# Housing Deposit Channel of Monetary Policy and Housing Price Double-Dip

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**Abstract.** "Rising interest rates lead to a gradual decline in housing prices.", a widely held belief; this paper challenges this conventional wisdom both empirically and theoretically by showing that sustained restrictive monetary policy results in a double-dip in housing prices: an initial dip immediately upon impact, followed by an upswing, and then a subsequent second dip in the longer run. The findings identify a novel monetary transmission channel, i.e., the housing deposit channel, as the driver of this phenomenon, employing lag-augmented local projection techniques for estimation. In addition, a general equilibrium model provides theoretical insights into the underlying mechanisms that influence housing prices through the housing deposit channel. By shedding light on the transmission channels of interest rate hikes to the economy, our study outlines the effects of monetary policy on the housing market, and reciprocally, how the housing market influences the efficacy of monetary policy, particularly in achieving inflation targeting objectives.

**Keywords:** Monetary transmissions, deposit channel, housing market, lag-augmented local projection
**JEL Classification:** C32, E43, E52, G2

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# 1. Introduction

This paper addresses the following key questions: (i) How does monetary policy affect the housing market in the short, medium and in the longer run? and in turn, (ii) What is the impact of housing channels on the effectiveness of monetary policy? The existing literature suggests that a contractionary monetary policy decreases housing demand, which consequently reduces housing prices. For instance, a meta-analysis by Ehrenbergerova et al. (2023), assessing vector autoregression (VAR) estimates, observes that on average a 1 percentage point increase in interest rates causes a 0.7 percent median decrease in housing prices at the one-year horizon and 0.9 percent at the two-year horizon. Figure 1 replicates their results using a U.S.-specific sub-sample from this study[1], showing a significant decline in U.S. housing prices following an interest rate increase.

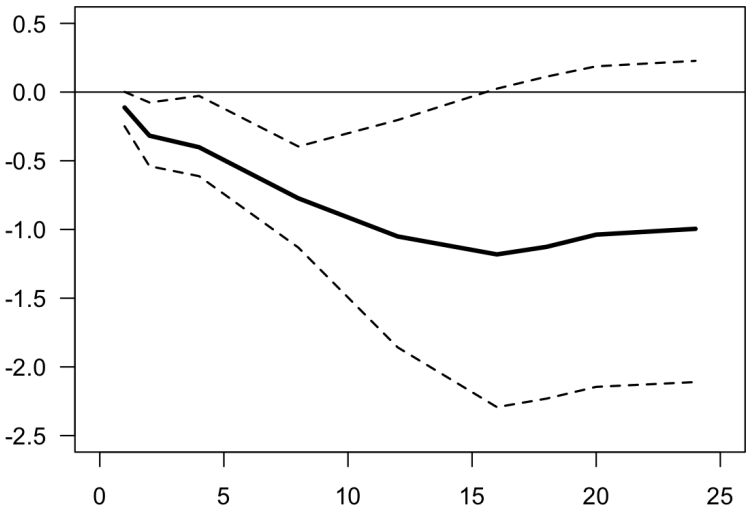



Figure 1: Mean response of US housing prices to interest rate shock. Solid line represents mean impulse response after 100bp increase in interest rates. The dashed lines indicate one standard deviation confidence intervals. Horizontal axis measures the number of quarters and vertical axis measures the percentage change.

Contrary to the existing literature, our paper indicates both empirically and theoretically that rising interest rates do not always lead to a persistent decline in housing demand and housing prices. Our paper is the first to unveil a "housing price double-dip" as a result of contractionary monetary policy; there is an initial decline in housing prices immediately after impact (which is shown in the previous literature, for instance Ehrenbergerova et al. (2023)), followed by a subsequent period of increase, and finally a second dip in the longer run during which housing prices decline again. Our paper brings these findings into line with a novel transmission channel in the literature, the "housing deposit channel of monetary policy". The three distinct phases above and the channel are explained as follows:

[1]Ehrenbergerova et al. (2023) collect 1,447 estimates of the effect of changes in short-term interest rates on housing prices from 31 VAR studies. In Figure 1, IRFs are collected manually, with the responses and their confidence intervals measured using pixel coordinates, following Rusnák et al. (2013). From the studies mentioned in Ehrenbergerova et al. (2023) that focus solely on the US housing market, we collect the responses of housing prices to a change in the short-term interest rate after one, two, four, eight, twelve, and sixteen quarters, as well as the longest time period. For each observation, the change in the interest rate is standardized to an increase of 100 basis points. The confidence bands are set at ± 1SE. The nine studies used in this analysis are the following: Assenmacher-Wesche & Gerlach (2008), Calza et al. (2007), Coibion et al. (2017), Gupta, Jurgilas, Kabundi & Miller (2012), Gupta, Jurgilas, Miller, Van Wyk et al. (2012), Jarociński & Smets (2008), McDonald & Stokes (2013), Musso et al. (2011), Sousa (2014).

