon housing market could be rented for €6,600 per year on average. If it were transferred to the tourism short-term rental market, it could generate €28,383 in gross income on average. Allowing the transfer of existing real estate from the housing market to the tourism market illustrates the Law of Unanticipated Consequences (Merton, 1936), where immediate interests override long-term interests (higher rental yields in the tourism market than in the housing market overriding affordable housing). If the transfer of real estate use from housing to a higher-yielding service leads to an increase in housing prices, policymakers should beware legislation that eases that transfer but should ease the legislation that allows the transfer of real estate use from other services to housing.

We believe there is much more to investigate regarding the impact on prices of the transfer of real estate use. Further research may include repeating the study with larger samples and investigating the transfer of real estate use from services such as offices to housing, to test whether those transfers are linked to supply shocks in either market.

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