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The effect of property taxes on house prices: Evidence from the 1993 and the 2012 reforms in Italy

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Melisso Boschi
Senate of the Republic of Italy
Centre for Applied Macroeconomic Analysis, ANU

Valeria Bevilacqua
Senate of the Republic of Italy

Carla Di Falco
Senate of the Republic of Italy

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Keywords

Policy effect evaluation, Property tax, House prices, Property tax capitalization, Tax reforms, Interrupted Time Series Analysis

JEL Classification

C32, H20, R21, R31

Address for correspondence:

(E) cama.admin@anu.edu.au

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The effect of property taxes on house prices: Evidence from the 1993 and the 2012 reforms in Italy[1]

Melisso Boschi,[2] Valeria Bevilacqua,[3] Carla Di Falco[4]

Abstract

We quantify the effect of property tax reforms implemented in Italy in 1993 and 2012 on property prices. We focus on the Italian house prices index using the Interrupted Time Series Analysis (ITSA), a statistical approach that proves to be useful when a counterfactual scenario for policy evaluation is difficult to create due to the universality of intervention. The hypothesis under test is that the two reforms caused a statistically significant discontinuity in the house prices index dynamics. We estimate two alternative versions of the ITSA model – one including only Italy, and another one including also similar European countries as control terms (France, Germany, Spain, and the UK). Property tax changes effects are reform-specific. As for the 1993 reform effect on real house prices, we estimate a 13-14 percent decrease of the mean level and a 1 percentage points (p.p.) increase of the rate of growth. As for the 2012 reform, depending on the model chosen, we estimate a 3-5 percent decrease, or a 4 percent increase, in level, as well as a 3-4 p.p. decrease, or a 2 p.p. increase, of the rate of growth.

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[2] Corresponding author. Senate of the Republic of Italy, Research Service and Working Group on Public Policies Analysis and Evaluation; Centre for Applied Macroeconomic Analysis (CAMA)

[3] Senate of the Republic of Italy, Budget Service and Working Group on Public Policies Analysis and Evaluation

[4] Senate of the Republic of Italy, Committee Service and Working Group on Public Policies Analysis and Evaluation

1. Introduction and literature review

Theory suggests that the real estate market works like any other financial market. According to the property tax capitalization hypothesis, in particular, a house's equilibrium price, as for any other asset, equals the present value of the after-tax flow of rents from owning it, given other housing market characteristics. Taxpayers would therefore anticipate fiscal liabilities pricing them in equilibrium prices. The most often cited study by Oates (1969) argues that, from a general equilibrium perspective, at the local level, the property price depend on the present value of the future stream of benefits from public services relative to the present value of future tax payments. More generally, at the national level, using an asset market approach, the seminal paper by Poterba (1984) argues that a house's price must equal the present discounted value of its net future service flow, the latter given by the rental service value minus depreciation, tax, and maintenance costs. Any change in property taxes, therefore, must affect property prices, at least theoretically, although the size of this effect must be determined empirically.

In this paper, we contribute to the literature on the relationship between property taxation and property prices by focusing on the effects of two crucial reforms implemented in Italy. The first one took effect in 1993, when Italy introduced the Municipal Property Tax (*Imposta comunale sugli immobili* – ICI from here onwards), and allocated its revenues to municipalities, as the tax name suggests. Except for small adjustments, this tax remained unaltered until 2012. The 2007-2008 international financial crisis and the recession immediately following it had strong repercussions on public finances across Europe, and especially so in Italy. As the financial crisis spread to the European bond *market triggering a sovereign debt crisis, the Italian government approved in 2012 a comprehensive fiscal austerity plan to stabilize public debt and avoid fiscal default (see Figari and Fiorio, 2015, for a survey of the main consolidation measures adopted and their budget effects). One of the main components of the plan was the anticipated implementation in 2012 of the Single Municipal Property Tax (*Imposta municipale unica* – IMU from here onwards) – the second reform considered – originally due in 2014.

The large empirical literature that followed Oates' study acknowledge that property tax liabilities negatively affect future property prices, but the extent of such capitalization remains a matter of contention. Most earlier studies focus on local level analyses (see Sirmans *et al*., 2008, for a review), but there are recent studies that take a national (see, for example, Elinder and Persson, 2017, Oliviero and Scognamiglio, 2019) or international view (see Oliviero *et al.*, 2019).

This literature, however, must deal with a serious endogeneity problem deriving from the relationship existing between any tax revenue and its tax base, as well as with issues of spurious correlation between unexplained variation in local public services and tax rates, and the measurement of effective tax rates as opposed to stated tax rates.

To deal with these issues, we resort to an empirical model by which house prices (as a proxy for property prices) do not depend on property tax revenues. We estimate the effect of the property tax reforms of 1993 and 2012 on property prices using the Interrupted time series analysis (ITSA) evaluation approach. ITSA has been widely used to infer the causal effect of a policy when a counterfactual scenario based on untreated individual is difficult to construct because the intervention is universal as it involves the entire population. In such a situation, ITSA exploits the discontinuity between pre- and post-intervention to recover the counterfactual scenario by projecting the pre-intervention trend over the intervention period.

To give more robustness to our evaluation, we estimate the model over different sample lengths and compare, within the ITSA model, the experience of Italy to that of a selection of countries having similar economic characteristics.

The key results indicate that, depending on the model considered, the property tax reforms enacted in Italy prompted on impact a reduction of real house prices level within the range 13-14 percent in 1993. As for 2012, depending on the model we find a 4 percent increase with respect to France or a 5 percent decrease with respect to Germany. As for the 2008 financial crisis, the single country model indicates a real house prices level decrease of 9 percent on impact, while the multiple countries model indicates a decrease of 10 percent in comparison to Germany and an increase within the range 7-12 percent in comparison to Spain and the UK.

With regard to the real house prices time trend, results indicate an increase by 1 p.p. in the post-1993 rate of growth with respect to the pre-intervention one. The post-2012 trend, by contrast, shows, depending on the model, a reduction in the rate of growth by 3-4 p.p. with respect to Spain or an increase by 2 p.p. in comparison to France. As for the post-2008 crisis trend, the reduction with respect to the pre-crisis trend ranges, depending on the model, between 1 p.p. and 2 p.p. if France and Germany are the control countries, while the trend increases by 3 p.p. with respect to Spain.

These results might have important policy implications. A large body of research has argued that shifting some of the tax burden away from labour and capital to immovable property would reduce the distortionary effects of taxation on work and investment decisions. This would increase economic efficiency and, therefore, economic growth (OECD, 2010, is the standard reference). Arnold *et al.* (2011) even find that the following ranking – from the most to the least harmful for economic growth – would hold: corporate taxes, personal income taxes, consumption taxes and finally property taxes (especially recurrent taxes on immovable property).

This evidence forms the basis for tax policy recommendations offered by international organizations to member countries (see European Commission, 2020a, for a recent example), and especially to Italy (see IMF, 2019, and European Commission, 2021).

Some authors, however, question this relationship between tax structure and economic growth by stressing that it eventually depends on the sample countries and years under investigation or methodology used (see, for example, Xing, 2012, Baiardi *et al.*, 2019). Xing (2012), in particular, finds that property taxes contribute to long-run GDP growth only in few countries (Finland, Ireland, and the United Kingdom).

