# Does Collateral Value Affect Asset Prices? Evidence from a Natural Experiment in Texas

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**Abstract**

This paper identifies the impact of collateral value on house prices, exploiting law changes in Texas which legalized home equity loans in 1998. The impact of this credit expansion was positive, heterogeneous and direct. The laws increased Texas house prices 3.8%; this is price-based evidence that households are credit constrained. Prices rose more in locations with inelastic supply, higher pre-law house prices, population, income and employment. These estimates reveal that wealthier households value the option to pledge their home as collateral more strongly. Further estimates indicate that the effect was direct, as variables related to house prices were unaffected.

# 1 Introduction

Households and firms often borrow against assets. Some assets are better collateral than others. Real estate is by far the largest source of collateral used by American households (FRBNY, 2016). A mortgage can be obtained at purchase and in the future through a home equity loan (HEL)[1]. This paper refers to the latter benefit, the real option to pledge a home as collateral in the future (and extract equity), as the "collateral value." An asset's price should reflect all of its benefits including the options to pledge it as collateral. However, the collateral value has never been estimated (for any asset) as it is difficult to separately identify from the cash flow.

There is strong evidence that households benefit from the option to pledge their home as collateral. Schmalz, Sraer and Thesmar, 2017 found that homeowners are more likely to become entrepreneurs than renters when local house prices rise. Hryshko, Luengo-Prado and Sørensen, 2010 found that homeowners smooth consumption better than renters after job loss or disability when local house prices are rising. Markwardt, Martinello and Sándor, 2014 found that demand for unemployment insurance fell after HELs became available in Denmark in 1992. They concluded that private insurance through housing collateral is a substitute for public insurance.

Texas has the strictest mortgage laws in the US dating back to the state's founding constitution in 1845 (McKnight, 1983; Texas Legislative Council, 2016). Before 1998, a home in Texas could only be pledged as collateral for a purchase mortgage or a mechanic's lien. Refinance mortgages were allowed up to the balance, permitting home owners to take advantage of a fall in interest rates but not to borrow an additional amount. Texas was the only state in the US with these restrictions. Proposition 8 greatly expanded the set of mortgages available to Texans beginning in 1998. Texan home owners gained access to HELs, cash-out refinance mortgages and reverse mortgages[2]. However, the total value of all liens on the home after purchase could not exceed 80% of its appraised price[3].

These law changes provide a unique source of exogenous variation; they expanded

[1]Other products that allow future home equity extraction, including cash-out refinance mortgages and reverse mortgages, will be referred to collectively as "HELs" for ease of exposition.

[2]See Appendix D for a time-line of relevant laws. Home Equity Lines of Credit were not fully legalized until 2004 and are not studied in this paper.

[3]The 80% CLTV limit is based on the appraised value of the property at the time of the future equity extraction via HEL etc. Kumar, 2017 showed these limits on home equity borrowing in Texas lowered mortgage default during the housing bust.

future HEL debt capacity without affecting purchase mortgage debt capacity. Abdallah and Lastrapes, 2012 used these law changes as a natural experiment to estimate the impact of credit constraints on consumption. They found that Texas retail sales increased significantly after the laws, lending support to the credit-constraint hypothesis. Stolper, 2015 showed that Texan homeowners spent more on their children's college education, relative to renters and homeowners in other states after the passage of the laws. While these papers studied the impact of the law changes on consumption and investment, the current paper examines the impact on prices.

In light of evidence that homeowners benefit from the option to pledge their home as collateral, this paper investigates if, and to what extent, house prices reflect this benefit. The impact of collateral value is disentangled from rents (or service flows) by exploiting this plausibly exogenous law change. The estimation requires detailed Texas house price data, which is notoriously hard to find because Texas is a non-disclosure state; recently available house price data from the FHFA (17,936 5-digit zip codes), however, makes this analysis possible.

Difference-in-differences estimates show that the impact of this credit expansion on Texas house prices was (1) positive, (2) heterogeneous and (3) direct. The law change raised Texas house prices 3.6%-4%. This result is robust across specifications and sample restrictions. The rise in house prices was gradual, and there is no evidence of an effect before implementation, as pre-trends are parallel.

The treatment effect was heterogeneous along several dimensions. The effect was smaller in elastic locations; each unit of the Saiz, 2010 measure of supply elasticity corresponds to a 1% lower rise in prices. Furthermore, zip codes with higher pre-law[4] house prices, population, income and employment saw a greater rise in prices. This is evidence that households in locations with stronger economic conditions valued this option more. While it has been shown that an expansion in purchase mortgage debt capacity has a greater impact on ex-ante lower priced properties (Landvoigt, Piazzesi and Schneider, 2015), this paper shows that an expansion in future HEL debt capacity has a greater impact on ex-ante higher priced properties.

The treatment effect could have occurred through two channels:

1. Direct Channel: the law caused a rise in demand for homeownership due to the new

[4] The values of these variables is averaged before the law change.

option allowing homes to be pledged as collateral.

2. Indirect Channel: the law affected other variables which affect house prices. For example, if the law increased consumption and investment enough to stimulate the local economy, this increase could have raised demand for housing, consequently raising the price.

Estimates show that the law change did not affect several variables related to prices including rent, population and income. This provides evidence that most of the effect was through the direct channel. However, it is impossible to know for certain whether the law affected house prices through other unobserved variables.

Finally, the law change should have increased demand for owner occupied housing since owners can pledge their home as collateral but renters cannot. On the other hand, however, the rise in house prices from this law change should have caused a reduction in demand. Estimates show Texas home ownership rates, single family building permits and population growth were unaffected by the law. These results indicate that house prices rose enough to keep the marginal buyer indifferent between owning and renting.

This paper makes three contributions:

1. It is the first to empirically identify how the option to pledge an asset – any asset – as collateral in the future affects its price.

2. It is the first *price-based* evidence that households are borrowing constrained. The treatment effect is a measure of the extent to which households value the option to borrow in the future.

3. It helps explain the high ownership rate despite the mediocre financial return. Households are willing to pay a higher price for housing (relative to the present value of its direct cash flow) because they value the option to pledge it as collateral.

While many papers study the impact of credit constraints on consumption and investment, this paper studies the impact on prices.

The remainder of the paper is organized as follows. Section 2 discusses the relevant streams of literature. Section 3 outlines a conceptual framework for interpreting how collateral value affects housing demand and prices. Section 4 discusses the empirical strategy, defends the identification and summarizes the data. Section 5 presents and discusses the results. Section 6 concludes and discusses directions for future research.

## 2 Literature Review

Many strands of literature study borrowing constraints. One strand explains why borrowers often pledge collateral and why this affects the economy. Theoretical literature has shown that information problems restrict borrowing and that pledging collateral can allow households and firms to borrow more and thus consume or invest more (Barro 1976; Hart and Moore, 1994). Furthermore, collateralized borrowing has been shown to affect the economy because falling asset prices reduce the amount households and firms can borrow, reducing consumption and investment (Bernanke and Gertler, 1989; Kiyotaki and Moore, 1997; Shleifer and Vishny, 1992).

A second strand of literature studies the relationship between housing, credit constraints and consumption[5]. Households who are, or fear they will be, credit constrained should have a stronger demand for assets that facilitate their future ability to borrow. If prices reveal information (Hayek, 1945), the magnitude of the treatment effect estimated in this paper can be seen as a measure of the extent to which households are constrained.

A third strand of literature studies the relationship between property prices and firm investment through the collateral channel. There is evidence that rises and declines in property values, which affect debt capacity, amplify firm investment in the US and Japan but not in China[6]. While firms should prefer to own assets that have better collateral value, none of these papers empirically identifies if and to what extent asset prices reflect this.

A fourth strand of literature compares the collateral value of different assets. The interest rate borrowers pay in the repo market depends on the type of collateral they use (Bartolini et al, 2011). For example, borrowers who use treasuries as collateral often borrow at "special" repo rates. Duffie, 1996 showed that "specialness" should raise the price of the underlying security by the present value of interest rate savings. The estimates below can be interpreted through the lens of Duffie, 1996: the collateral value of housing should reflect the interest rate savings on a HEL relative to an unsecured loan.

However, in addition to the interest rate savings, the collateral value should also reflect the greater debt capacity of housing via a lower margin requirement (higher LTV) relative

[5](Agarwal and Qian, 2016; Agarwal, Hadzic, and Yildirim, 2016; Bhutta and Keys, 2015; Calomiris, Longhofer, and Miles, 2013; Chen, Michaux, and Roussanov, 2013; Cooper, 2013; DeFusco, 2016; Hurst and Stafford, 2004; Leth-Petersen, 2010; Mian and Sufi, 2011)

[6](Chaney, Sraer and Thesmar, 2012; Cvijanovic, 2014; Gan, 2007; Schmalz, Sraer and Thesmar, 2017; Wu, Gyourko and Deng, 2015)

to unsecured credit and debt secured by other forms of collateral. In related work, Gârleanu and Pedersen, 2011 showed that a security's margin affects its return.

The collateral value of other assets – such as gold, patents and fine art – has also been studied (Huang, 2016; Mann, 2016; McAndrew and Thompson, 2007). In particular, Huang, 2016 contended that gold is a better source of collateral than platinum for historical and institutional reasons. For example, gold is formally recognized as collateral by the Basel Accords and is accepted as collateral by broker dealers, while platinum is not. He argued that, in times when the probability of a consumption disaster is high, agents prefer gold for its collateral benefits, which is reflected in the price.

A fifth strand of literature studies the impact of purchase mortgage leverage on house prices[7]. It is worth emphasizing that purchase mortgage leverage is different from the type of leverage studied here. A purchase mortgage can be used to buy a house and to otherwise smooth consumption (since money is fungible). The laws studied here did not change a household's ability to finance the original purchase, but rather its ability to pledge its home as collateral in the future. There is evidence that HEL debt capacity increased purchase mortgage debt capacity for some households who used second liens ("piggy-back mortgages") to avoid mortgage insurance and obtain bigger loans (Lee, Mayer and Tracy, 2012). This was not a relevant factor in Texas because of the 80% LTV limit for all liens after purchase.

# 3 Conceptual Framework

This section constructs a model to help think about how collateral value affects housing demand and house prices to help interpret the results below. Consider a household that values non-durable consumption $c_t$ and durable housing $h_t$, which depreciates at rate $\delta_t$. The household can borrow up to $\kappa_t \equiv \min\{\overline{\kappa}_t, LTV_t\}$ of its home equity, where $\overline{\kappa}_t$ is the

[7](Adelino, Schoar and Severino, 2012; An and Yao, 2016; Anenberg, Hizmo, Kung, and Molloy, 2017; Di Maggio and Kermani, 2014; Favara and Imbs, 2015; Labonne and Welter-Nicol, 2016)

legal limit and $LTV_t$ is the most lenders will lend at time $t$[8]. It solves the following problem.

$$\max\ \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t u(c_t, h_t) \qquad \text{s.t.}$$

$$c_t + p_t h_{t+1} + a_{t+1} \leq y_t + p_t h_t(1-\delta_t) + (1+r_t)a_t \tag{DBC $\lambda_t$}$$

$$-a_{t+1} \leq \kappa_t p_t h_t \tag{CC $\mu_t$}$$

where $u(c_t, h_t)$ is the household's flow utility from consumption and housing, assumed to be twice continuously differentiable and strictly concave in each argument. $\lambda_t$ is the multiplier on the dynamic budget constraint (DBC) and $\mu_t$ is the multiplier on the collateral constraint (CC)[9]. $c_t$ is numéraire and $p_t$ is the price per unit of housing. $a_t$ is the amount saved or borrowed at interest rate $r_t$ and $y_t$ is income.

The solution to the household's problem (derived in Appendix E) implies:

$$\underbrace{p_t}_{\text{price}} = \mathbb{E}_t\left[\underbrace{\beta\frac{u_c(t+1)}{u_c(t)}}_{\text{discount factor}} \times \left(\underbrace{\frac{u_h(t+1)}{u_c(t+1)}}_{\text{service flow (rent)}} + \underbrace{\frac{\mu_{t+1}\kappa_{t+1}p_{t+1}}{u_c(t+1)}}_{\text{collateral value}} + \underbrace{(1-\delta_{t+1})p_{t+1}}_{\text{resale price}}\right)\right] \tag{1}$$

Equation 1 comes from the first order conditions of the household. Following Favilukis, Ludvigson and Van Nieuwerburgh, 2017 the periodic service flow from housing can be interpreted as a cash flow equal to rent. If a household is collateral constrained, then $\mu_{t+1} > 0$, and its FOCs reflects the collateral value. Any model with a collateral constraint has this collateral value term[10] (denoted $CV_t$). Even though it is ubiquitous, this term has never been estimated in the literature as it is difficult to separately identify the service flow (or cash flow if the property is rented) from the collateral value.