First, our paper distinguishes between lender (patient) households which buy houses outright, and borrower (impatient) households, which borrow from banks to satisfy their housing demand. As shown by Mian & Sufi (2009) and Mian & Sufi (2011), housing price cycles are generally considered to arise from significant exogenous shocks, typically related to the effective demand for housing. An initial interest rate hike reduces the effective demand for housing as borrower households decrease their mortgage demand. In turn, this causes a temporary bust in prices immediately after impact. This first bust is observed in most of the existing literature and is also captured by the empirical and theoretical models in our paper.

Second, our paper goes further and observes that as interest rates continue to rise, lender household demand for housing and thus housing prices begin to rise as monetary policy is transmitted through the housing deposit channel in the following way: Drechsler et al. (2017) observe that households have a preference for liquidity, which they can obtain from cash and deposits. The Federal Funds Rate (FFR) is the cost of holding cash, while the difference between the FFR and the deposit rate (the deposit spread) is the cost of holding deposits. When the central bank raises the policy rate, as in the case of a contractionary monetary policy, cash becomes more expensive to hold. Conversely, having market power in their local deposit markets, banks tend to raise deposit spreads without losing deposits to cash. Lender households respond by reducing their deposit holdings at banks, and deposits flow out of the banking system.

These deposits are likely to flow into the housing market as housing represents the largest asset on household balance sheets in many developed economies (Piazzesi & Schneider 2016), while providing substantial utility. This pattern is illustrated in Figure 2, where lender households increasingly perceive the housing market as an investment opportunity following a rate hike, albeit with a lag. This channel, i.e., the housing deposit channel of monetary policy, finances lender households' housing demand in response to lower prices, ultimately contributing to the recovery of housing prices.

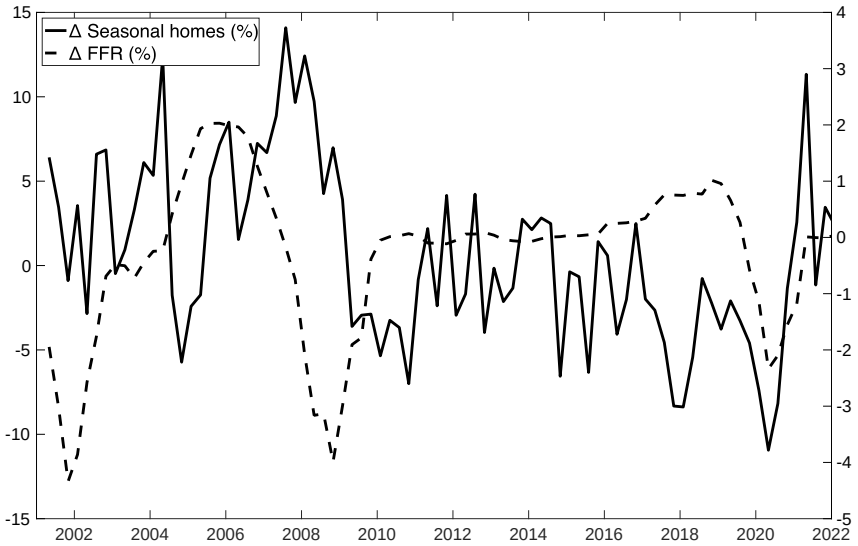



Figure 2: Seasonal homes and interest rate growth. Seasonal housing serves a proxy for second homes.[2]Left vertical axis indicates year-over-year growth of seasonal housing units in %. Right vertical axis indicates year-over-year growth of FFR in %. Data is provided by the US Census Bureau and the Federal Reserve System. A positive relationship between the FFR and the number of seasonal housing units is observed. This suggests that the demand for second homes as an investment opportunity for lender households increases with a lag after a rate hike.

Third, banks, experiencing deposit outflows, seek to attract more deposits to meet mortgage demand and strengthen their net worth. This marks the beginning of the second dip. The rise in deposit spreads prior to the recovery period, which led to large deposit outflows, adds to the already high liquidity needs of lender households, thereby driving up demand for bank deposits. As lender households start to prioritize bank deposits, their demand for housing in the markets declines, which subsequently exerts downward pressure on housing prices again, although not to the same degree as in the first dip.

The existing empirical literature is mostly unable to capture these dynamics in housing prices beyond the first dip, as the effects of monetary policy on housing prices are primarily analyzed by estimating structural VAR models (SVARs).[3] As a result, SVARs are often unable to capture transitory states and long-term fluctuations since they are inherently restricted in their dynamics and, by definition, intentionally over-optimistic with regard to the rate of decay of the impulse responses, assuming the VAR is stable (Jordà et al. 2023). This leaves the medium and longer-run behavior of housing prices mostly unexplored.