A shift of the taxation burden to immovable property, in fact, might have an undesirable negative effect on aggregate economic activity through the real estate market. An important strand of literature points towards a correlation between real estate property prices and macroeconomic conditions (economic growth, consumption, investment, and employment).

A decrease in property values may entail a negative wealth effect that, in turn, could reduce consumption, access to credit, and investment. Rising house prices may stimulate consumption by increasing households' perceived wealth, or by relaxing borrowing constraints, as Campbell and Cocco (2006) find using UK micro data. These authors argue, however, that wealth effects are heterogeneous across age groups − older homeowners show larger responses of consumption to house prices changes in comparison to younger renters. Predictable changes in house prices are correlated with predictable changes in consumption, particularly for households that are more likely to be borrowing constrained. Along the same line, Surico and Trezzi (2019), using data from the Survey on Household Income and Wealth and exploiting the 2011 property tax reform enacted in Italy, find that a tax hike on the main dwelling leads to large expenditure cuts among mortgagors, who hold low liquid wealth despite owning sizable illiquid assets. The effect on other residential properties affect, by contrast, affluent households, so that the impact on consumer spending is modest. Chaney *et al*. (2012) focus instead on the collateral channel through which shocks to the value of real estate can have a large impact on aggregate investment. Collateral pledging enhances a firm's financial capacity, but this implies that asset liquidation values play a key role in the determination of a firm's debt capacity. From this derives that business downturns will deteriorate assets values, thus reducing debt capacity and depressing investment, which will amplify the downturn. Therefore, abruptly declining real estate prices can reduce investment through this "collateral channel". Chaney *et al*. (2012) show that over the 1993–2007 period, a $1 increase in collateral value leads the representative US public corporation to raise its investment by $0.06. The same collateral channel also works for small business employment. Adelino *et al*. (2015), in fact, show that small business in areas with greater increases in house prices experience stronger growth in employment when compared to large firms in the same areas and industries so that the collateral channel explains 15-25 percent of employment variation.

The remainder of the article proceeds as follows. In section 2 we summarize the development in the property taxation in Italy since the beginning of the '90s. In sections 3 and 4 we illustrate the empirical methodology and data used. In section 5 we discuss our results for both models estimated − that focusing on Italy and that including other countries and robustness checks. In section 6 we close with some concluding remarks.

2. Background: Recent evolution of the Italian property tax system

In this section we briefly overview the evolution of the immovable property tax system in Italy in the latter 30 years, concentrating on the recurrent property tax component.

In 1993, the government implemented ICI – the first recurrent wealth tax specifically intended to hit immovable property in Italy. The property owner was liable for ICI to the local municipality (*Comune*), which had to approve yearly the relevant tax rate and possible deductions within a certain range.[5] The tax base consisted of the cadastral property value, computed by multiplying the cadastral return, revalued by 5 percent, by a factor that varied according to the property type.

In 2008, the government exempted from ICI the taxpayer's main residence, with the exception of luxury properties.

In a quest for public finance sustainability, at the end of 2011 the Italian government introduced a package of measures aimed to counteract the sovereign crisis through fiscal consolidation. Within this package, the property tax reform had an important role. With effect from 2012, the government

[5] From 1993 to 2008, a deduction on the tax liability due on the main residence was in place.

replaced ICI with IMU, which differed from the previous system under four main aspects. First, it repealed the 2008 main residence exemption. Second, the tax base increased by 60 p.p.. Third, it set the tax rate on the main residence equal to 0.4 percent and the rate on secondary dwellings equal to 0.76 percent. Finally, it introduced a deduction of 200 euros for the main residence, plus 50 euro per household member younger than 26 (up to a maximum deduction of 400 euro).[6] Local municipalities were free to modify the first rate by +/-0.2 p.p. and the second rate by +/-0.3 points by the end of October 2012. Municipalities were also free to modify main residence deductions, but only a small fraction of them set the property tax rates below the statutory levels.

In 2013, the government exempted from IMU some types of properties (e.g. properties built by building cooperatives, social housing properties, properties used by research institutions) while it partially exempted main residences.

In 2014, IMU on the main residence was abolished and a new tax on local services (*Tributo per i servizi indivisibili* – TASI from here onwards) was introduced on both main residences and secondary dwellings. TASI, however, resembled so much IMU in terms of both the tax base and rates that total fiscal revenues on main residences were almost unaffected (see Messina and Savegnago, 2014).

Finally, TASI was abolished first on main residences (exception for luxury ones) with effect from 2016, and then altogether, with effect from 2020. Although at the time the government claimed the IMU reform to be transitory, most Italian taxpayers perceived it as permanent (see Oliviero and Scognamiglio, 2019), as it eventually turned out to be.

We now turn to an overview of the dynamics of tax revenues from ICI and IMU.

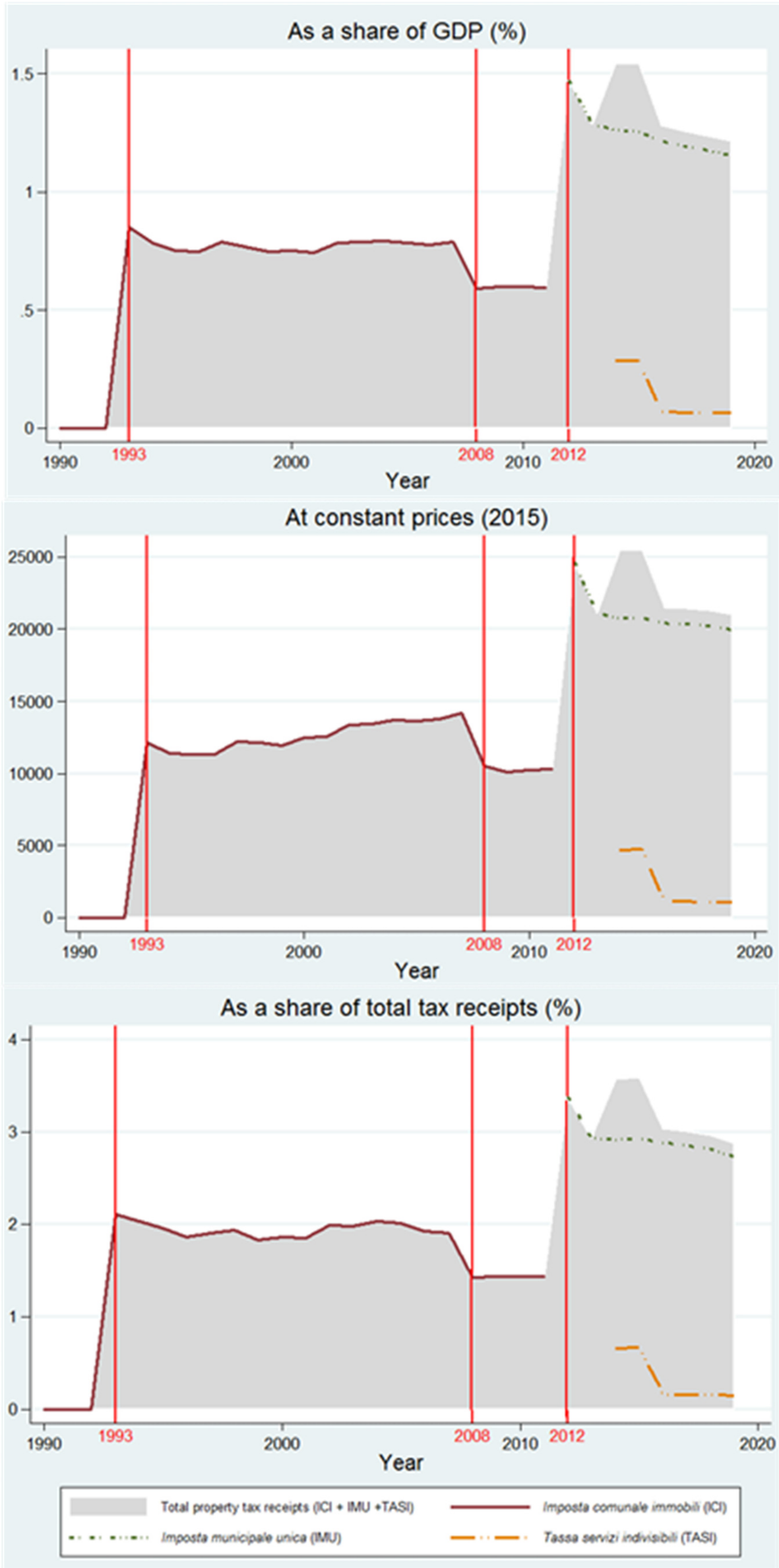
Specifically, the top graph in figure 1 shows the course of revenues, as a share of GDP, from each of the three taxes ICI, IMU, and TASI, as well as their total, at current prices over the period 1990-2019. Data on the municipal real estate tax (ICI from 1993 to 2011 and IMU from 2012 to 2019) and on the tax on indivisible services (TASI) come from the OECD database, along with data on GDP.