In equilibrium, the only way collateral value can be positive is if there is at least one unconstrained household doing the lending with $\mu_{t+1} = 0$. Hence a general equilibrium model requires heterogeneity for $CV_{t+1} > 0$. A convenient way to model this heterogeneity is with an impatient borrower and a patient lender (Iacoviello, 2005; Kiyotaki and Moore, 1997). An alternative way to model positive collateral value is to allow interest rates to

[8]The loan to value (LTV) ratio is determined endogenously in credit markets (Geanakoplos 2009).

[9]The collateral constraint can be written in different ways. See Appendix C for a comparison.

[10]For example, see Bianchi, Boz and Mendoza, 2012, Greenwald, 2016, He, Wright and Zhu, 2015, Iacoviello, 2005 and the recent survey Piazzesi and Schneider, 2016.

be exogenous and to assume that the representative agent borrows from a deep-pocket, risk-neutral international lender (Bianchi, Boz and Mendoza, 2012).

The collateral value can be decomposed into three parts from equation 1:

$$\underbrace{CV_{t+1}}_{\text{collateral value}} = \frac{1}{u_c(t+1)} \times \underbrace{\mu_{t+1}}_{\text{desire to borrow}} \times \underbrace{\kappa_{t+1}}_{\text{fraction you can borrow}} \times \underbrace{p_{t+1}}_{\text{value of collateral}} \tag{2}$$

This decomposition shows that the value of being able to pledge housing as collateral depends on the desire to borrow and the debt capacity $\kappa_{t+1}p_{t+1}$ (the amount a home owner can borrow).

The component $\mu_{t+1}$ (the multiplier on the collateral constraint) is a measure of the desire to borrow. If there is no demand for HELs (or if markets are complete) then $\mu_{t+1} = 0$ making $CV_{t+1} = 0$. However, $CV$ can still affect property prices today if there is a desire for HELs in the future.

The component $\kappa_{t+1} \equiv \min\{\overline{\kappa}_{t+1}, LTV_{t+1}\}$ depends on (a) the amount households can legally borrow ($\overline{\kappa}_{t+1}$) and (b) the amount lenders are willing to lend ($LTV_{t+1}$). If HELs are illegal, ($\overline{\kappa}_{t+1} = 0$) equation 2 implies $CV_{t+1} = 0$, as in Texas before 1998[11]. Abdallah and Lastrapes, 2012 showed that HELs were indeed utilized by Texans after 1998.

The model above does not distinguish between the ability to pledge an asset as collateral at the time of purchase and in the future. A decomposition of purchase mortgage collateral value from HEL collateral value[12] would require both long term and short term loans as well as frictions such as adjustment and transaction costs which is beyond the scope of this paper, but is studied in Zevelev, 2017. In such a model, the borrower's HEL debt capacity is equal to $\kappa_t p_t h_t - B_t^{pm}$, where $B_t^{pm}$ is the remaining balance on the purchase mortgage used to buy the home. For example, consider a household who owns a home worth $p_t h_t = \$100k$ with a remaining balance on its purchase mortgage of $B_t^{pm} = \$50k$. If this household can borrow up to $\kappa_t = 80\%$ of the price, then its HEL debt capacity would be $80\% \times \$100k - \$50k = \$30k$.

[11] The Texas constitution set $\overline{\kappa}_t = 0$ prior to 1998, and $\overline{\kappa}_t = 0.80$ after.

[12] He, Wright and Zhu, 2015 have a model where housing can only be used as collateral for non housing consumption in a separate "KM" market with frictions.

# 4 Empirical Strategy, Identification and Data

This paper estimates the impact of the law change on various outcome variables with a generalized difference-in-differences (DID) methodology. The main analysis uses three geographically nested samples. In all samples the treated group is Texas and the control group includes all locations outside Texas. The US sample includes all zip codes with data 6 years before and after the treatment year 1998. The border sample includes zip codes near the Texas border. The border sample is constructed by finding all zip codes within a 50 mile radius of each Texas zip code, using distance data from the NBER and Census[13]. Only zip codes which include both Texas and control locations within a 50 mile radius are kept. The third and most local sample uses all zip codes in the border city Texarkana, which is at the intersection of Texas and Arkansas.

The local samples can help control for unobserved heterogeneity if houses near each other are more likely to be affected by the same variables. In particular, the border city Texarkana, can be viewed as one economy. So if the law change affected the economy on the Texas side of Texarkana, it should have also affected the economy on the Arkansas side.

This paper estimates:

$$y_{i,s,t} = \alpha_i + 1_t + \beta_{DID} Post_t Texas_s + \Gamma X_{i,s,t} + u_{i,t} \tag{static DID}$$

$$y_{i,s,t} = \alpha_i + 1_t + \sum_{k \neq 1997} \eta_k Texas_s 1_k + \Gamma X_{i,s,t} + u_{i,t} \tag{dynamic DID}$$

where $y_{i,s,t}$ is the outcome variable. In the main regressions, it is the log of real house prices. In further regressions, log real rent, log population, log real income per capita, unemployment rate, home ownership rate and log single family building permits are also considered. The index $i$ corresponds to the most local level of the outcome variable. For house price regressions, $i$ is the zip code. For population, income, employment and permits regressions, $i$ is the county Federal Information Processing Standard (FIPS) code. For rent and homeownership rate regressions, $i$ is the Metropolitan Statistical Area (MSA). The index $s$ corresponds to the level of treatment, which is the state in all regressions. $Post_t$ is the treatment period indicator variable, which is equal to 1 for $t \geq 1998$. $Texas_s$ is the treatment group indicator variable, which is equal to 1 if $s =$ "$Texas$".

There are two categories of controls: internal and external. The internal controls

[13] http://www.nber.org/data/zip-code-distance-database.html

include location and time fixed effects, as well as time trends and a lagged dependent variable (LDV). External controls include national oil prices and interest rates interacted with state dummies. Local data is not used for external controls in order to avoid the risk of including "bad controls" (Angrist and Pischke, 2009) that might be affected by the treatment.

This paper also investigates heterogeneity in the treatment effect – that is, whether the law change had a different impact in different locations. To study the sensitivity of the effect to various observable measures $H_i$, this paper estimates:

$$\begin{aligned} y_{i,s,t} = {} & \beta_0 + \beta_1 Texas_s + \beta_2 Post_t + \beta_3 H_i \\ & + \beta_{H,0} Post_t Texas_s + \beta_5 Texas_s H_i + \beta_6 Post_t H_i \\ & + \beta_H Post_t Texas_s H_i + \Gamma X_{i,s,t} + u_{i,t} \end{aligned} \tag{DDD}$$

In this specification the average treatment effect (ATE) is an affine function of $H_i$

$$\text{ATE}\left(H_i\right) = \beta_{H,0} + \beta_H H_i$$

The coefficient $\beta_{H,0}$ is the estimated average treatment effect if $H_i = 0$ and $\beta_H = \frac{\partial \text{ATE}(H_i)}{\partial H_i}$ is the sensitivity of the average treatment effect to a rise in $H_i$. Specifications with quadratic ATEs are also considered.

For example, theory predicts that a rise in demand should have a smaller impact on house prices in elastically supplied locations where it is easier to build real estate (Figure 4). This corresponds to the hypothesis $\beta_{elasticity} < 0$. The coefficient $\beta_{elasticity,0}$ is the estimated impact of the law change on prices in a hypothetical location where the asset (housing) is in perfectly inelastic supply.

This paper investigates treatment effect heterogeneity in elasticity, *pre-law* house prices, population, income and employment. Pre-law variables are set equal to their average value before 1998 to ensure they are unaffected by the treatment[14].

[14] The analysis was also conducted setting pre-law variables equal to their value in the first year of the sample. The estimates are identical.

## 4.1 Identification

The identifying assumption is parallel trends; i.e., that price changes in Texas would have been the same as in the control group, on average, over the treatment period if the law had not been passed. Letting $D$ be the treatment indicator, the identifying assumption can be written:

$$\underbrace{\mathbb{E}\left[y_{\text{TX, post}} - y_{\text{TX, pre}} | D = 0, X\right]}_{\text{unobserved: Texas (TX) if not treated}} = \underbrace{\mathbb{E}\left[y_{\text{control, post}} - y_{\text{control, pre}} | D = 0, X\right]}_{\text{observed: control group}} \tag{3}$$

If the law change was correlated with an expected rise in house prices, then the left side of 3 would be larger than the right side and the identification breaks down.

Abdallah and Lastrapes, 2012 linked these law changes to three major factors. First, the Tax Reform Act of 1986 which made mortgage interest the only form of interest on consumer credit which is tax deductible. This tax shield made HELs more attractive than other forms of debt. Second, a Fifth Circuit ruling in 1994 that federal regulations superseded the Texas constitution, temporarily overturned Texas restrictions on HELs. Even though subsequent actions by congress quickly reestablished the restrictions, this ruling brought attention and publicity to the issue. A third factor linked to these law changes was growing Republican control in Texas culminating in their 1997 return to being the majority party in the state senate. Republicans opposed these borrowing restrictions (and government regulation in general).

The identifying assumption can be defended because these three factors are not clearly linked to Texas house prices and other outcome variables studied in this paper.

Under additional assumptions, the regression coefficients are tied to the model in Appendix E.6. If the law had no indirect impact on prices, the coefficients can be interpreted as a measure of the willingness to pay for this embedded real option:

$$\eta_{1998} = \left(\frac{PDV_{98}(CV)}{p_{97}}\right)_{\text{Texas}} = \frac{\mathbb{E}_{98}\left[\sum_{j=1}^{\infty}(1-\delta)^{j-1} M_{99,98+j} \times CV_{98+j}\right]}{p_{97}}$$

$\eta_{1998}$ is the imputed fraction of its house price that a household would pay for the real option to pledge its home as collateral in the future. For example, if $\hat{\eta}_{1998} = 1\%$, then the average household is willing to pay 1% of its house price for this option.

## 4.2 Data

Data used in this paper are summarized in Table 1. The main outcome variable used in this study is the log real house price index. Detailed Texas house price data is notoriously hard to find as Texas is a non-disclosure state. Household level datasets such as CoreLogic do not have good Texas data for the relevant time periods. This paper measures house prices using the recently available Federal Housing Finance Agency (FHFA) 5-digit zip code data, which contains 17,936 zip codes in the US, including 918 zip codes in Texas. The trade-off for this lower level of geographic aggregation (zip code) is a higher level of time aggregation (annual). Like the S&P/CoreLogic/Case-Shiller home price indices, the FHFA series corrects for the changing quality of houses being sold at any point in time by estimating price changes with repeat sales. For details about the construction of this new dataset see Bogin, Doerner and Larson, 2016.

The Zillow house value index (ZHVI) is not used in the main analysis because the data begins too late in April 1996. However, Zillow data is used in the heterogeneity analysis and as a robustness test. BLABLABLA

Data used for controls, heterogeneity analysis and other outcome variables come from several different sources. Supply elasticity data is available at the MSA level from Saiz, 2010. Employment data at the county level is from the BLS. Median house price data at the zip code level and rent data at the MSA level is from Zillow. Income data at the county level is from the BEA. Population, single family building permit and homeownership data are from the census. The population and permit data is at the county level, whereas the homeownership rate data is at the MSA level. One must be careful in merging the data since the same zip code can be located in more than one county. This paper assigns each zip code to the county with the maximum allocation factor. US oil price data is from the EIA. US interest rates are constructed as in Himmelberg, Mayer and Sinai, 2005 by correcting the 10 year Treasury bond rate for inflation with the Livingston Survey of inflation expectations. Nominal variables are deflated using the CPI for all urban consumers from the BLS as in Glaeser, Gottlieb and Gyourko, 2012.

# 5 Estimates

This section presents and discusses the estimates. Table 2 provides summary statistics comparing the treatment group (Texas) to the control group in the three geographically nested samples. Texas had lower mean real house price growth than the rest of the US (1.1% vs 2.42%), but also lower house price volatility (3.61% vs 5.19%). Texas had lower pre-law median house prices than the US, but higher median prices than the control zip codes near its border. Consistent with its reputation of being a large state with a lot of space, Texas MSAs have a higher average supply elasticity than MSAs in the control group. Texas MSAs had lower pre-law home ownership rates than the control group (56% vs 63%).

Figure 1 plots the demeaned annual percent change in real house prices using raw data in Texas, the US, and four states bordering Texas. The series are plotted between 1992 and 2004, 6 years before and after the treatment. The only time in this sample when Texas had greater real house price growth than either the full US or its border states was for a few years after the law change in 1998.

## 5.1 Impact of the Credit Expansion on Texas House Prices

This section presents the main results of this paper, namely the impact of the credit expansion on Texas house prices. Estimates are presented for all three geographically nested samples and the data is weighted by the inverse of the number of zip codes in each state. The baseline specification contains year and location dummies using the most local fixed effects possible (zip code, county or MSA). It also includes a state time trend. The results are robust to whether the trend is interacted with treatment, state or MSA dummies (or to whether the time trend is linear or quadratic). Estimates from the baseline specification will be presented first, followed by an exploration of robustness to the method used for estimating standard errors and to specification.