This paper fills this gap by employing lag-augmented local projection (LA-LP) techniques where additional lags of the variables are included in the regression as controls. This paper follows Montiel Olea & Plagborg-Møller (2021) who demonstrate that LA-LP are particularly well suited to drawing robust inference with regard to impulse responses at long horizons. Consequently, these techniques help to avoid the issue of overparameterization inherent in VARs,[4] while enabling an analysis of housing price responses in the longer run.

Moreover, standard confidence intervals based on LA-LP are considered to have correct asymptotic coverage uniformly over the persistence in the data generating process and over a wide range of horizons.[5] In addition to robustifying inference, LA-LP simplifies the computation of standard errors by avoiding the adjustment for serial correlation in the residuals. Instead of computing heteroscedasticity and autocorrelation consistent/robust (HAC/HAR) standard errors, as proposed in most of the LP literature (see Jordà 2005, Ramey 2016), this paper follows Montiel Olea & Plagborg-Møller (2021) who show that the classical heteroscedasticity-robust standard errors are also sufficient for lag-augmented local projections, while possibly having even better coverage/length properties.[6]

The empirical identification strategy in this paper is based on the observation that deposit spreads and short-term deposits are seen to respond immediately to changes in the policy rate, which helps to distinguish between shocks transmitted through a housing deposit channel and those that are not. To do so, following Romer & Romer (2004) and Coibion et al. (2017), this paper uses unanticipated interest rate surprises around narrow windows on Federal Open Market

[2]This is because data on second homes is not available for a frequent time period. The National Association of Home Builders (NAHB) bases its estimate of the number of second homes, i.e., non-rental property not classified as the taxpayer's principal residence, on, among other things, seasonal or vacation homes.

[3]Standard VAR models are used in the majority of the existing empirical work on monetary policy shocks and housing price dynamics, as they are widely considered to be standard for analyzing the propagation of shocks at business cycle frequencies.

[4]The issue of overparameterization occurs as the number of component series (variables or lags) is increased. This might bias the estimation results and compromise analysis derived from more distant impulse responses.

[5]This means that confidence intervals remain valid even if the data exhibits unit roots, and even at horizons $h$ that are allowed to grow with the sample size $T$.

[6]Montiel Olea & Plagborg-Møller (2021) provide simulations which show that lag-augmented local projections with heteroscedasticity-robust standard errors have better coverage/length properties than more standard non-augmented local projections with "off-the-shelf HAR standard errors."

Committee (FOMC) meeting days as a variable for a monetary policy shocks. Zero-maturity deposits are then checked for changes in the week following the rate surprises. Interest rate changes that affect the housing market through a deposit channel are identified using the "poor man's" sign-restriction approach of Jarociński & Karadi (2020): Only interest rate surprises that are accompanied by opposite-sign deposit changes in the week following the FOMC are considered. Implementing this shock in the LA-LP framework, the impulse responses show a two-dip dynamic of housing prices along the horizon while strongly indicating a housing deposit channel of monetary policy.

To confirm the soundness of our empirical findings, we conduct several robustness checks. We, first, distinguish endogenous fluctuations in the FFR from exogenous monetary policy shocks using Romer & Romer (2004). The resulting IRFs closely align with the baseline estimates. Second, incorporating alternative variable choices and housing price indexes, as in Calza et al. (2007) and Calza et al. (2013), reaffirms the double-dip pattern in housing prices following interest rate hikes. We then show that shocks triggering an initial decline in housing prices, similar to an interest rate hike, do not produce the same effect. For instance, credit risk and uncertainty shocks lead to a decline in housing prices but show no evidence of a second dip in the longer term.

To support the empirical findings and provide theoretical insights into the housing deposit channel, this paper develops a general equilibrium model based on Iacoviello (2015). The model includes a housing market, a financial sector, and two types of households: lenders and borrowers. By incorporating the interactions between borrower and lender households, as well as the financial sector's response to monetary policy shocks, the model effectively captures the housing price double dip, reflecting key dynamics observed in our empirical analysis.

The model indicates that there is a difference between the reactions of borrower and lender households following an interest rate increase. Borrowers' consumption and housing demands typically fall as interest rates rise due to the wealth effect. This leads to an initial drop in housing prices i.e., the first dip. The model outlines that the response of lenders is more complex, as it depends on how banks pass on interest rates to deposits. Therefore, lender households may shift funds from deposits to the housing market, activating the housing deposit channel, which leads to an increase in housing prices. Finally, the model outlines the mechanisms driving the second dip, as banks respond to deposit outflows by attempting to reverse the trend and attract more deposits to meet regulatory obligations. The model illustrates that housing prices are affected not only by the simple relationship between banks and borrowers through mortgages, but also by the complex dynamics between banks and lenders through both the deposit and housing deposit channels of monetary policy.