The graph in the middle of figure 1 shows the same revenues in absolute value at constant 2015 prices. To express revenues data at constant 2015 prices, we used the GDP deflator from the national accounts of the OECD database. Finally, the bottom graph shows the trend of the same revenues as a share of total tax revenues.

As a share of GDP (figure 1, top graph), total property tax revenues (ICI+IMU+TASI depending on years) decrease from a yearly average amount of 0.78 percent over the period 1993-2007 (before the main residences exemption from ICI) to an average of 0.60 percent over the years 2008-2011, and then increase again to 1.36 percent after the introduction of IMU (2012-2018). More in detail, revenues amounted to 0.79 percent of GDP in 2007, then 0.59 percent in 2008, and 1.47 percent in 2012. The significant decrease of ICI revenue in 2008 is mainly due to the main residence exemption. Analogously, the decrease in 2013, smaller than the former, was mainly due to the partial exemption of some types of properties.

[6] See Oliviero and Scognamiglio (2019) for more details.

Figure 1 – Property tax revenues in Italy



Notes: the top graph shows the course of revenues from ICI, IMU, and TASI, as well as their total, at current prices as a share of GDP. The graph in the middle shows the same revenues in absolute value at constant 2015 prices. The bottom graph shows the same revenues as a share of total tax receipts. Data on tax revenues and on GDP come from the OECD database.

In absolute value, at 2015 constant prices (figure 1, middle graph), total revenues from ICI, IMU, and TASI decrease from an average yearly amount of 12,7 billion euro over the years 1993-2007 to more than 10 billion euro over the period 2008-2011, and then increase again to slightly less than 23 billion euro yearly average over the period 2012-2018.

With respect to total tax receipts (figure 1, bottom graph), property tax revenues decreased from a yearly average of 1.94 percent over the period 1993-2007 to 1.43 percent in 2008-2011, and then increased again to 3.17 percent over the period 2012-2018.

Clear breaks in the revenues' dynamics stand out in years 1993, 2008, 2012, 2013, 2014, and 2016 in correspondence of the ICI reform, financial crisis, the IMU reform, the introductions of partial exemptions from IMU, the introduction and abolition of TASI on main residences.

3. Empirical methodology

When a policy intervention involves the entire population (the sample size is $N = 1$), one cannot construct a proper counterfactual scenario based on untreated individuals, i.e. individuals not subject to the policy measures. In this case, if a sufficiently long sample of observations on the variable of interest (outcome) in the pre- and post-intervention period is available, interrupted time series analysis (ITSA) represents a quasi-experimental method with a potentially high degree of internal validity (see Campbell and Stanley, 1966, Shadish *et al*., 2001) to infer the causal effects of a policy.[7] A strong internal validity of the ITSA approach, even in the absence of a comparison group, also derives from its control over the effects of regression to the mean (see Campbell and Stanley, 1966).

More specifically, ITSA exploits the time discontinuity between pre- and post-intervention to project the pre-intervention trend over the post-intervention period to serve as a counterfactual scenario[8]. The policy impact estimation results from the comparison of the outcome variable with the counterfactual scenario over the post-intervention period.

As we briefly anticipated above, we must emphasize that this method is better suited to estimate the impact of property taxes than those that include property tax revenues among the independent variables, since the latter approach may entail endogeneity problems that bias results. Within these models, in fact, house prices are a function of tax revenues, which, in turn, could depend on house prices themselves as long as these represent the main component of the property tax base. Lutz (2008), for example, estimates a 0.4 percent elasticity of property tax revenues with respect to house prices changes in the United States. This endogeneity could undermine the fundamental hypotheses of the OLS regression model, although this problem could be less serious in those tax

[7] For examples of ITSA applications to policy evaluation, see Nunn and Newby (2011) for traffic regulation, Bernal *et al*. (2017) for health, Humphrey *et al*. (2017) for self-defence, De Jorge-Huertas and De Jorge-Moreno (2020) for house prices.

[8] The term "interrupted" derives from the expectation of such discontinuity in the level or trend of the time series subsequent to the intervention (see Campbell and Stanley, 1966, Shadish *et al*., 2001).

systems in which the tax base is calculated basing on the cadastral rent rather than the property market value (see Oliviero *et al*. 2019). Even in these cases, however, one has to consider that new houses' cadastral rents are determined, at least initially, basing on the current market value so that the endogeneity problem emerges again. In the ITSA model, we do not need to include the property tax revenues among the model independent variables.

We also notice that our methodology gives an estimate of the overall effect of each property tax reform introduced in Italy. The existing literature, by contrast, (Oliviero and Scognamiglio, 2019, and Oliviero *et al*., 2019, being the most direct reference) gives results related to the effect of an increase of the growth rate of property tax revenues or one standard deviation increase in property tax intensity on house prices. When trying to isolate fiscal reforms effects, Oliviero *et al*. (2019) need to resort to the comparison of 3-year moving averages with long-run historical averages of tax revenues, which might easily blur the effect of specific reform events. Although this kind of results could seem appealing in view of its supposed external validity, one should bear in mind the inherent complexity of such important reforms, whose design and implementation inevitably imply so many institutional features that make their effect difficult to generalize to external settings and scenarios. Therefore, our approach focused on the overall reform effect seems to us more pragmatic and adherent to the real world. Results corroborate this approach as they show reform-specific effects of property tax changes on property prices. Elinder and Persson's (2017) Differences-in-Differences framework is more akin to ours since it allows focusing on the 2008 Swedish property tax reform. In our paper, however, we compare more reform events.