Estimates from static regressions across the three samples (Table 3: columns 1-3) show the legalization of HELs raised Texas house prices $3.5\% - 5\%$. Estimates from dynamic regressions (Table 3: column 4-6) show parallel pre-trends across samples before the law $\eta_k \approx 0$ and a positive treatment effect after $\eta_k > 0$. The pre-trends show no evidence of anticipation before the law was implemented. The effect was gradual in the US and border samples but instantaneous in the Texarkana sample. In the first year in the US sample, $\hat{\eta}_{1998} \approx 1.5\%$ . In the following years, $\hat{\eta}_k \approx 3.36\% - 5.5\%$. The coefficients and 95%

confidence intervals from the dynamic regressions for all three samples are plotted in Figure 2 which shows parallel pre-trends and a statistically significant effect after the law. In the border sample, the effect was positive but did not become statistically significant until 1999.

These results answer the question posed in the title: yes, collateral value does affect asset prices. If households are or fear they will be credit constrained, they should value assets that facilitate their future ability to borrow. Hence, these estimates can be interpreted as price-based evidence that households are credit constrained.

Estimates in Table 4 explore whether the treatment effect remains significant when standard errors are conventional, robust, or clustered at the 5-digit zip code, 3-digit zip code, county FIPS code, MSA or state level. All estimates are statistically significant, however estimates are omitted for the smaller samples where there are not enough clusters to estimate standard errors.

Estimates in Table 5 investigate whether the treatment effect in Table 3 is robust to the inclusion of controls for oil prices, interest rates and a Lagged Dependent Variable (LDV). The treatment effect is statistically significant and very similar to the effect estimated in Table 3 across all samples and specifications. Since the specifications in columns 4-6 include both an LDV and fixed effects, they might suffer from the Nickell, 1981 bias. To correct for this bias, these specifications are estimated using the Arellano and Bond, 1991 (AB) method with all available lags as instruments.

The LDV coefficient is statistically significant in the US and border samples but not in the Texarkana sample which is too small. The estimated LDV coefficients $.59\% - .98\%$ provide evidence that house prices are auto-correlated. These results are consistent with other estimates in the literature[15].

## 5.2 Treatment Effect heterogeneity

This section explores if and in what ways the effect differed across treated zip codes. Figure 3 presents histograms and summary statistics of the treatment effect for each treated zip code in the three geographically nested samples. These estimates are from regressions in the baseline specification (Table 3 columns 1-3) except the term Texas $\times$ Post is interacted with an indicator for each zip code. The figure shows that the average treatment effect in each sample is the same as the corresponding estimate in Table 3 and there is considerable

[15] (Case and Shiller, 1989; Chinco and Mayer, 2016; Guren, 2016)

heterogeneity in the effect across zip codes.

Table 6 presents estimates from triple-difference regressions to investigate treatment effect heterogeneity along five dimensions. Estimates using the Saiz, 2010 measure of supply elasticity[16] (Table 6: column 1) show that zip codes in more elastic locations saw a smaller rise in prices. This is consistent with predictions from a partial equilibrium model; a rise in housing demand should have a bigger impact on prices in locations where it is relatively hard to build (Figure 4). The rise in house prices was 1.01% lower per unit of elasticity. If housing supply was perfectly inelastic, the average treatment effect would be the intercept 6.63%. The estimated treatment effect is plotted against elasticity for each Texas MSA in Figure 5. The treatment effect varies considerably from 5.45% in the most inelastic MSA (Galveston) to 1.86% in the most elastic MSA (Sherman).

Table 6, column 2 looks at heterogeneity by pre-law log population. Zip codes in counties with higher populations saw a greater treatment effect: a 1% larger pre-law population corresponds to a 0.009% larger treatment effect. Table 6, column 3 looks at heterogeneity by pre-law log real income per capita. Zip codes in higher income counties saw larger treatment effects. A 1% higher pre-law real income per capita corresponds to a 0.11% larger treatment effect. Table 6, column 4 looks at heterogeneity by the pre-law unemployment rate. Zip codes in counties with higher unemployment rates saw smaller treatment effects. A 1% higher pre-law unemployment rate corresponds to a 0.7% smaller treatment effect.

Table 6, column 5 investigates heterogeneity by each zip code's pre-law log real median house price level. Zillow estimated median house prices are used because the level of the FHFA index is not very informative about median price levels. The model estimates a quadratic function of pre-law median prices to allow for non-linearity. The effect in a linear triple-difference model is positive but not statistically significant, providing evidence of nonlinearity.

Ex ante pricier zip codes saw a bigger treatment effect. Figure 6 presents the fitted values for the predicted treatment effect which ranges from 2.08% to 16.7%. This might be because households in locations with higher pre-law house price levels were more likely to have a disproportionate fraction of their wealth stuck in their illiquid homes. These households are popularly referred to as "house rich, but cash poor" or as "The Wealthy Hand to Mouth" (Kaplan, Violante and Weidner, 2016). These wealthier households might

[16]While this measure of elasticity is widely used in the literature as an instrumental variable for house prices (Mian and Sufi, 2011), not all authors agree it is ideal (Davidoff, 2016).

also be more likely to use their home equity for entrepreneurship, and thus have a stronger demand for HELs. The coefficient on Texas × Post is not interesting in this specification because there were no observations with zero house price levels.

These results might appear to be in contrast with Landvoigt, Piazzesi and Schneider, 2015 (LPS), but they are not inconsistent. While LPS, 2015 find that the credit expansion during the housing boom had a bigger impact on ex-ante lower priced homes in San Diego, this paper finds that the (exogenous) HEL expansion had a bigger impact on ex-ante higher priced homes in Texas. The credit expansion in LPS, 2015 increased access to borrowers seeking all loans secured by housing, including both purchase mortgages and HELs. The Texas laws did not affect access to purchase mortgages, but rather the ability of existing home owners to extract equity by borrowing after they were already home owners.

The triple-difference regressions provide evidence that households in zip codes with stronger economic conditions (with *ex ante* higher prices, income and employment) value the option to pledge their home as collateral more strongly. This suggests that households prefer to borrow against their home equity for reasons unrelated to bad economic shocks (e.g. education or entrepreneurship). In fact, a good economic shock can increase investment opportunities, raising demand for HELs by home owners looking to start a business. This might also be due to the fact that lenders have debt to income underwriting requirements which make it harder for unemployed home owners to be approved for HELs.

### 5.3 Channels

This section investigates the channels behind the treatment effect by studying the impact of the credit expansion on various other outcome variables. The treatment effect could have occurred through two different channels:

1. Direct Channel: the law caused a rise in demand for homeownership due to the new option allowing homes to be pledged as collateral.

2. Indirect Channel: the law affected other variables that affect house prices. For example, if the law increased consumption and investment enough to stimulate the local economy, this could have raised demand for housing, thus raising the price.

If variables known to affect house prices such as rent, population, income and employment were not affected by the law change, that would provide evidence that the direct channel drove the treatment effect.

Estimates presented in Table 7 show that the law did not lead to economically or statistically significant changes in real rents, population or real income per capita. Economic theory implies that the rent regressions are particularly informative as the price of housing should have been equal to the present discounted value of rents before the legalization of HELs $p_t = \sum_{j=1}^{\infty} \frac{Rent_{t+j}}{(1+r)^j}$. After the HEL credit expansion, the price should reflect both rents and the collateral value $p_t = \sum_{j=1}^{\infty} \frac{Rent_{t+j}+\mathbf{CV_{t+j}}}{(1+r)^j}$. The rent regression helps reassure us that the treatment effect is due to the demand for HELs and not indirect effects on rent. There was a slight rise in the unemployment rate, but this works against the indirect channel. Together, these estimates suggest that the law change did not have a significant impact on variables related to house prices, providing evidence that the treatment effect occurred mainly through the direct channel.

Various concerns regarding the channels behind the treatment effect are considered and addressed below:

1. Home Improvement Loans: if Texans used HELs to improve their homes, the rise in house prices might be due to the higher quality of the properties and not due to the demand for HELs.

   A: Home improvement loans were available before 1998. In fact, home improvement loans were the only option for equity extraction before proposition 8.

2. Piggyback Loans: HELs could have increased purchase mortgage debt capacity for households who used second liens ("piggy-back mortgages") to avoid mortgage insurance and obtain bigger loans (Lee, Mayer and Tracy, 2012).

   A: This was not a relevant factor in Texas because of the 80% LTV limit for all liens obtained after purchase.

3. Unobserved economic impact: the HEL legalization could have increased consumption, investment and entrepreneurship, stimulating stimulating house prices in unobserved ways.

   A: The local samples can help control for unobserved heterogeneity if houses near each other are more likely to be affected by the same variables. In particular, the border city Texarkana, which is at the intersection of Texas and Arkansas, can be viewed as one economy. So if the law change affected the economy on the Texas side of Texarkana, it should have also affected the economy on the Arkansas side.

While this section provides evidence that the effect on house prices was direct, it is impossible to be certain as the law change could have affected other unobserved variables related to house prices not captured in the border and Texarkana samples.

### 5.4 The Marginal Buyer

Finally, this section investigates whether the law changes affected home ownership in Texas. The law change created a new benefit for home ownership. Hence, if households value this option and if house prices were held constant, demand for owner occupied housing would be expected to rise. However, in equilibrium, the rise in demand should have raised house prices until the marginal buyer was indifferent between owning and renting. The same logic applies to potential owners deciding whether or not to live in Texas (or a nearby state).

Estimates presented in Table 7, columns 2, and 5-6 show the law did not lead to economically or statistically significant changes in population, home ownership or single family building permits. These estimates provide evidence that the rise in house prices was sufficient to offset the rise in demand for ownership, keeping the marginal buyer indifferent between renting and owning.

## 6 Conclusion

A large body of literature studies the impact of credit constraints on borrowing, consumption and investment. This paper finds there is also an impact on prices. Estimates using zip code data show the law change raised Texas house prices by 3.8%. If households fear they will be credit constrained, they should value assets that facilitate their future ability to borrow. Hence, the treatment effect estimated in this paper can be interpreted as price based evidence that households are credit constrained.

Prices rose more in locations with inelastic supply, higher pre-law house prices, population, income and employment. This offers evidence that wealthier households and households in locations with stronger economic conditions value the option to pledge their home as collateral more strongly. The law change did not affect variables known to be related to house prices such as rent, population and income. This indicates that the treatment effect was mainly driven by the direct channel. Moreover, the border and Texarkana samples, which help control for local unobserved heterogeneity, provide further evidence that the effect was direct. Finally, the law change did not affect Texas population, ownership or construction. This offers evidence that the rise in price was sufficient to keep the marginal home buyer indifferent between owning and renting.

There are several promising avenues for future work. It would be interesting to see how pledgability affects the price of other assets such as stocks and Treasury bonds. Since 1974, regulation T has set the minimum initial margin requirement for stocks to 50%. However, finding a good source of exogenous variation to identify the collateral value of stocks might prove difficult. It also would be interesting for a macroeconomic life cycle model to provide an estimate of how much a household would be willing to pay for the option to pledge its home as collateral in the future. These structural estimates could be compared with the results in this paper. Finally, loans secured by housing have two benefits: a lower interest rate and a higher debt capacity. It would be interesting to separately identify the fraction of the collateral value that is due to the interest rate savings compared to the debt capacity.

In conclusion, owner occupied housing comes with a valuable option to pledge the home as collateral in the future. Prices reflect this.

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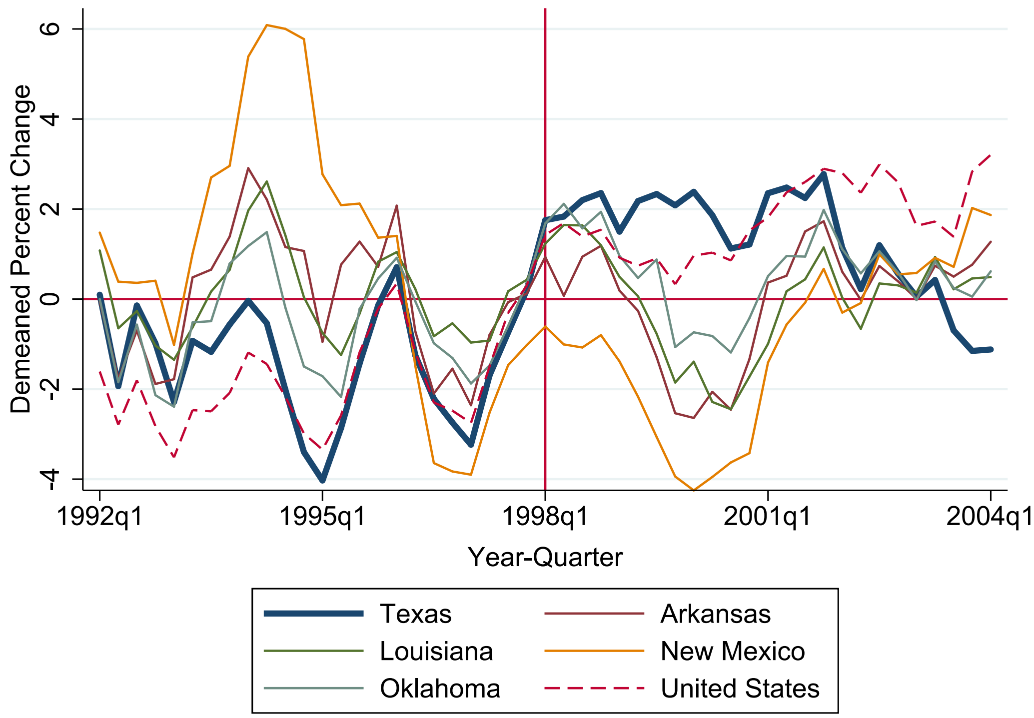
# A Appendix: Figures

## A.1 Percent Change in Real House Prices

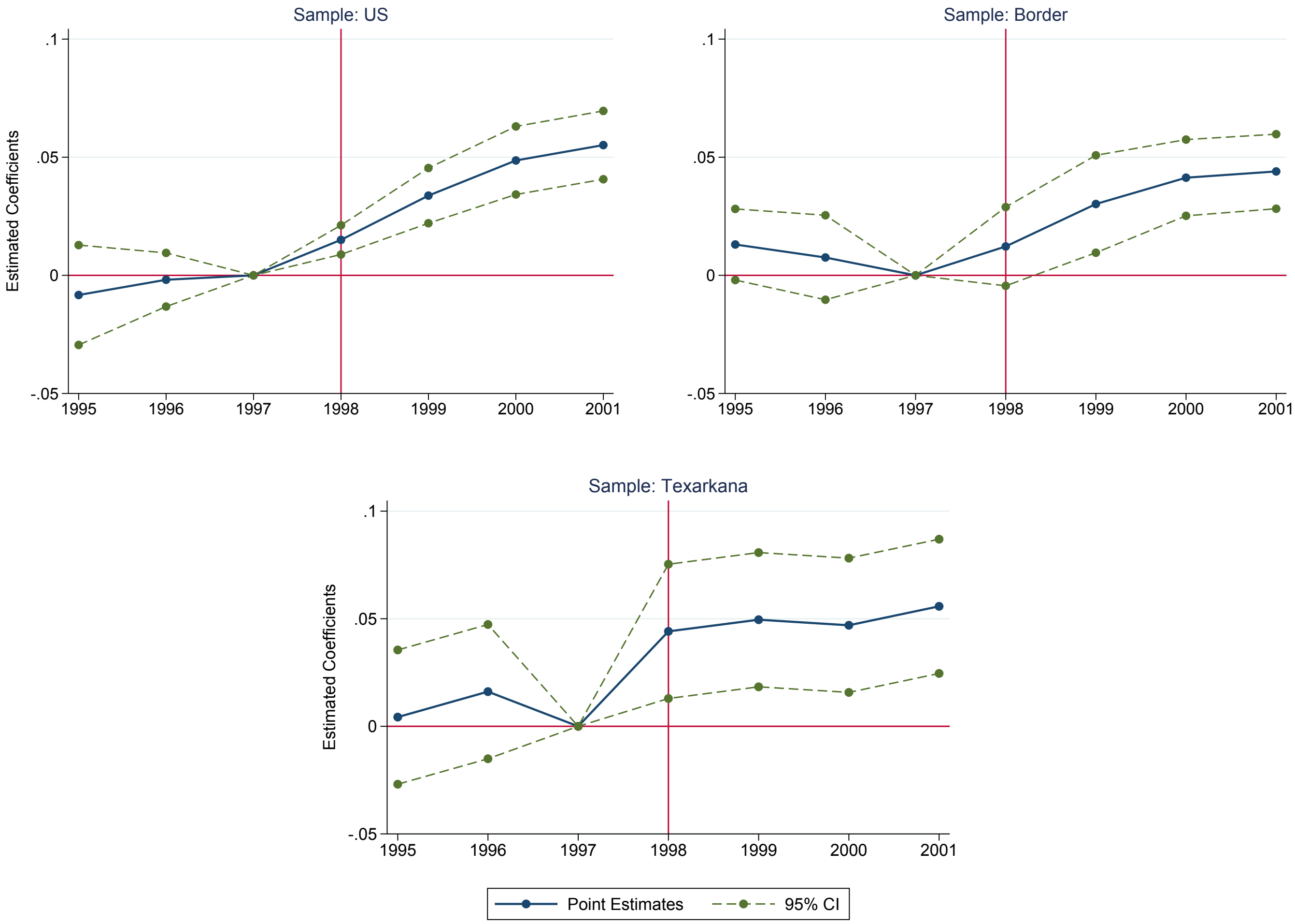
**Figure 1:** Annual Percent Change in Real House Prices (demeaned) in the US, Texas and Border States




*Note.* This figure plots the annual demeaned percent change in real house prices in the US, Texas and its four border states. The sample period is six years before and after the legalization of home equity loans in Texas in 1998. The house price data is from the FHFA AT Index. It is deflated by the CPI-U as explained in the paper. Data sources can be found in Table 1.

## A.2 Pre-trends

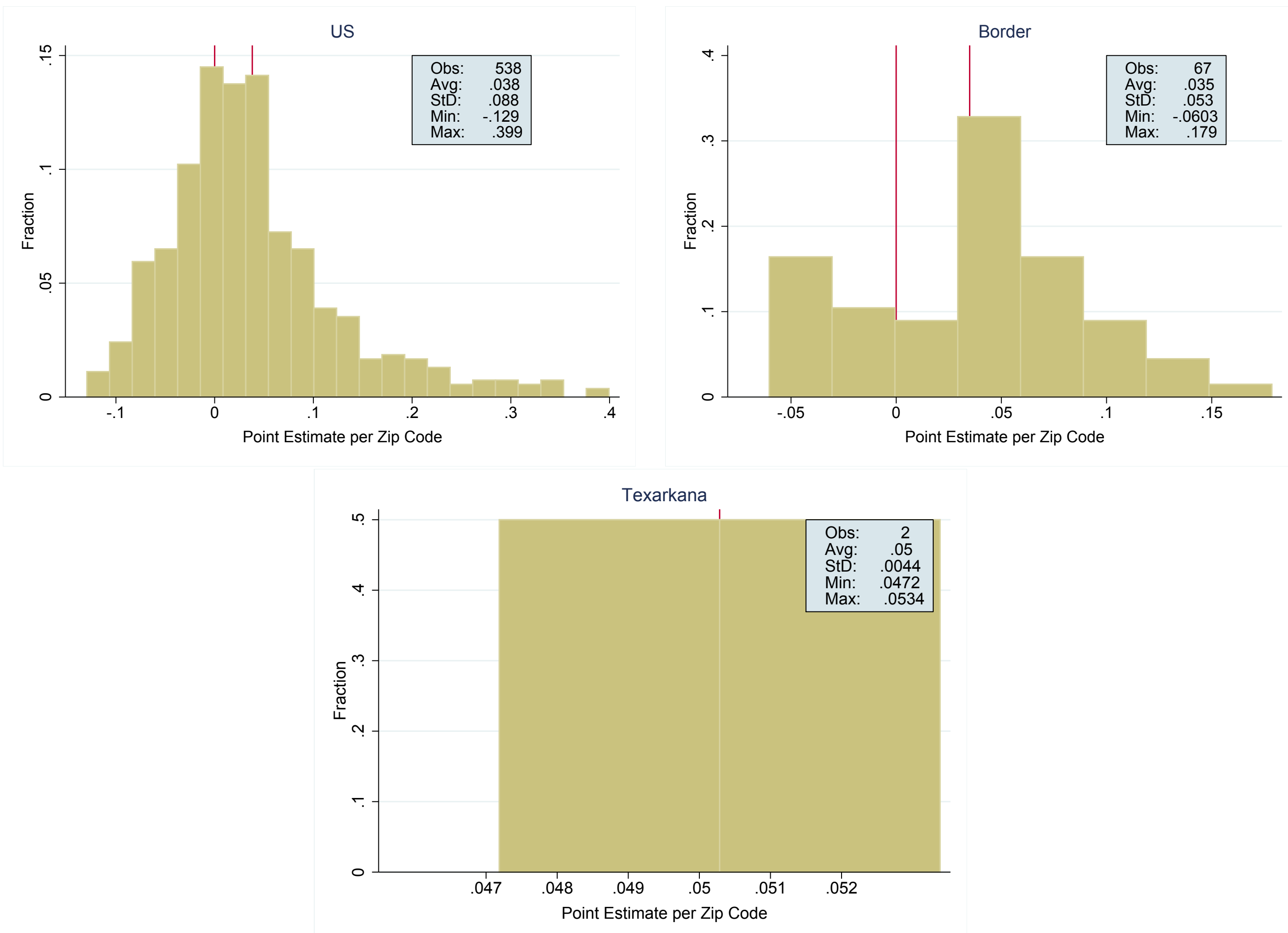
**Figure 2:** Impact of HEL Legalization on Texas House Prices in Three Geographically Nested Samples




*Note.* This figure plots point estimates $\hat{\eta}_k$ and 95% confidence intervals from the dynamic regression in Table 3: columns 4-6. There is a vertical red line in 1998, the year of the law change. Data sources can be found in Table 1.

## A.3 Treatment Effect Heterogeneity by Zip Code

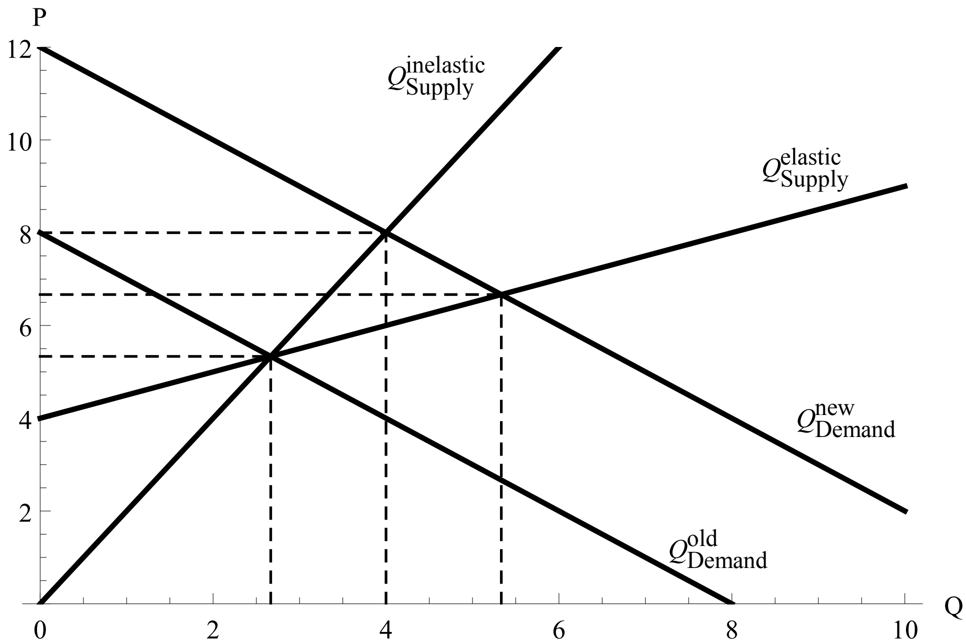
**Figure 3:** Heterogeneity in the Impact of HEL Legalization on Texas House Prices Across Zip Codes




*Note.* This figure presents histograms and summary statistics of the treatment effect for each zip code in three geographically nested samples. The mean treatment effect in each sample is exactly equal to the coefficient on Texas × Post in Table 3: columns 1-3. Data sources can be found in Table 1.