Moreover, our approach allows circumventing limitations due to the presence of spurious correlation between the unexplained variation in the quantity and quality of local public services and the tax rates, as well as issues related to the difficulty of measuring effective property tax rates (see Palmon and Smith, 1998, for a thorough explanation of these issues).

Furthermore, in contrast to Oliviero et al. (2019), the ITSA method allows us estimate level effects as well as the trend effect.

Finally, the ITSA method is better suited to determine specific policy driven shifts, so that we estimate the effects of specific reforms rather than relationships estimated by averaging across time, countries, and reform episodes.

To sum up, we assume that the outcome variable (the house prices index as a proxy for overall property prices) evolves according to the following model:

$$y_t = \beta_0 + \beta_1 T_t + \beta_{2,year} X_{t,year} + \beta_{3,year} X_{t,year} T_t + \epsilon_t \tag{1}$$

where $y_t = \ln(Y_t)$ indicates the natural logarithm of the house price index level – the aggregated outcome variable – at time t, T_t indicates the time variable, $X_{t,year}$ is a binary (dummy) variable taking value zero over the pre-intervention years and 1 otherwise, $X_{t,year}T_t$ is an interaction term changing with the intervention period, and ϵ_t is a stochastic error term.

As a further benefit, the ITSA method can accommodate for multiple treatment periods of the outcome variable dynamics, either due to policy interventions or to other exogenous events. The 2007-2008 financial crisis had a relevant effect on the Italian real estate market, as well as the financial and credit system (see Baldini and Poggio, 2014, for an illustration). The 2012 property tax reform took place in an economic environment already heavily shocked by the financial crisis. We thus consider year 2008 – conventionally the starting year of the financial crisis in Italy, only a few months after the crisis began in the US – together with years 1993 and 2012 – in which the property tax reforms took effect – as the discontinuity or intervention years. From here onwards, in this paper we consider "intervention" either the introduction of a new policy measure and an exogenous event (like the financial crisis) whose effect we want to measure.

In our model specification, therefore, we have 3 intervention variables: $X_{t,1993}$, $X_{t,2008}$ e $X_{t,2012}$.

The intercept β_0 represents the starting (conditional) mean level of the outcome variable while the coefficient β_1 represents the slope with respect to time T_t – that is the average annual percent change – of the outcome variable in the pre-intervention period. More interestingly, the coefficient $\beta_{2,year}$ indicates the change in the outcome variable mean level immediately after an intervention (compared with the counterfactual), and $\beta_{3,year}$ indicates the difference between pre-intervention and post-intervention slopes of the outcome.

A statistically significant value of β_2 would imply an immediate impact of the intervention on the house price level, while a statistically significant value of β_3 would imply that a house prices index change took place over time.

Therefore, a statistically significant value of $\beta_{2,1993}$ indicates that the ICI reform had an immediate impact of $\beta_{2,1993}$ percent points on house prices. A statistically significant value of $\beta_{3,1993}$, instead, indicates that the ICI reform had an impact on the average annual percent change of house prices of $\beta_{3,1993}$ percent points. A similar interpretation holds for coefficients $\beta_{2,2008}$ and $\beta_{3,2008}$, related to the financial crisis, and for $\beta_{2,2012}$ e $\beta_{3,2012}$, related to the IMU reform. Clearly, coefficients $\beta_{2,year}$ and $\beta_{3,year}$ measure changes in the intercept and slope with respect to the counterfactual.

We have thus four different model specifications depending on the value that the relevant dummy variable assumes in each period. The following table 1 reports a summary of the four model specifications.

Table 1 – The model specification in each sub-sample

Sub-sample	Dummy variables value	Relevant model
1970-1992	$X_{t,1993}=0,\ X_{t,2008}=0, X_{t,2012}=0$	$y_t=\beta_0+\beta_1 T_t+\epsilon_t$
1993-2007	$X_{t,1993}=1,\ X_{t,2008}=0, X_{t,2012}=0$	$y_t = \beta_0+\beta_1 T_t + \beta_{2,1993}X_{t,1993} + \beta_{3,1993}X_{t,1993}T_t+\epsilon_t$
2008-2011	$X_{t,1993}=1,\ X_{t,2008}=1, X_{t,2012}=0$	$y_t = \beta_0+\beta_1 T_t + \beta_{2,1993}X_{t,1993} + \beta_{3,1993}X_{t,1993}T_t + \beta_{2,2008}X_{t,2008} + \beta_{3,2008}X_{t,2008}T_t+\epsilon_t$
2012-2019	$X_{t,1993}=1, X_{t,2008}=1, X_{t,2012}=1$	$y_t = \beta_0+\beta_1 T_t + \beta_{2,1993}X_{t,1993} + \beta_{3,1993}X_{t,1993}T_t + \beta_{2,2008}X_{t,2008} + \beta_{3,2008}X_{t,2008}T_t + \beta_{2,2012}X_{t,2012} + \beta_{3,2012}X_{t,2012}T_t+\epsilon_t$

Note: This table shows, for each sub-sample, the model specification corresponding to the value of the intervention dummy.

When a comparison group is available, the researcher can further enhance the ITSA methodology's internal validity by controlling for confounding omitted variables (see Linden, 2015 and 2017).

To obtain results that are more robust, therefore, we extend the model illustrated so far to include other countries to serve as comparison terms. The ITSA model becomes as follows:

$$\boldsymbol{y}_t=\beta_0+\beta_1 T_t+\beta_{2,year}X_{t,year}+\beta_{3,year}X_{t,year}T_t+\beta_4 Z+\beta_5 ZT_t+\beta_{6,year}ZX_{t,year}+\beta_{7,year}ZX_{t,year}T_t+\epsilon_t \qquad (2)$$

where $\boldsymbol{y}_t$ is a vector of the logarithm of house prices' indices of Italy and the other countries, the dummy variable Z denotes the assignment to the treatment cohort ($Z=1$ for Italy in our case) or to the control cohort ($Z=0$ for the other country), while the terms ZT_t, ZX_t, and ZX_tT_t are interaction terms between the variables already described above. In this model, therefore, coefficients β_0, β_1, $\beta_{2,year}$, and $\beta_{3,year}$refer to the control country, while coefficients β_4, β_5, $\beta_{6,year}$, and $\beta_{7,year}$ refer to the treatment country. More specifically, the coefficient β_4 represents the difference between the two countries' intercept in the pre-intervention period, β_5 represents the difference between the two countries' time slope in the pre-intervention period, $\beta_{6,year}$ indicates the difference between the two countries' intercept immediately after the date in which the intervention takes place, and $\beta_{7,year}$ indicates the difference between the two countries' time slope after the intervention date compared with pre-intervention. This last coefficient is similar to the slope in a difference-in-differences model. Therefore, the counterfactual construction method illustrated above along with the inclusion of other countries as controls makes the ITSA method similar to a difference-in-differences model.

As Linden (2015) explains, an ITSA model with control individuals proves especially useful when there is an exogenous event that affects all groups, as the financial crisis is in our set up. One crucial hypothesis on which the whole analysis is based upon is that the change in the outcome variable intercept or time trend would have taken place in the same way both in the control country and in the treatment country in case the latter had not been treated. This requires that the two countries, treated and control, are structurally similar to each other, at least with regard to the sector related to the outcome variable (the real estate sector in our case) and that any differences is only induced by the intervention.

In order to check for residual autocorrelation, we use the Cumby-Huizinga (C-H, 1992) general test for autocorrelation.[9]

4. Data

Our outcome variable, Y_t, is the OECD Residential Property Prices Index (RPPI) – also named House prices index (HPI) – an index number, with 2015 as base year that measures the prices of residential properties, both old and newly built, over time.[10] House prices data are in real terms – we deflate nominal prices using the private consumption expenditure deflator from the national account statistics – at annual frequency, and seasonally adjusted. We use the HPI as a proxy for the overall property prices.

We use the entire available sample period 1970-2019 in order to allow for a sufficiently long series before the three interventions take place (1993, 2008, and 2012) to estimate an accurate trend behaviour. Ending the sample in 2019 allows us to avoid the COVID-19 pandemic crisis and its disrupting effects on the economy, which could distort the intervention effects.

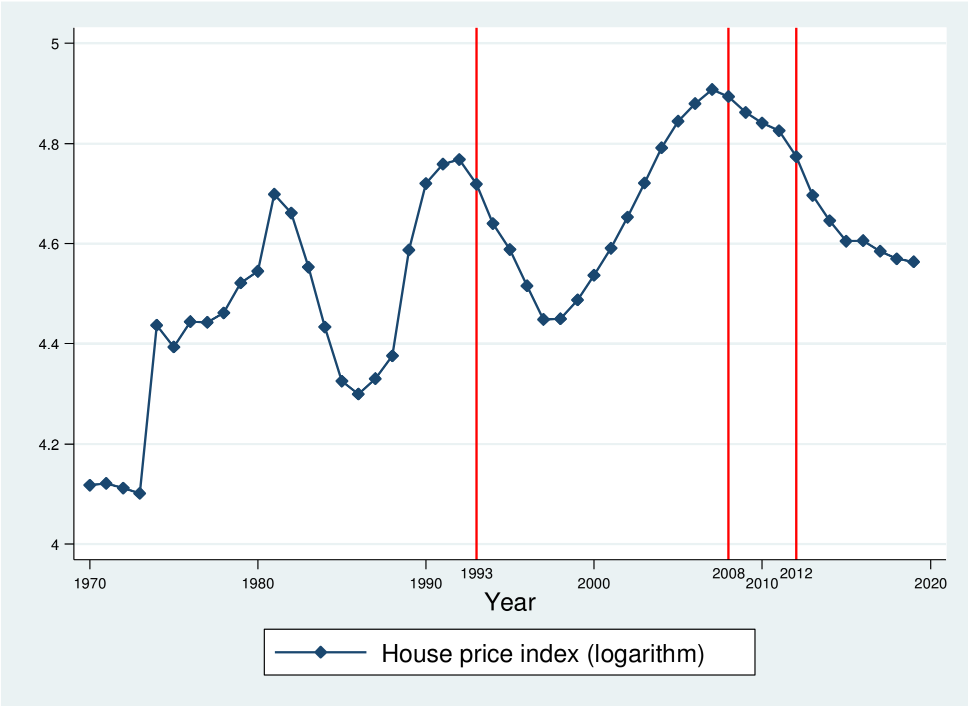
Figure 2 below shows the course of HPI together with real GDP over the last five decades, while figure 3 shows HPI in log-level.

[9] Following Linden (2015), we use the *actest* Stata module developed by Baum and Schaffer (2013).
[10] The data are downloaded from the OECD database website
http://stats.oecd.org/Index.aspx?DataSetCode=HOUSE_PRICES

Figure 2. The house prices index and GDP in Italy

Notes: this figure shows the course of the real house prices index (HPI) and real GDP in Italy over the period 1970-2019. The vertical red lines indicate the intervention years 1993, 2008, and 2012

Figure 3. The logarithm of the real house prices index in Italy



Notes: This figure shows the course of the log-level real house prices index (HPI) in Italy over the period 1970-2019. The vertical red lines indicate the intervention years 1993, 2008, and 2012

The real house prices index level cyclical pattern appears much more marked than the GDP one. The cyclical turning points of the two series, however, appear to coincide in time.

Summary statistics reported in Table 2 show that, over the longest available period (50 observations between 1970 and 2019) the index level ranges between 60.4 and 135.4, with a mean value of 98.4 and a standard deviation of 19.6. Table 2 also reports the summary statistics of the natural logarithm of the real house prices index used to estimate the ITSA model, as well as the summary statistics, for the HPI level and logarithm, for each of the relevant sample sub-periods.

According to the unit root ADF (Augmented Dickey-Fuller) and KPSS (Kwiatkowski, Phillips, Schmidt e Shin, 1992) tests (unreported to save space), the log index appears to be non-stationary. Given, however, that we include no independent variables in the model other than the dummy variables, a time trend, and the lagged dependent variable, we prefer to use the logarithm of real HPI instead of its first difference to avoid wasting important statistical information. This also allows us to interpret results, quite conveniently, as percent values. Moreover, using logarithm facilitates results interpretation and comparability with previous literature.

Table 2. Summary statistics of the logarithm of the real house prices index for Italy

Sub sample	Variable	Number of observations	Mean	Standard deviation	Minimum	Maximum
1970-2019	Prices (level)	50	98.5	19.6	60.4	135.4
	Prices (log)	50	4.6	0.2	4.1	4.9
1970-1992	Prices (level)	23	86.8	17.6	60.4	117.7
	Prices (log)	23	4.44	0.2	4.1	4.7
1993-2007	Prices (level)	15	105.9	16.6	85.5	135.4
	Prices (log)	15	4.7	0.2	4.4	4.9
2008-2011	Prices (level)	4	128.5	3.8	124.7	133.4
	Prices (log)	4	4.9	0.0	4.8	4.9
2012-2019	Prices (level)	8	102.8	7.7	96.0	118.4
	Prices (log)	8	4.6	0.1	4.6	4.8

Notes: This table reports the summary statistics of the Italian real house price index (in level and logarithm) over the period 1970-2019 as well as the other relevant sub-periods used in the analysis

We see that the cyclical behaviour of log prices appears to mimic the economy business cycle phases. The cyclical turning points correspond to years 1993, 1998, and 2008.[11] Three red vertical lines indicate the 2008 financial crisis cyclical turning point and the 1993 and 2012 policy intervention years on which we focus in this study.

5. Results

In this paragraph, we present OLS estimation results for models (1) and (2).[12]

5.1 The ITSA model for Italy

We first estimated model (1), which includes only data for Italy, without any comparison with other similar countries.