## A.4 Supply Elasticity Theory

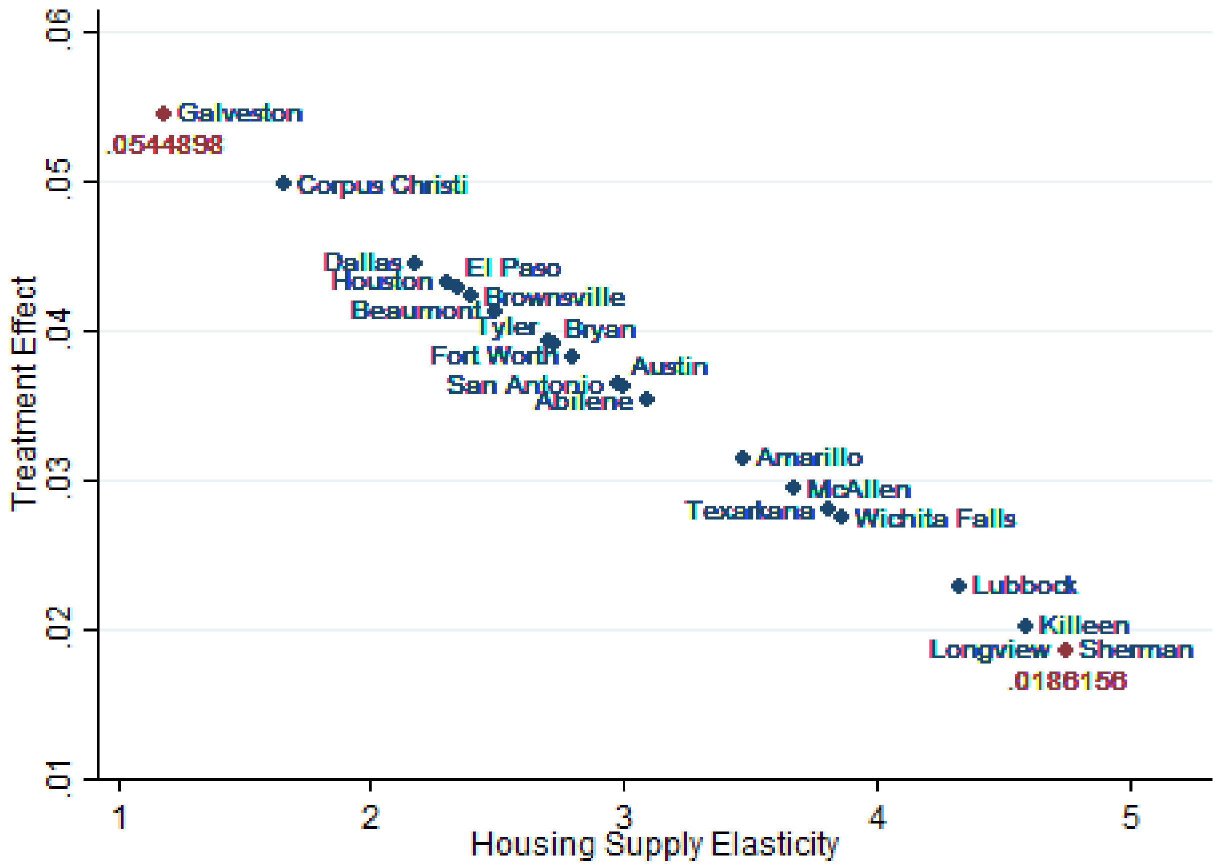
**Figure 4:** The impact of a rise in demand on house prices in cities with different supply elasticities




*Note.* This figure compares the impact of a rise in housing demand on house prices in two cities with different supply elasticities. Price (P) is on the y-axis and quantity (Q) in on the x-axis. Initially, the price of housing is the same in both cities. A rise in demand causes prices to rise more in the inelastic city relative to the elastic city.

## A.5 Elasticity Fitted Values

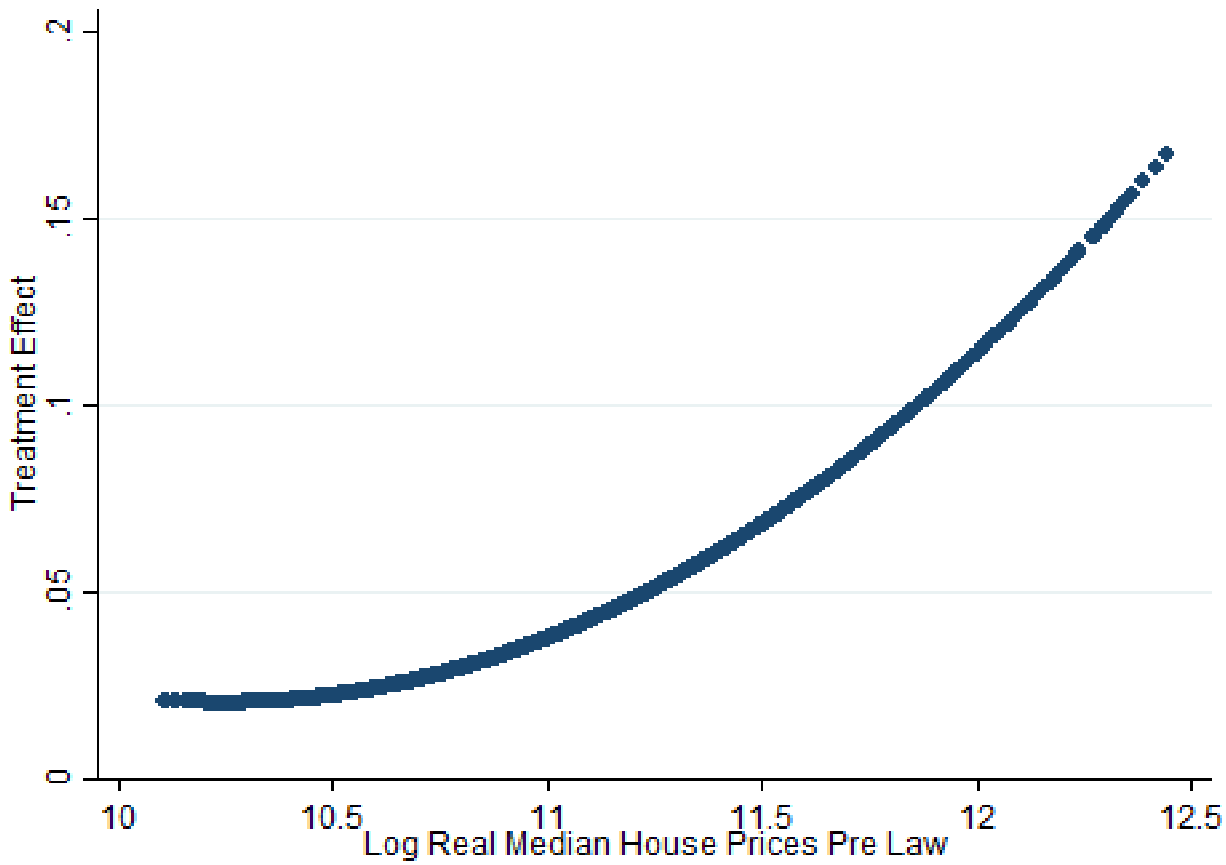
**Figure 5:** Impact of HEL Legalization on Texas House Prices by MSA Elasticity




*Note.* This figure plots fitted values for the treatment effect for each MSA against housing supply elasticities from triple-difference regressions presented in Table 6. This figure includes all 21 Texas MSAs. Galveston, the most inelastic MSA, had a 5.4% effect, whereas Sherman, the most elastic MSA, had a 1.86% effect. The measure of housing supply elasticity is from Saiz, 2010. Data sources can be found in Table 1.

## A.6 HP Heterogeneity Fitted Values

**Figure 6:** Impact of HEL Legalization on Texas House Prices by pre-law House Prices




*Note.* This figure plots fitted values for the treatment effect for each merged zip code against log pre-law real median house prices from triple-difference regressions presented in Table 6. Pre-law median price data is from Zillow. It is deflated by the CPI-U. Data sources can be found in Table 1.

# B Appendix: Tables

## B.1 Datasets

**Table 1:** DATASETS

| Variable | Level | Source | # Locations | TX |
|---|---|---|---|---|
| House Price Index | 5-digit zip code | FHFA | 17,936 zips | 918 zips |
| House Price Index | state | FHFA | 51: US & DC | 1 |
| Median House Prices | 5-digit zip code | Zillow | 12.665 zips | 576 zips |
| Median Rent (historic) | MSA | Zillow | 313 MSAs | 21 MSAs |
| Supply Elasticity | MSA | Saiz, 2010 | 269 MSAs | 21 MSAs |
| Employment | county | BLS | 2,581 fips | 153 fips |
| Income | county | BEA | 2,581 fips | 153 fips |
| Population | county | Census | 2,581 fips | 153 fips |
| Building permits | county | Census | 3,072 fips | 229 fips |
| Homeownership Rate | MSA | Census | 75 MSAs | 5 MSAs |
| Oil Prices | US | EIA | - | - |
| 10 year Treasury | US | Treasury | - | - |
| Inflation Expectations | US | Livingston survey | - | - |
| CPI-U | US | BLS | - | - |

*Note.* This table lists sources for the different variables used in this paper. All data are annual except for the FHFA state house price index, Zillow rent index, and the Zillow median price index.

## B.2 Summary Statistics

**Table 2:** Summary Statistics

| Sample: | US | | Border | | Texarkana | |
|---|---|---|---|---|---|---|
| Variable | Texas | Control | Texas | Control | Texas | Control |
| Real House Price Growth | 1.10 | 2.42 | 0.51 | 1.18 | 0.66 | 0.52 |
| | (3.61) | (5.19) | (3.22) | (2.95) | (1.93) | (2.14) |
| Real Median House Price | 107,517.60 | 133,946.00 | 95,822.70 | 80,018.36 | 76,367.96 | |
| | (48,058.05) | (72,570.51) | (36,849.00) | (29,386.70) | (15,377.16) | |
| Home Ownership Rate | 56.00 | 63.00 | 56.00 | | | |
| | (3.60) | (9.00) | (3.90) | | | |
| Real Rent | 483.00 | 482.00 | 495.00 | 386.00 | | |
| | (34.20) | (115.00) | (31.00) | (56.80) | | |
| Elasticity | 3.06 | 2.49 | 3.40 | 3.07 | 3.82 | 3.82 |
| | (1.01) | (1.46) | (1.21) | (1.06) | | |
| Single Family Permits | 798.46 | 575.71 | 651.08 | 191.54 | 128.67 | 87.83 |
| | (1,785.26) | (1,214.28) | (1,450.20) | (257.73) | (8.43) | (24.43) |
| Real Income Per Capita | 12,586.89 | 13,616.35 | 12,752.73 | 11,562.53 | 12,382.60 | 10,922.95 |
| | (2,333.05) | (2,826.71) | (2,280.64) | (1,231.80) | (252.44) | (546.46) |
| Unemployment Rate | 16.00 | 11.00 | 7.40 | 6.00 | 8.90 | 6.70 |
| | (4.70) | (7.10) | (3.10) | (2.10) | (0.78) | (1.00) |

*Note*. This table presents summary statistics for the treatment group (Texas) and control group (all locations outside Texas) in the three geographically nested samples used in the analysis. The standard deviation of each variable is in parentheses below the average. All nominal variables are adjusted for inflation as explained in Section 4. All variables except real house price growth and elasticity use pre-law average values as discussed in the paper. Both the Texas and Arkansas sides of the border city Texarkana have the same elasticity as estimated by Saiz, 2010. Median price, ownership and rental data are not available in some samples. Data sources can be found in Table 1.

## B.3 Main

**Table 3:** Impact of HEL Legalization on Texas House Prices

| VARIABLES | (1) US | (2) Border | (3) Texarkana | (4) US | (5) Border | (6) Texarkana |
|---|---|---|---|---|---|---|
| Texas × Post | 0.038*** | 0.035*** | 0.050*** | | | |
| | (0.007) | (0.009) | (0.012) | | | |
| Texas × 1995 | | | | -0.008 | 0.013* | 0.004 |
| | | | | (0.011) | (0.008) | (0.014) |
| Texas × 1996 | | | | -0.002 | 0.008 | 0.016 |
| | | | | (0.006) | (0.009) | (0.014) |
| Texas × 1998 | | | | 0.015*** | 0.012 | 0.044*** |
| | | | | (0.003) | (0.008) | (0.014) |
| Texas × 1999 | | | | 0.034*** | 0.030*** | 0.050*** |
| | | | | (0.006) | (0.010) | (0.014) |
| Texas × 2000 | | | | 0.049*** | 0.041*** | 0.047*** |
| | | | | (0.007) | (0.008) | (0.014) |
| Texas × 2001 | | | | 0.055*** | 0.044*** | 0.056*** |
| | | | | (0.007) | (0.008) | (0.014) |
| Observations | 147,069 | 1,430 | 39 | 147,069 | 1,430 | 39 |
| R-squared | 0.613 | 0.424 | 0.922 | 0.614 | 0.439 | 0.960 |
| std-err | state | zip5 | conventional | state | zip5 | conventional |

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

*Note.* This table reports estimates of the effect of a law change (which legalized HELs in Texas) on house prices in three geographically nested samples. Each column reports a separate regression estimated at the zip code year level where the dependent variable is the log of the real house price index. In columns 1-3, coefficients are reported for the interaction of the Texas dummy with an indicator for whether the year of observation falls on or after 1998. In columns 4-6, coefficients are reported for interactions of the Texas dummy with year indicators. All specifications include zip code and year fixed effects, along with state time trends. Standard errors are are reported in parentheses. Data sources can be found in Table 1.