In order to improve the model fit to data and obtain satisfactory residuals in terms of autocorrelation and normality properties, we included among the independent variables a dummy variable that takes value 1 for years 1983 and 1984 and 0 otherwise, as well as a lag for the dependent (outcome) variable.[13][14]

In order to understand for how long the policy intervention produced its effects, if any, we estimated the model on a rolling sample, adding one period at a time from 2013 to 2019, so that we eventually

[11] Bulligan (2010) shows that, although the Italian real house prices are uncorrelated with GDP at lag 0, they follow the economic cycle with a two-year delay.

[12] We use the *itsa* command for Stata developed by Linden (2015). The code of *itsa* relies on OLS rather than ARIMA models because of the flexibility and applicability the former allows in an interrupted time-series context (see Linden, 2015).

[13] The ITSA model is based on the assumption that any other context variables that affects the outcome variable change slowly over time around the intervention period. This is obviously a bold assumption that one could avoid by including a number of control variables that account for the macroeconomic environment in which the real estate market operates. However, this appears impossible given the limited number of observations (only 50 annual data points) and the already large number of estimated coefficients (10).

[14] In order to try and take into account the relevant cyclical component of house prices data, we also estimated a version of the model including a quadratic trend, which, however, turned out not to be statistically significant.

obtained seven OLS regression results, one for each sample length following the last policy intervention year (2012), as shown in table 2.

Table 3 – Results of model (1) estimation

	Model 1 (1970 - 2013)	Model 2 (1970-2014)	Model 3 (1970-2015)	Model 4 (1970-2016)	Model 5 (1970-2017)	Model 6 (1970-2018)	Model 7 (1970-2019)
$\hat{\beta}_0$	0.69	0.69	0.69	0.69	0.69	0.69	0.69
	(0.49)	*(0.49)*	*(0.49)*	*(0.49)*	*(0.48)*	*(0.48)*	*(0.48)*
$\hat{\beta}_1$	0.00	0.00	0.00	0.00	0.00	0.00	0.00
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*
$\hat{\beta}_{2,1993}$	-0.13**	-0.13**	-0.13**	-0.13***	-0.13***	-0.13***	-0.13***
	(0.05)	*(0.05)*	*(0.05)*	*(0.05)*	*(0.05)*	*(0.05)*	*(0.05)*
$\hat{\beta}_{3,1993}$	0.01**	0.01**	0.01**	0.01**	0.01**	0.01**	0.01**
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*
$\hat{\beta}_{2,2008}$	-0.09***	-0.09***	-0.09***	-0.09***	-0.09***	-0.09***	-0.09***
	(0.02)	*(0.02)*	*(0.02)*	*(0.02)*	*(0.02)*	*(0.02)*	*(0.02)*
$\hat{\beta}_{3,2008}$	-0.01***	-0.01***	-0.01***	-0.01***	-0.01***	-0.01***	-0.01***
	(0.01)	*(0.01)*	*(0.01)*	*(0.01)*	*(0.01)*	*(0.01)*	*(0.01)*
$\hat{\beta}_{2,2012}$	-0.03***	-0.04***	-0.04***	-0.05***	-0.05***	-0.05***	-0.05***
	(0.01)	*(0.01)*	*(0.01)*	*(0.02)*	*(0.02)*	*(0.01)*	*(0.01)*
$\hat{\beta}_{3,2012}$	-0.03***	-0.01	-0.00	0.01	0.01	0.01	0.01
	(0.00)	*(0.01)*	*(0.01)*	*(0.01)*	*(0.01)*	*(0.00)*	*(0.00)*
$\hat{\beta}_1 + \hat{\beta}_{3,1993}$	0.01***	0.01***	0.01***	0.01***	0.01***	0.01***	0.01***
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*
$\hat{\beta}_1 + \hat{\beta}_{3,1993} + \hat{\beta}_{3,2008}$	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00	-0.00
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*	*(0.00)*
$\hat{\beta}_1 + \hat{\beta}_{3,1993} + \hat{\beta}_{3,2008} + \hat{\beta}_{3,2012}$	-0.03***	-0.01	-0.00	0.01	0.00	0.00	0.00
	(0.01)	*(0.01)*	*(0.01)*	*(0.01)*	*(0.01)*	*(0.01)*	*(0.00)*
$y_t(-1)$	0.85***	0.84***	0.84***	0.85***	0.85***	0.84***	0.84***
	(0.11)	(0.11)	(0.11)	(0.11)	(0.11)	(0.11)	(0.11)
$dummy\_83-84$	-0.14***	-0.14***	-0.14***	-0.14***	-0.14***	-0.14***	-0.14***
	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)	(0.02)
Number of observations	43	44	45	46	47	48	49
F	194.05***	188.17***	275.46***	201.87***	231.18***	291.13***	392.81***
R^2	0.90	0.90	0.90	0.90	0.90	0.90	0.90

Note: this table reports the relevant estimated coefficients of model (1) for Italy. Each column corresponds to a post-2012 sample length varying from 1 (sample stops in 2013) to 7 (sample including observations from 2013 to 2019). The rows 9, 10, and 11 report the value of the time coefficient referred to periods 1993-2007, 2008-2011, and 2012-final year corresponding to the column. The symbols "*", "**" and "***" indicate a level of statistical significance equal to, respectively, 10%, 5%, and 1%. The bottom two rows of the table report the main statistics of the model goodness of fit to data

Each column reports the estimation results for each version of the model with the sample length indicated in the column heading. The F and R^2 statistics reported at the bottom of the table show a satisfactory goodness of fit. The inclusion of the lagged outcome variable allows to account for the correct autocorrelation features of the models, as shown by the C-H test performed with *actest* (Baum and Schaffer, 2013), according to which autocorrelation is present at lag 1 but not at any higher lag orders (up to the six lags tested).[15]

As shown in table 3, regardless of the post-intervention sample length, the pre-intervention conditional mean level of the log real HPI is 0.69 – but not statistically significant – while the model slope is basically zero.

All the impact coefficients estimates ($\hat{\beta}_{2,year}$) are statistically significant, although their economic dimension varies according to the specific intervention period considered. The estimated changes in slope ($\hat{\beta}_{3,year}$) are mostly statistically significant, except for coefficient $\beta_{3,2012}$ in models estimated on the sample ending in 2014, 2015, 2016, 2017, 2018, and 2019.

The real estate market replied to the ICI reform of 1993 with a significant decrease in real house prices of 13 percent on impact – that estimate remains stable across the various sample length specifications.[16] As for the slope, the post-1993 average growth rate of real HPI is 1 p.p. higher than the counterfactual one. This might be due to the strong real estate market expansion that took place between the end of the '90s and 2008, highlighted in figure 2, prevailing statistically over the negative impact of the 1993 reform – after the reform's negative impact, the house market recovered at a sustained pace to regain lost positions.

The financial crisis of 2008 brought about a statistically significant negative effect, both on impact (9 percent decrease in the first year of intervention) and on the annual trend (1 p.p. lower than the counterfactual). After the 2008 crisis, therefore, the annual trend of house prices got back to the pre-1993 level. The financial crisis effect seems to have prevailed over the main residence exemption from ICI enacted in 2008.

Finally, the real estate market reacted on impact to the 2012 IMU reform with a statistically significant decrease within the range 3-5 percent, depending on the model sample length. The slope effect is statistically significant only in the model whose sample ends in 2013, with an estimated annual trend lower than the counterfactual one by 3 p.p.. Over the bigger sample length, the slope effect vanishes. This might be due to the recent real estate market expansion.

5.2 Robustness check: The ITSA model with control countries

In this subsection, we switch to a multi-group design to provide a robustness check by illustrating results from the estimation of model (2).

As a preliminary issue, one has to choose correctly the control countries − that is countries that have not experienced similar property tax reforms in the same period and are comparable to Italy on both baseline level and trend of the outcome variable, as explained in section 3. In the context of the ITSA approach applied in this paper, therefore, comparable countries are those for which the

[15] These results are unreported to save space, but available on request.

[16] As it is reasonable to expect, increasing the length of the sample with the most recent data observations affects mainly the estimate of the most recent intervention considered - the 2012 reform.

estimation of model (2) gives a value of $\hat{\beta}_4$ and $\hat{\beta}_5$ not statistically different from zero at the 5 percent significance level (i.e. having a *p*-value greater than 0.05). In this case, in fact, the model would indicate that the treated country (Italy) is comparable to the control country as for the mean and annual rate of change of the house prices index in the pre-intervention period (Linden, 2015). The underlying assumption is that other relevant factors affect the property market in the two countries in similar fashion, so that only the policy (property tax reform) under study differentiates the outcome variable dynamics.[17]

Following this procedure, we chose the comparison countries by estimating model (2) for each of the OECD countries considered as a control. The following countries were good comparison terms with respect to Italy over the pre-intervention period – that is, the corresponding models show values of $\hat{\beta}_4$ and $\hat{\beta}_5$ not statistically different from zero at the 5 percent significance level: Australia, Belgium, Canada, Finland, France, Germany, Ireland, Japan, Netherlands, New Zealand, Norway, Spain, Sweden, Switzerland, the UK, and the US. To keep the analysis manageable, we narrow the choice further by picking those countries that appear to be better comparable to Italy in terms of economic dimension, trade, and financial linkages, and because of their membership of the European Union. We therefore concentrate on France, Germany, Spain, and the UK. Table 4 shows the estimation results of model (2) considering each of these control countries in turn.

The models show an adequate goodness of fit as measured by the statistics F and R^2. The results of the C-H test (unreported) indicate that the lag of order 1 of the dependent variable is sufficient to solve autocorrelation problems present in the models. The (log) levels of control countries' HPI, as well as the Italian one, appear integrated of order 1. However, as in the case of the model (1) estimation, we reckoned that using variables in first difference could waste useful information and could blur the readability of results. Besides, a cointegration analysis would be unfeasible since regressors include only dummy variables, a time trend, and an intercept.

We illustrate results reported in table 4 by focusing on the coefficients that are most relevant for the comparison with the control countries, i.e. $\hat{\beta}_6$ − difference between the two countries' intercept immediately after the intervention year, and $\hat{\beta}_7$ − difference between the two countries' time slope after the intervention year compared with pre-intervention. We analyse these coefficients for each of the intervention dates considered, i.e. 1993, 2008, and 2012, thus complementing results reported in table 3.

The Italian real estate market reacted sharply to the 1993 reform − with a 14 percent decrease of the real HPI on impact − only if compared to Germany (see the value of $\hat{\beta}_{6,1993}$). The reform effect on the real HPI level is, by contrast, statistically insignificant if appraised with respect to France, Spain, and the UK. Again in comparison with Germany, however, the Italian real HPI shows a statistically significant 1 p.p. increase in post-1993 rate of growth with respect to pre-intervention (see the value of $\hat{\beta}_{7,1993}$). In the post-1993 period, Italy's real HPI rate of growth is 2 p.p. larger than in Germany, while it is more than 1 p.p. lower than in Spain and the UK (see the post-1993 trend difference indicated by the coefficient $\hat{\beta}_5 + \hat{\beta}_{7,1993}$).

[17] Other, more sophisticated, approaches could be the "synthetic control" method introduced by Abadie and Gardeazabal (2003), which we leave for future reasearch.

As for the financial crisis of 2008, the estimation of coefficient $\hat{\beta}_{6,2008}$ points to an immediate impact, albeit with different signs, only in comparison with Germany, Spain, or the UK. Specifically, Italy's real HPI after 2008 is 10 percent lower than in Germany, while it is 7 percent higher than in Spain and 12 percent higher than in the UK. As for the post-2008 trend effect (given by $\hat{\beta}_{7,2008}$), results point to a real HPI rate of growth in Italy lower than pre-intervention by 2 p.p. if compared to France and Germany, while it is higher by 3 p.p. if compared to Spain. The overall difference in the post-2008 trend between Italy and the control countries (given by $\hat{\beta}_5 + \hat{\beta}_{7,1993} + \hat{\beta}_{7,2008}$) is -3 p.p. for France, -1 p.p. for Germany, and +2 p.p. for Spain.

In comparison to 1993 and 2008 interventions, the 2012 IMU reform had a smaller impact on the Italian real HPI. If one uses Germany as a control country the estimated impact is -5 percent. When compared to France, the real HPI increases in Italy by 4 percent (see $\hat{\beta}_{6,2012}$). The post-2012 rate of growth is 2 p.p. higher than pre-2012 if compared to France, while it is 4 p.p. lower when compared to Spain (see $\hat{\beta}_{7,2012}$). The effect is null with respect to Germany and the UK. The overall difference between the post-2012 trend in Italy and the control countries is significant only if we look at Spain − in Italy the rate of change is 2 p.p. lower than Spain (see $\hat{\beta}_5 + \hat{\beta}_{7,1993} + \hat{\beta}_{7,2008} + \hat{\beta}_{7,2012}$).

It is worth noticing that the main residence exemption from ICI in 2008 may have helped the real HPI to get back to an increasing path after the 2008 financial crisis. By contrast, the main residence taxation at the beginning of the 2012 IMU reform implementation may have exacerbated the HPI dynamics afterwards.

Comparing these results with the existing literature, even with the closest one that focuses on national level relationships between property taxes and prices, is difficult. As anticipated in the introduction, Oliviero and Scognamiglio (2019) and Oliviero *et al*. (2019), as closest examples to our study, give results related to the effect of an increase of the growth rate of property tax revenues or one standard deviation increase in property tax intensity on house prices. When trying to isolate fiscal reforms effects, Oliviero *et al*. (2019) resort to the comparison of 3-year moving averages with long-run historical averages of tax revenues. In this paper, by contrast, we estimate the effect of specific reforms (1993 and 2012) or exogenous events (2008), rather than an average effect. However, our results appear broadly in line with the studies mentioned above, since Oliviero *et al*. (2019) find that a significant increase in the property tax revenues growth rate makes house prices grow at a rate 2.5-3.1 p.p. lower than normal. Oliviero and Scognammiglio (2019) find that a one standard deviation increase in property tax intensity reduces property values by 2.7 percent in the 2012, the year of the IMU reform. By contrast, Elinder and Persson (2017), using an approach closer to our own, find no support for the property tax capitalization hypothesis since they show that Swedish house prices did not respond to a large property tax cut enacted in Sweden in 2008.