## B.4 Standard Error Robustness

**Table 4:** Impact of HEL Legalization on Texas House Prices, Standard Error Robustness

| Method | US | Border | Texarkana |
|---|---|---|---|
| conventional | 0.038<br>(0.005)*** | 0.035<br>(0.009)*** | 0.050<br>(0.012)*** |
| robust | 0.038<br>(0.002)*** | 0.035<br>(0.009)*** | 0.050<br>(0.009)** |
| cluster zip5 | 0.038<br>(0.002)*** | 0.035<br>(0.009)*** | 0.050<br>(0.009)** |
| cluster zip3 | 0.038<br>(0.008)*** | 0.035<br>(0.013)** | |
| cluster FIPS | 0.038<br>(0.009)*** | 0.034<br>(0.013)** | |
| cluster MSA | 0.038<br>(0.010)*** | 0.035<br>(0.014)** | |
| cluster state | 0.038<br>(0.007)*** | | |

*** p<0.01, ** p<0.05, * p<0.1

*Note*. This table repeats the estimates in Table 3: columns 1-3, using a variety of different methods to estimate standard errors. Each column corresponds to the sample and each row corresponds to the method for estimating standard errors. Estimates of standard errors clustered at the state level are not not reported for the Border sample as there are too few clusters. Similarly estimates are not reported for the Texarkana sample for clusters more local than the 5-digit zip code level. Estimates clustered at the FIPS and MSA levels used slightly different samples as a few zip codes could not be matched. Data sources can be found in Table 1.

## Specification Robustness

**Table 5:** Impact of HEL Legalization on Texas House Prices, Specification Robustness

**Panel A: US Sample**

| VARIABLES | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| Texas × Post | 0.038*** | 0.036*** | 0.036*** | 0.040*** | 0.038*** | 0.039*** |
| | (0.007) | (0.008) | (0.007) | (0.003) | (0.004) | (0.004) |
| LDV | | | | 0.978*** | 0.945*** | 0.884*** |
| | | | | (0.035) | (0.037) | (0.042) |
| Observations | 147,069 | 147,069 | 147,069 | 134,070 | 134,070 | 134,070 |
| R-squared | 0.613 | 0.685 | 0.735 | | | |
| Covariates | N | Oil | Oil & Interest Rates | N | Oil | Oil & Interest Rates |

**Panel B: Border Sample**

| VARIABLES | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| Texas × Post | 0.035*** | 0.048*** | 0.047*** | 0.037*** | 0.039*** | 0.039*** |
| | (0.009) | (0.010) | (0.010) | (0.008) | (0.008) | (0.008) |
| LDV | | | | 0.602*** | 0.586*** | 0.602*** |
| | | | | (0.071) | (0.075) | (0.081) |
| Observations | 1,430 | 1,430 | 1,430 | 1,302 | 1,302 | 1,302 |
| R-squared | 0.424 | 0.435 | 0.439 | | | |
| Covariates | N | Oil | Oil & Interest Rates | N | Oil | Oil & Interest Rates |

**Panel C: Texarkana Sample**

| VARIABLES | (1) | (2) | (3) | (4) | (5) | (6) |
|---|---|---|---|---|---|---|
| Texas × Post | 0.050*** | 0.052*** | 0.051*** | 0.053*** | 0.064*** | 0.063*** |
| | (0.012) | (0.013) | (0.013) | (0.013) | (0.017) | (0.018) |
| LDV | | | | 0.113 | 0.068 | 0.089 |
| | | | | (0.181) | (0.191) | (0.196) |
| Observations | 39 | 39 | 39 | 35 | 35 | 35 |
| R-squared | 0.922 | 0.922 | 0.927 | | | |
| Covariates | N | Oil | Oil & Interest Rates | N | Oil | Oil & Interest Rates |

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

*Note.* This table reports estimates of the effect of a law change (which legalized HELs in Texas) on house prices in three geographically nested samples. Each column reports a separate regression estimated at the zip code year level where the dependent variable is the log of the real house price index. Specifications in columns 2-3 and 5-6 control for US oil prices and interest rates interacted with state dummies as described in the bottom row. Specifications in columns 4-6 include a Lagged Dependent Variable (LDV). These models are estimated with the Arellano and Bond, 1991 (AB) estimator to correct for the Nickell, 1981 bias using all available lags as instruments. Data sources can be found in Table 1.

## B.5 Treatment Effect Heterogeneity

**Table 6:** Impact of HEL Legalization on Texas House Prices, Treatment Effect Heterogeneity

| VARIABLES | (1) | (2) | (3) | (4) | (5) |
|---|---|---|---|---|---|
| Texas × Post | 0.066*** | -0.090** | -1.013*** | 0.079*** | 3.218** |
| | (0.011) | (0.042) | (0.246) | (0.015) | (1.325) |
| Texas × Post × Elasticity | -0.010*** | | | | |
| | (0.004) | | | | |
| Texas × Post × LogPopulationPre | | 0.009** | | | |
| | | (0.004) | | | |
| Texas × Post × LogRealIncomePre | | | 0.110*** | | |
| | | | (0.026) | | |
| Texas × Post × UnemploymentRatePre | | | | -0.007*** | |
| | | | | (0.002) | |
| Texas × Post × LogRealMedianHousePricesPre | | | | | -0.624** |
| | | | | | (0.245) |
| Texas × Post × LogRealMedianHousePricesPre2 | | | | | 0.030** |
| | | | | | (0.011) |
| Observations | 111,878 | 142,961 | 142,194 | 143,793 | 99,216 |
| R-squared | 0.654 | 0.614 | 0.619 | 0.614 | 0.632 |
| Max $H_i$ | 4.749 | 14.93 | 9.949 | 19.95 | 12.44 |
| Min $H_i$ | 1.18 | 9.467 | 8.936 | 1.95 | 10.11 |
| TE Max $H_i$ | .0186 | .0479 | .0815 | -.056 | .167 |
| TE Min $H_i$ | .0545 | -.00251 | -.03 | .0656 | .0208 |
| std-err | state | state | state | state | state |

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

*Note*. This table reports estimates of the effect of a law change (which legalized HELs in Texas) on house prices. Each column reports a separate regression in which the treatment effect is allowed to vary based on five measures of heterogeneity: supply elasticity, pre-law log population, log real income per capita, the unemployment rate and log real median house prices. Pre-law variables are set equal to their average before 1998. The dependent variable is the log of the real house price index. The first row reports coefficients for the interaction of the Texas dummy with an indicator for whether the year of observation falls on or after 1998. This represents the treatment effect if the measure of heterogeneity is equal to zero. The second to seventh rows report the coefficient on the triple interaction between the Texas dummy, an indicator for the treatment period and one of five measures of heterogeneity (and its square for median prices). All specifications include zip code and year fixed effects, along with state time trends. Standard errors are clustered at the state level and are reported in parentheses. Data sources can be found in Table 1.

## B.6 Other Outcome Variables

**Table 7:** Impact of HEL Legalization on Other Outcome Variables

| VARIABLES | (1)<br>Log<br>Real Rent | (2)<br>Log<br>Population | (3)<br>Log<br>Real Income<br>Per Capita | (4)<br>Unemployment<br>Rate | (5)<br>Home<br>Ownership<br>Rate | (6)<br>Log<br>Single Family<br>Permits |
|---|---|---|---|---|---|---|
| Texas × Post | 0.003 | 0.000 | -0.001 | 0.003** | 0.003 | 0.048 |
| | (0.005) | (0.007) | (0.005) | (0.001) | (0.004) | (0.033) |
| Observations | 16,276 | 40,690 | 40,404 | 21,645 | 947 | 36,614 |
| R-squared | 0.303 | 0.254 | 0.634 | 0.391 | 0.442 | 0.063 |
| std-err | state | state | state | state | state | state |

Robust standard errors in parentheses
*** p<0.01, ** p<0.05, * p<0.1

*Note.* This table reports estimates of the effect of a law change (which legalized HELs in Texas) on six outcome variables: log real rent, log population, log real income per capita, the unemployment rate, the home ownership rate and log single family building permits. Coefficients are reported for the interaction of the Texas dummy with an indicator for whether the year of observation falls on or after 1998. All specifications include year fixed effects along with location fixed effects at the most local level possible. Standard errors are clustered at the state level and are reported in parentheses. Data sources can be found in Table 1.

## C For Online Publication: Comparison of collateral value models

The collateral constraint can be written in the following ways:

$$
\begin{aligned}
-a_{t+1} &\leq \kappa_t p_t h_t \\
-a_{t+1} &\leq \kappa_t \mathbb{E}_t[p_{t+1}] h_t \\
-a_{t+1} &\leq \kappa_t p_t h_{t+1} \\
-a_{t+1} &\leq \kappa_t \mathbb{E}_t[p_{t+1}] h_{t+1}
\end{aligned}
$$

Depending on whether lenders let you borrow against the (1) present value of the asset you currently own, (2) expected future value of the asset you currently own, (3) present value of the asset you bought for tomorrow, (4) expected future value of the asset you bought for tomorrow.

Households borrow against the present value of housing they have today.

$$
c_t + p_t h_{t+1} + a_{t+1} \leq y_t + p_t h_t (1 - \delta_t) + (1 + r_t) a_t \tag{DBC $\lambda_t$}
$$

$$
-a_{t+1} \leq \kappa_t p_t h_t \tag{CC $\mu_t$}
$$

$$
\underbrace{p_t}_{\text{price}} = \mathbb{E}_t \left[ \underbrace{\beta \frac{u_1(t+1)}{u_1(t)}}_{sdf} \times \left( \underbrace{\frac{u_2(t+1)}{u_1(t+1)}}_{\text{service flow}} + \underbrace{\frac{\mu(t+1)}{u_1(t+1)} \kappa_{t+1} p_{t+1}}_{\text{collateral value}} + \underbrace{(1 - \delta_{t+1}) p_{t+1}}_{\text{resale price}} \right) \right]
$$

Households borrow against the expected future value of new housing.

$$
c_t + p_t h_{t+1} + a_{t+1} \leq y_t + p_t h_t (1 - \delta_t) + (1 + r_t) a_t \tag{DBC $\lambda_t$}
$$

$$
-a_{t+1} \leq \kappa_t \mathbb{E}_t[p_{t+1}] h_{t+1} \tag{CC $\mu_t$}
$$

$$\underbrace{p_t}_{\text{price}} = \mathbb{E}_t \left[ \underbrace{\beta \frac{u_1(t+1)}{u_1(t)}}_{sdf} \times \left( \underbrace{\frac{u_2(t+1)}{u_1(t+1)}}_{\text{service flow}} + \underbrace{\frac{\mu_t \kappa_t}{\beta u_1(t+1)} p_{t+1}}_{\text{collateral value}} + \underbrace{(1-\delta_{t+1})p_{t+1}}_{\text{resale price}} \right) \right]$$

Bianchi, Boz and Mendoza, 2012

Households borrow against the present value of land they buy at $t$ for $t+1$.

$$q_t k_{t+1} + c_t + \frac{b_{t+1}}{R_t} \leq q_t k_t + b_t + \varepsilon_t Y(k_t) \qquad \text{(DBC } \lambda_t\text{)}$$

$$-\frac{b_{t+1}}{R_t} \leq \kappa_t q_t k_{t+1} \qquad \text{(CC } \mu_t\text{)}$$

$$\underbrace{q_t}_{\text{price}} = \underbrace{\frac{q_t \mu_t \kappa_t}{u'(t)}}_{\text{collateral value}} + \mathbb{E}_t^s \left[ \underbrace{\beta \frac{u'(t+1)}{u'(t)}}_{\text{sdf}} \times \left( \underbrace{\varepsilon_{t+1} Y_k(k_{t+1})}_{\text{service flow}} + \underbrace{q_{t+1}}_{\text{resale price}} \right) \right]$$

This equation is not in their paper, this is a rearrangement of their FOC to illustrate collateral value.

Iacoviello, 2005

Households borrow against the expected future value of new housing collateral purchased at time $t$.

$$c_t + q_t h_t + \frac{R_{t-1} b_{t-1}}{\pi_t} + w_t' L_t \leq \frac{Y_t}{X_t} + q_t h_{t-1} + b_t \qquad \text{(DBC)}$$

$$b_t \leq m \mathbb{E}_t \left[ \frac{q_{t+1} h_t \pi_{t+1}}{R_t} \right] \qquad \text{(CC } \lambda_t\text{)}$$

$$\underbrace{q_t}_{\text{price}} = \mathbb{E}_t \left[ \underbrace{\gamma \frac{u'(c_{t+1})}{u'(c_t)}}_{\text{sdf}} \times \left( \underbrace{\nu \frac{Y_{t+1}}{X_{t+1} h_t}}_{\text{service}} + \underbrace{q_{t+1}}_{\text{resale price}} \right) + \underbrace{\lambda_t m \pi_{t+1} q_{t+1}}_{\text{collateral value}} \right]$$

Kiyotaki and Moore, 1997
Farmers borrow against the land they buy at $t$ at next period's price.

$$x_t + q_t k_t + Rb_{t-1} \leq (a + c)k_{t-1} + q_t k_{t-1} + b_t \tag{DBC $\lambda_t$}$$

$$Rb_t \leq q_{t+1} k_t \tag{CC $\mu_t$}$$

Fostel and Geanakoplos, 2008
Equation (9)
The collateral value of asset $j$ in state $s$ to agent $i$ is the marginal benefit from being able to take out loans backed by asset $j$

$$CV^i_{s,j} = \left[\frac{1}{1+r_s} - \frac{1}{1+\omega^i_s}\frac{1}{1+r_s}\right]\phi^i_{s,j} = \frac{1}{1+r_s}\frac{\omega^i_s}{1+\omega^i_s}\phi^i_{s,j}$$

Where $\omega^i_s$ is the liquidity wedge and $\phi^i_{s,j}$ is the collateral capacity.