Table 4 – Results of model (2) estimation

	France (1970 - 2019)	Germany (1970-2019)	Spain (1971-2019)	United Kingdom (1969-2019)
$\hat{\beta}_0$	0.58**	0.76	0.50**	0.70***
	(0.28)	*(0.46)*	*(0.24)*	*(0.22)*
$\hat{\beta}_1$	0.00	-0.00	0.01*	0.01
	(0.00)	*(0.00)*	*(0.00)*	*(0.00)*
$\hat{\beta}_4$	0.09	-0.04	0.13	0.23*
	(0.07)	*(0.04)*	*(0.09)*	*(0.13)*
$\hat{\beta}_5$	0.00	0.00	-0.00	-0.00
	(0.00)	*(0.00)*	*(0.01)*	*(0.01)*
$\hat{\beta}_{2,1993}$	-0.07**	0.02	-0.11**	-0.06
	(0.03)	*(0.01)*	*(0.05)*	*(0.05)*
$\hat{\beta}_{3,1993}$	0.02***	-0.00**	0.01***	0.02***
	(0.00)	*(0.00)*	*(0.00)*	*(0.01)*
$\hat{\beta}_{6,1993}$	-0.05	-0.14***	-0.01	-0.07
	(0.05)	*(0.05)*	*(0.06)*	*(0.07)*
$\hat{\beta}_{7,1993}$	-0.01	0.01***	-0.00	-0.01
	(0.00)	*(0.00)*	*(0.01)*	*(0.01)*
$\hat{\beta}_{2,2008}$	-0.14***	0.02*	-0.16***	-0.20***
	(0.04)	*(0.01)*	*(0.02)*	*(0.03)*
$\hat{\beta}_{3,2008}$	0.01	0.01***	-0.04***	-0.01
	(0.01)	*(0.00)*	*(0.01)*	*(0.02)*
$\hat{\beta}_{6,2008}$	0.05	-0.10***	0.07**	0.12***
	(0.04)	*(0.03)*	*(0.03)*	*(0.03)*
$\hat{\beta}_{7,2008}$	-0.02**	-0.02***	0.03***	-0.01
	(0.01)	*(0.01)*	*(0.01)*	*(0.01)*
$\hat{\beta}_{2,2012}$	-0.09***	0.00	-0.05	0.02
	(0.02)	*(0.01)*	*(0.04)*	*(0.05)*
$\hat{\beta}_{3,2012}$	-0.02*	0.01	0.05***	-0.00
	(0.01)	*(0.00)*	*(0.01)*	*(0.02)*
$\hat{\beta}_{6,2012}$	0.04*	-0.05***	-0.01	-0.07
	(0.02)	*(0.02)*	*(0.04)*	*(0.05)*
$\hat{\beta}_{7,2012}$	0.02**	0.00	-0.04***	0.01
	(0.01)	*(0.01)*	*(0.01)*	*(0.02)*
$\hat{\beta}_5 + \hat{\beta}_{7,1993}$	-0.00	0.02***	-0.01*	-0.01*
	(0.00)	*(0.00)*	*(0.00)*	*(0.01)*
$\hat{\beta}_5 + \hat{\beta}_{7,1993} + \hat{\beta}_{7,2008}$	-0.03***	-0.01*	0.02***	-0.02
	(0.01)	*(0.00)*	*(0.01)*	*(0.01)*
$\hat{\beta}_5 + \hat{\beta}_{7,1993} + \hat{\beta}_{7,2008} + \hat{\beta}_{7,2012}$	-0.00	-0.01	-0.02***	-0.01
	(0.00)	*(0.01)*	*(0.01)*	*(0.01)*
$y_t(-1)$	0.85***	0.84***	0.86***	0.79***
	(0.07)	*(0.10)*	*(0.07)*	*(0.07)*
$dummy\_83-84$	-0.14***	-0.13***	-0.14***	-0.13***
	(0.02)	*(0.02)*	*(0.02)*	*(0.03)*
Number of observations	98	98	97	99
F	949.96***	587.19***	748.76***	796.92***
R^2	0.98	0.91	0.97	0.98

Note: this table reports the relevant estimated coefficients of model (2) for Italy. Each column corresponds to a control country. The sample final year is 2019 for every control country. The coefficients $\hat{\beta}_5 + \hat{\beta}_{7,1993}$, $\hat{\beta}_5 + \hat{\beta}_{7,1993} + \hat{\beta}_{7,2008}$, and $\hat{\beta}_5 + \hat{\beta}_{7,1993} + \hat{\beta}_{7,2008} + \hat{\beta}_{7,2012}$ indicate the difference between the pre- and post-intervention trends for each intervention year. The symbols "*", "**" and "***" indicate a level of statistical significance equal to, respectively, 10%, 5%, and 1%. The bottom two rows of the table report the main statistics of the model goodness of fit to the data

Summary and conclusions

In this paper, we document the effect on property prices (as proxied by the house prices index - HPI) of two major property tax reforms introduced in Italy in 1993 and 2012. We exploited the time discontinuity generated by reforms using the Interrupted time series analysis (ITSA), a statistical method especially appropriated to evaluate policy interventions that involve the whole population so that to overcome the lack of control groups.

To obtain robust results, we estimated two models, the first one considering only Italy's real HPI, while the second one including also data for four comparison countries – France, Germany, Spain, and the UK.

Focusing on the most conservative estimation, results indicate that, depending on the model considered, the property tax reforms enacted in Italy prompted on impact a reduction of real house prices within the range 13-14 percent in 1993. As for 2012, model (2) gives a 4 percent increase with respect to France and a 5 percent decrease with respect to Germany. As for the 2008 financial crisis, the single country model indicates a real HPI decrease of 9 percent on impact, while the multiple countries model indicates a decrease of 10 percent in comparison to Germany and an increase within the range 7-12 percent in comparison to Spain and the UK.

With regard to the effect on the real HPI time trend, the two models agree in indicating an increase by 1 p.p. in the post-1993 rate of growth with respect to the pre-intervention one, at least in comparison to Germany as for model 2. The post-2012 trend, by contrast, shows, with respect to pre-intervention, a lower rate of growth by 3 p.p. in model 1 and 4 p.p. in model 2 when Spain is the control country, whereas the rate of growth is higher by 2 p.p. in comparison to France. As for the post-2008 crisis trend, the financial crisis effect with respect to the pre-crisis trend is included between 1 p.p. reduction in model 1 and 2 p.p. in model 2 if France and Germany are the control countries, while the HPI trend increases by 3 p.p. with respect to Spain.

The neater and more economically relevant results for the 1993 reform with respect to the 2012 one is likely to reflect the innovative role that the first reform had for the Italian tax structure, in which it introduced for the first time a recurrent wealth tax specifically intended to hit immovable property. Compared to this innovation, the 2012 reform represents an adaptation of the existing property tax to the new necessities of the Italian public finances equilibrium.

These results, broadly in line with the existing comparable literature, bring support to the property tax capitalization hypothesis and might help to analyse the relationship between alternative tax structures and economic growth.

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