He, Wright and Zhu, 2015
What this paper calls "collateral value", they call "Liquidity Value".
Equation (5)

$$\underbrace{\psi_t}_{\text{price}} = \underbrace{\beta}_{\textit{sdf}} \times \left( \underbrace{U_2\left(x_{t+1}, h_{t+1}\right)}_{\text{service flow}} + \underbrace{\alpha D_1 \psi_{t+1} \lambda\left(y_{t+1}\right)}_{\text{collateral value}} + \underbrace{\psi_{t+1}}_{\text{resale price}} \right)$$

Their work is closely related to the current paper.
Like here housing is used as collateral for future non-housing consumption, not to buy housing.
When the liquidity value is positive, they find that house prices can display fascinating dynamics. Figure 2 in their paper shows that steady state house prices are hump shaped in the LTV ratio.
In section 6 they allow houses to depreciate and construction.
Equation (13)

$$\underbrace{\psi_t}_{\text{price}} = \underbrace{\beta}_{\textit{sdf}} \times \left( \underbrace{\Omega\left(h_{t+1}\right)}_{\text{service flow}} + \underbrace{\alpha D_1 (1-\delta) \psi_{t+1} \Lambda\left[(1-\delta)\psi_{t+1} h_{t+1}\right]}_{\text{collateral value}} + \underbrace{(1-\delta)\psi_{t+1}}_{\text{resale price}} \right)$$

# D For Online Publication: Texas Legal Time Line

This section provides a brief time line of relevant laws regarding home equity borrowing in Texas (Abdallah and Lastrapes, 2012; McKnight, 1983; Stolper, 2015; Texas Legislative Council, 2016).

- Article XVI, Section 50 of Texas Constitution of 1876 protects homesteads from foreclosure for failure to pay all debts except the purchase loan, property taxes or a mechanic's lien.

- This section has only been amended twice before 1997; the first amendment extended protections to single adults in 1973 and the second amendment of 1995 related to divorce proceedings.

- December 1994, Texas Senate Interim Committee on Home Equity Lending comes out in strong support of easing restrictions on lending with limits on home equity lending to protect consumers, and called for the amendment proposal to be put on the ballot for voters to decide. The proposal did not gather the two-thirds majority in the House of Representatives, but it did pass the Senate, which was a first for such a proposal. The committees report was incorporated into House Joint Resolution 31 which passed the Texas House and Senate in May 1997.

- Voters approved Proposition 8 on 11-4-1997 with almost 60% voting yes of the 1.17 million votes cast, begins **1-1-1998**, legalized Home Equity Loans without restrictions on how the money could be used, however the total value of all liens on the home **cannot exceed 80% of the fair market value**. Reverse mortgages and cashout refinance mortgages were also legalized.

- Several problems with prop 8 in the first year.
  HELs not allowed for those who live on more than one acre of land.
  Reverse mortgage rules made the loans ineligible for purchase by Fannie Mae.

- Proposition 2 on the November 2, 1999 ballot corrected the shortcomings.
  Article XVI, Section 50 was amended to clear up the issues with reverse mortgages.
  Article XVI, Section 51 was amended by increasing the acreage limit used to define urban households to 10 acres.

# E For Online Publication: Collateral Constraint (present value) derivation

$$\max\ \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t u(c_t, h_t)$$

$$\text{s.t.}$$

$$c_t + p_t h_{t+1} + a_{t+1} \leq y_t + p_t h_t(1-\delta_t) + (1+r_t)a_t \qquad \text{(DBC } \lambda_t\text{)}$$

$$-a_{t+1} \leq \kappa_t p_t h_t \qquad \text{(CC } \mu_t\text{)}$$

Savings: $a_{t+1}$ at time $t$. If $a_{t+1} > 0$, save, if $a_{t+1} < 0$, borrow.
Multiplier on the dynamic budget constraint: $\lambda(s^t)\beta^t\pi(s^t)$.
The multiplier on the collateral constraint is: $\mu(s^t)\beta^t\pi(s^t)$
Complementary slackness: $\mu(s^t)\beta^t\pi(s^t)\left[\kappa(s^t)p(s^t)h_t(s^{t-1}) + a_{t+1}(s^t)\right] = 0$

State variables in $s^t$: $y(s^t)$, $h_t(s^{t-1})$, $a_t(s^{t-1})$, $\delta(s^t)$, $\kappa(s^t)$.
Choice variables in $s^t$: $c(s^t)$, $h_{t+1}(s^t)$, $a_{t+1}(s^t)$.

$$\begin{aligned}
\mathcal{L} = \sum_{t=0}^{\infty} \sum_{s^t \in S^t} & \beta^t \pi(s^t) u(c(s^t), h(s^t)) \\
& + \lambda(s^t)\beta^t\pi(s^t)\left[y(s^t) + p(s^t)h_t(s^{t-1})(1-\delta(s^t)) + \left(1 + r_t(s^{t-1})\right)a_t(s^{t-1})\right] \\
& - \lambda(s^t)\beta^t\pi(s^t)\left[c(s^t) + p(s^t)h_{t+1}(s^t) + a_{t+1}(s^t)\right] \\
& + \mu(s^t)\beta^t\pi(s^t)\left[\kappa(s^t)p(s^t)h_t(s^{t-1}) + a_{t+1}(s^t)\right]
\end{aligned}$$

In $s^t$ we have: $y(s^t)$, $h_t(s^{t-1})$, $a_t(s^{t-1})$, $\delta(s^t)$, $\kappa(s^t)$.
In $s^t$ we choose: $c(s^t)$, $h_{t+1}(s^t)$, $a_{t+1}(s^t)$. (Any two will pin down the third.)

$$\begin{aligned}
\mathcal{L}_{c(s^t)} &= \beta^t\pi(s^t)u_1(s^t) + \lambda(s^t)\beta^t\pi(s^t)\left[-1\right] \\
&= 0 \\
&\Leftrightarrow \\
\lambda(s^t) &= u_1(s^t)
\end{aligned}$$

$$\begin{aligned}
\mathcal{L}_{h_{t+1}(s^t)} &= \sum_{s^{t+1}|s^t} \beta^{t+1}\pi(s^{t+1})u_2(s^{t+1}) \\
&+ \lambda(s^t)\beta^t\pi(s^t)\left[-p(s^t)\right] \\
&+ \sum_{s^{t+1}|s^t} \lambda(s^{t+1})\beta^{t+1}\pi(s^{t+1})\left[p(s^{t+1})(1-\delta(s^{t+1}))\right] \\
&+ \sum_{s^{t+1}|s^t} \mu(s^{t+1})\beta^{t+1}\pi(s^{t+1})\left[\kappa(s^{t+1})p(s^{t+1})\right] \\
&= 0
\end{aligned}$$

$$\begin{aligned}
\mathcal{L}_{a_{t+1}(s^t)} &= \lambda(s^t)\beta^t\pi(s^t)\left[-1\right] \\
&+ \mu(s^t)\beta^t\pi(s^t)\left[1\right] \\
&+ \sum_{s^{t+1}|s^t} \lambda(s^{t+1})\beta^{t+1}\pi(s^{t+1})\left(1+r_{t+1}(s^t)\right) \\
&= 0 \\
&\Leftrightarrow \\
\lambda(s^t)\beta^t\pi(s^t) &= \mu(s^t)\beta^t\pi(s^t) + \sum_{s^{t+1}|s^t} \lambda(s^{t+1})\beta^{t+1}\pi(s^{t+1})\left(1+r_{t+1}(s^t)\right) \\
&\Leftrightarrow \\
\lambda(s^t) &= \mu(s^t) + \beta\left(1+r_{t+1}(s^t)\right)\mathbb{E}_t[\lambda(s^{t+1})] \\
\text{recall } \lambda(s^t) &= u_1(s^t) \text{ (consumption foc)} \\
&\Leftrightarrow \\
u_1(s^t) &= \mu(s^t) + \beta\left(1+r_{t+1}(s^t)\right)\mathbb{E}_t[u_1(s^{t+1})]
\end{aligned}$$

This is the famous liquidity-constrained euler equation! If the collateral constraint is not binding ($\mu = 0$) or doesn't exist, then this collapses to the usual frictionless euler equation.

Using $\lambda(s^t) = u_1(s^t)$, we combine the consumption and housing FOCs.

$$\begin{aligned}
u_1(s^t)\beta^t\pi(s^t)\left[p(s^t)\right] = &\sum_{s^{t+1}|s^t} \beta^{t+1}\pi(s^{t+1})u_2(s^{t+1}) \\
&+ \sum_{s^{t+1}|s^t} u_1(s^{t+1})\beta^{t+1}\pi(s^{t+1})\left[p(s^{t+1})(1-\delta(s^{t+1}))\right] \\
&+ \sum_{s^{t+1}|s^t} \mu(s^{t+1})\beta^{t+1}\pi(s^{t+1})\left[p(s^{t+1})\kappa(s^{t+1})\right]
\end{aligned}$$

$$\begin{aligned}
u_1(s^t)\left[p(s^t)\right] &= \mathbb{E}_t\beta u_2(s^{t+1}) \\
&+ \mathbb{E}_t u_1(s^{t+1})\beta\left[p(s^{t+1})(1-\delta(s^{t+1}))\right] \\
&+ \mathbb{E}_t \mu(s^{t+1})\beta\left[p(s^{t+1})\kappa(s^{t+1})\right]
\end{aligned}$$

$$\begin{aligned}
p(s^t) &= \mathbb{E}_t\beta\frac{u_2(s^{t+1})}{u_1(s^t)} \\
&+ \mathbb{E}_t\beta\frac{u_1(s^{t+1})}{u_1(s^t)}\left[p(s^{t+1})(1-\delta(s^{t+1}))\right] \\
&+ \mathbb{E}_t\beta\frac{\mu(s^{t+1})}{u_1(s^t)}\left[p(s^{t+1})\kappa(s^{t+1})\right]
\end{aligned}$$

$$\begin{aligned}
p(s^t) &= \mathbb{E}_t\beta\frac{u_1(s^{t+1})}{u_1(s^t)}\frac{u_2(s^{t+1})}{u_1(s^{t+1})} \\
&+ \mathbb{E}_t\beta\frac{u_1(s^{t+1})}{u_1(s^t)}\left[p(s^{t+1})(1-\delta(s^{t+1}))\right] + \mathbb{E}_t\beta\frac{u_1(s^{t+1})}{u_1(s^t)}\frac{\mu(s^{t+1})}{u_1(s^{t+1})}\left[p(s^{t+1})\kappa(s^{t+1})\right]
\end{aligned}$$

$$p(s^t) = \mathbb{E}_t\left[\beta\frac{u_1(s^{t+1})}{u_1(s^t)}\left(\frac{u_2(s^{t+1})}{u_1(s^{t+1})} + \left[p(s^{t+1})(1-\delta(s^{t+1}))\right] + \frac{\mu(s^{t+1})}{u_1(s^{t+1})}\left[p(s^{t+1})\kappa(s^{t+1})\right]\right)\right]$$

$$\underbrace{p_t}_{\text{price}} = \mathbb{E}_t\left[\underbrace{\beta\frac{u_1(t+1)}{u_1(t)}}_{sdf} \times \left(\underbrace{\frac{u_2(t+1)}{u_1(t+1)}}_{\text{service flow}} + \underbrace{\frac{\mu(t+1)}{u_1(t+1)}\kappa_{t+1}p_{t+1}}_{\text{collateral value}} + \underbrace{(1-\delta_{t+1})p_{t+1}}_{\text{resale price}}\right)\right]$$

$$p_t = \mathbb{E}_t[M_{t+1} \times (s_{t+1} + CV_{t+1} + (1-\delta_{t+1})p_{t+1})]$$

We can use the equilibrium value of the multiplier:
$\mu(t+1) = u_1(t+1) - \beta\,(1 + r_{t+2}(s^{t+1}))\,\mathbb{E}_{t+1}[u_1(t+2)]$

If there is no collateral constraint or the collateral constraint never binds ($\mu(s^t) = 0$), then housing has no collateral value ($CV_t = 0$).

If housing doesn't enter the utility function ($u_2 = 0$), then housing has no service flow value $s_t = 0$. In this case housing is only useful for it's collateral value.

## E.1 Price Decomposition: multiplicative CV

Recall $CV_{t+1} = \frac{\mu(t+1)\kappa_{t+1}}{u_1(t+1)} p_{t+1}$
Rearrange the pricing equation:

$$\begin{aligned}
p_t &= \mathbb{E}_t[M_{t+1} \times (s_{t+1} + CV_{t+1} + (1-\delta_{t+1})p_{t+1})] \\
&= \mathbb{E}_t\left[M_{t+1} \times \left(s_{t+1} + \frac{\mu(t+1)\kappa_{t+1}}{u_1(t+1)} p_{t+1} + (1-\delta_{t+1})\, p_{t+1}\right)\right] \\
&= \mathbb{E}_t\left[M_{t+1} \times \left(s_{t+1} + \left(1 - \delta_{t+1} + \frac{\mu(t+1)\kappa_{t+1}}{u_1(t+1)}\right) p_{t+1}\right)\right] \\
&= \mathbb{E}_t\left[M_{t+1} \times (s_{t+1} + A_{t+1} p_{t+1})\right]
\end{aligned}$$

Observe that if $CV_{t+1} = 0$, then $A_{t+1} = (1-\delta_{t+1})$, however if $CV_{t+1} > 0$, then $A_{t+1} > (1-\delta_{t+1})$.
Define $M_{t,j} \equiv M_t \times M_{t+1} \cdots M_{j-1} \times M_j$
similarly: $A_{t,j} \equiv A_t \times A_{t+1} \cdots A_{j-1} \times A_j$

$$\begin{aligned}
p_t &= \mathbb{E}_t\left[M_{t+1} \times (d_{t+1} + A_{t+1} p_{t+1})\right] \\
&= \mathbb{E}_t\left[M_{t+1} \times (d_{t+1})\right] + \mathbb{E}_t\left[M_{t+1} A_{t+1} p_{t+1}\right] \\
&= \mathbb{E}_t\left[M_{t+1} \times (d_{t+1})\right] + \mathbb{E}_t\left[M_{t+1} A_{t+1} \mathbb{E}_{t+1}\left[M_{t+2} \times (d_{t+2} + A_{t+2} p_{t+2})\right]\right] \\
&= \mathbb{E}_t\left[M_{t+1} \times (d_{t+1})\right] + \mathbb{E}_t\left[M_{t+1} M_{t+2}\left(A_{t+1} d_{t+2} + A_{t+1,t+2} p_{t+2}\right)\right] \\
&= \mathbb{E}_t\left[M_{t+1} \times (d_{t+1})\right] + \mathbb{E}_t\left[M_{t+1,t+2}\left(A_{t+1} d_{t+2}\right)\right] + \mathbb{E}_t\left[M_{t+1,t+2} A_{t+1,t+2} p_{t+2}\right] \\
&= M_{t+1} d_{t+1} + \mathbb{E}_t\left[\sum_{j=1}^{\infty} M_{t+1,t+1+j} \times A_{t+1,t+j} d_{t+1+j}\right] \\
&= \mathbb{E}_t\left[\sum_{j=1}^{\infty} M_{t+1,t+j} A_{t+1,t+j-1} d_{t+j}\right] \qquad A_{t+1,t} \equiv 1
\end{aligned}$$

Written multiplicatively in this subsection, the equation above shows how collateral value amplifies the service flow from housing. For ease of exposition, the additive case is given below.

## E.2 Price Decomposition: additive CV

Define $M_{t,j} \equiv M_t \times M_{t+1} \cdots M_{j-1} \times M_j$
Define the total dividend $d_t \equiv s_t + CV_t$

$$
\begin{aligned}
p_t &= \mathbb{E}_t \left[ M_{t+1} \times (d_{t+1} + (1-\delta)p_{t+1}) \right] \\
&= \mathbb{E}_t \left[ M_{t+1} \times (d_{t+1}) \right] + (1-\delta)\mathbb{E}_t \left[ M_{t+1} \times p_{t+1} \right] \\
&= \mathbb{E}_t \left[ M_{t+1} \times (d_{t+1}) \right] + (1-\delta)\mathbb{E}_t \left[ M_{t+1} \times \mathbb{E}_{t+1} \left[ M_{t+2} \times (d_{t+2} + (1-\delta)p_{t+2}) \right] \right] \\
&= \mathbb{E}_t \left[ M_{t+1} \times (d_{t+1}) \right] + (1-\delta)\mathbb{E}_t \left[ M_{t+1} \times M_{t+2} \times (d_{t+2} + (1-\delta)p_{t+2}) \right] \\
&= \mathbb{E}_t \left[ M_{t+1} \times (d_{t+1}) \right] + (1-\delta)\mathbb{E}_t \left[ M_{t+1,t+2} \left( d_{t+2} \right) \right] + (1-\delta)^2 \mathbb{E}_t \left[ M_{t+1,t+2} p_{t+2} \right] \\
&= \mathbb{E}_t \left[ \sum_{j=1}^{\infty} M_{t+1,t+j} \times \left( d_{t+j} \right) (1-\delta)^{j-1} \right] \\
&= \mathbb{E}_t \left[ \sum_{j=1}^{\infty} M_{t+1,t+j} \times \left( s_{t+j} + CV_{t+j} \right) (1-\delta)^{j-1} \right] \\
&= \mathbb{E}_t \left[ \sum_{j=1}^{\infty} M_{t+1,t+j} \times \left( s_{t+j} (1-\delta)^{j-1} \right) \right] + \mathbb{E}_t \left[ \sum_{j=1}^{\infty} M_{t+1,t+j} \times \left( CV_{t+j} (1-\delta)^{j-1} \right) \right]
\end{aligned}
$$

Given a stochastic process $x \equiv \{x_{t+j}\}_{j=1}^{j=\infty}$ we can define the PDV-operator

$$
\begin{aligned}
PDV_t \left( x_{t+j} \right) &= \mathbb{E}_t \left[ M_{t+1,t+j} x_{t+j} \right] \\
PDV_t \left( x \right) &= \mathbb{E}_t \left[ \sum_{j=1}^{\infty} M_{t+1,t+j} \times \left( x_{t+j} \right) \right] \\
&= \sum_{j=1}^{\infty} PDV_t \left( x_{t+j} \right) \\
&= PDV_t \left( x_{t+1} \right) + PDV_t \left( \{x_{t+j}\}_{j=2}^{j=\infty} \right)
\end{aligned}
$$

$$
\begin{aligned}
p_t &= \mathbb{E}_t \left[ \sum_{j=1}^{\infty} M_{t+1,t+j} \times \left( s_{t+j} (1-\delta)^{j-1} \right) \right] + \mathbb{E}_t \left[ \sum_{j=1}^{\infty} M_{t+1,t+j} \times \left( CV_{t+j} (1-\delta)^{j-1} \right) \right] \\
\underbrace{p_t}_{\text{price}} &= \underbrace{PDV_t \left( \{s_{t+j}\}_{j=1}^{j=\infty} \right)}_{\text{pdv service flow}} + \underbrace{PDV_t \left( \{CV_{t+j}\}_{j=1}^{j=\infty} \right)}_{\text{pdv collateral value}}
\end{aligned}
$$

The process $s \equiv \{s_{t+j}(1-\delta)^{j-1}\}_{j=1}^{j=\infty}$ will be written $s \equiv \{s_{t+j}\}_{j=1}^{j=\infty}$ to save space.

## E.3 Price before the law

We can re-write the price before the law assuming no expectations of the law:

$$p_t^{NL} = PDV_t\left(s_{t+1}^{NL}\right) + PDV_t\left(\left\{s_{t+j}^{NL}\right\}_{j=2}^{j=\infty}\right)$$
$$p_{t+1}^{NL} = PDV_{t+1}\left(\left\{s_{t+j}^{NL}\right\}_{j=2}^{j=\infty}\right)$$

## E.4 Price change before the law

$$\begin{aligned} p_{t+1}^{NL} - p_t^{NL} &= PDV_{t+1}\left(\left\{s_{t+j}^{NL}\right\}_{j=2}^{j=\infty}\right) - \left(PDV_t\left(s_{t+1}^{NL}\right) + PDV_t\left(\left\{s_{t+j}^{NL}\right\}_{j=2}^{j=\infty}\right)\right) \\ &= \left(PDV_{t+1} - PDV_t\right)\left(\left\{s_{t+j}^{NL}\right\}_{j=2}^{j=\infty}\right) - PDV_t\left(s_{t+1}^{NL}\right) \\ &= \text{news about future cash flow} - \text{current period cash flow} \end{aligned}$$

## E.5 Price Change at the time of the law

The law is a surprise and occurs at $t+1$. Homeowners can borrow at $t+1$, but the price will only reflect CV starting at $t+2$.

$$\begin{aligned} p_{t+1}^{L} - p_t^{NL} &= PDV_{t+1}\left(\left\{CV_{t+j}\right\}_{j=2}^{j=\infty}\right) + PDV_{t+1}\left(\left\{s_{t+j}^{L}\right\}_{j=2}^{j=\infty}\right) \\ &- \left(PDV_t\left(s_{t+1}^{NL}\right) + PDV_t\left(\left\{s_{t+j}^{NL}\right\}_{j=2}^{j=\infty}\right)\right) \\ &= \text{PDV(CV)} + \text{news about future cash flow} - \text{current period cash flow} \end{aligned}$$

## E.6 Impact of law changes on house prices

Compare the price change if the law change occurred at $t+1$ versus if it didn't:

$$\begin{aligned} \Delta^{\text{Law}} p_{t+1} &\equiv p_{t+1}^{L} - p_t^{NL} \\ &= PDV_{t+1}\left(\left\{CV_{t+j}\right\}_{j=2}^{j=\infty}\right) + PDV_{t+1}\left(\left\{s_{t+j}^{L}\right\}_{j=2}^{j=\infty}\right) - PDV_t\left(\left\{s_{t+j}^{NL}\right\}_{j=1}^{j=\infty}\right) \\ \Delta^{\text{NoLaw}} p_{t+1} &\equiv p_{t+1}^{NL} - p_t^{NL} \\ &= \left(PDV_{t+1}\left(\left\{s_{t+j}^{NL}\right\}_{j=2}^{j=\infty}\right) - PDV_t\left(\left\{s_{t+j}^{NL}\right\}_{j=1}^{j=\infty}\right)\right) \\ \Delta^{\text{Law}} p_{t+1} - \Delta^{\text{NoLaw}} p_{t+1} &= p_{t+1}^{L} - p_{t+1}^{NL} \\ &= PDV_{t+1}\left(\left\{CV_{t+j}\right\}_{j=2}^{j=\infty}\right) + PDV_{t+1}\left(\left\{s_{t+j}^{L} - s_{t+j}^{NL}\right\}_{j=2}^{j=\infty}\right) \end{aligned}$$

If we assume the law had no impact on rents $s^L_{t+j} = s^{NL}_{t+j}$, then the second term cancels out

$$\Delta^{\text{Law}} p_{t+1} - \Delta^{\text{NoLaw}} p_{t+1} = PDV_{t+1}\left(\{CV_{t+j}\}_{j=2}^{j=\infty}\right)$$
$$\frac{\Delta^{\text{Law}} p_{t+1}}{p_t} - \frac{\Delta^{\text{NoLaw}} p_{t+1}}{p_t} = \frac{PDV_{t+1}\left(\{CV_{t+j}\}_{j=2}^{j=\infty}\right)}{p_t}$$

In the dynamic difference-in-differences regression the coefficient is

$$\begin{aligned}\eta_{1998} &= \frac{\Delta^{\text{Law}} p_{t+1}}{p_t} - \frac{\Delta^{\text{NoLaw}} p_{t+1}}{p_t} \\ &= \frac{PDV_{t+1}(CV)}{p_t} \\ &= \frac{\mathbb{E}_{t+1}\left[\sum_{j=1}^{\infty}(1-\delta)^{j-1} M_{t+1,j} \times CV_{t+1+j}\right]}{p_t}\end{aligned}$$

Hence, if we assume that the law change had no impact on rents then $\eta_{1998}$ is not only the impact of the law change on house prices, but also the percent of a house price due to the collateral option value.

If the law change increases the supply of housing and $h_{t+j}$ rises more than it would without the law, then rents should be lower $s^L_{t+j} \leq s^{NL}_{t+j}$ implying the coefficient underestimates the collateral option value

$$\begin{aligned}\eta_{1998} &= \frac{\Delta^{\text{Law}} p_{t+1}}{p_t} - \frac{\Delta^{\text{NoLaw}} p_{t+1}}{p_t} \\ &\leq \frac{PDV_{t+1}(CV)}{p_t} \\ &= \frac{\mathbb{E}_{t+1}\left[\sum_{j=1}^{\infty}(1-\delta)^{j-1} M_{t+1,j} \times CV_{t+1+j}\right]}{p_t}\end{aligned}$$

The more elastic housing supply is in a given location, the bigger the rise in $h_{t+j}$, the more we would be underestimating the collateral option value. However, regardless of the impact on rent (service flow), as long as the parallel trends assumption holds, we are still able to identify the total impact of the law on house prices.

In general

$$\eta_k = \mathbb{E}\left(y_{Texas,k} - y_{Control,k}\right) - \mathbb{E}\left(y_{Texas,97} - y_{Control,97}\right)$$
$$\eta_k = \frac{\Delta^{\text{treatment}} p_{t+k}}{p_t} - \frac{\Delta^{\text{control}} p_{t+k}}{p_t}$$