In addition, this paper offers new evidence to support the implementation of inflation-targeting monetary policy with broader perspective that includes both short-term and long-term impacts. While most inflation-targeting central banks have adopted forecast horizons of two or three years, we show that interest rate changes have lasting effects on housing prices over medium and long-term periods. This suggests that monetary policy should not focus solely on the immediate effects of housing prices on inflation and the real economy but should also account for their long-term dynamics.

Our findings challenge the effectiveness of a "leaning against the wind" policy (as proposed

by Borio & White 2004, Roubini 2006, Woodford 2012, Bank for International Settlements 2016, for instance), which involves actively using interest rates to curb financial imbalances in the housing market, for example. As our results show, while the housing prices recovers in the medium term after a contractionary monetary policy shock, there is a second, albeit smaller, dip in the longer run due to the prolonged impact of the initial shock. Policymakers should therefore not overreact to short-term fluctuations in housing inflation, as long-term considerations are crucial for formulating effective and stable monetary policy.[7]

**Related Literature** Our paper fuses various strands of the literature including housing, monetary policy and impulse response analysis at distant horizons. To the best of our knowledge, our paper is the first to empirically assess the response of housing prices to monetary policy over the medium to longer term using LA-LP methods. As a result, it is the first study to identify that housing prices fall twice, i.e., experiences a housing price double-dip, after a contractionary monetary policy. Moreover, the paper introduces a novel transmission channel of monetary policy, the housing deposit channel, shedding light on its potential significance for the effectiveness of monetary policy, particularly in terms of inflation targeting. Finally, the general equilibrium model presented in this paper is the first to incorporate the housing deposit channel, showing that when the central bank raises interest rates, the pass-through of these rates to depositors by banks affects consumption and housing demands, which may, in turn, lead to an increase in housing prices.

Several transmission channels of interest rate policy to the housing market and thus the economy are proposed in the literature (see Mishkin 2007, for instance).[8] More recently, Drechsler et al. (2017) have added an additional channel to the relevant literature: the deposit channel of monetary policy. Abadi et al. (2023), Choi & Rocheteau (2023) and Xiao (2020), among others, expand on the deposit channel to assess its implications for the efficacy of monetary policy. However, their analyses do not include the housing market, despite its significance as the largest asset on household balance sheets in developed countries. Our paper contributes to this literature by explicitly considering housing in the analysis, introducing the so-called housing deposit channel of monetary policy and thereby complementing the transmission channels mentioned above.

With regard to the empirical literature, most analyses of housing price dynamics and monetary policy rely on standard VAR techniques, (see Iacoviello & Minetti 2008, Goodhart & Hofmann 2008, Jarociński & Smets 2008, Calza et al. 2013, McDonald & Stokes 2013, for instance). As most of them do not explicitly account for the longer-run responses of housing prices, they observe the typical dip after interest rate hikes in the short run. However, as our paper examines longer-term dynamics, the use of VAR is problematic, as it is limited in that the use of a short lag length is required,[9] while being overly optimistic about the rate of impulse

[7] There is an ongoing discussion about the extent to which central banks should respond to inflationary phases, see Chung et al. (2020), Adam et al. (2022), Lane (2024) and Reifschneider (2024) to name but a few.

[8] Mishkin (2007) finds at least six channels: direct effects of interest rates on the user cost of capital, expectations of future housing price movements, and housing supply; and indirect through standard wealth effects of housing prices, balance sheet/credit channel effects on consumer spending, and balance sheet/credit channel effects on housing demand.

[9] A short lag length is essential in VARs to avoid the problem of overfitting and running into the so-called "curse of dimensionality," i.e., too many variables and/or too many lags in the model lead to overparameterization and imprecise estimates. In addition, autoregression (AR) inference methods that rely on the delta method break

response decay (given the stability of the VAR). As a result, the literature faces major challenges in analyzing impulse responses at longer horizons. For instance, Antolin-Diaz & Surico (2022) show that by employing Bayesian shrinkage and a sufficiently large sample size, it might be possible to uncover impulse responses at horizons that are usually overlooked in business cycle analysis.[10] Similar methods are employed in Carriero et al. (2023) to analyze the long-run effects of uncertainty shocks on economic output, finding shocks to produce two dips in output over the long horizon. Inoue & Kilian (2020) address some of the above issues by suggesting lag-augmentation in the AR context. However, as shown in Montiel Olea & Plagborg-Møller (2021), when the data exhibit near-unit-root behavior, the lag-augmented AR estimator of the impulse responses could be criticized as being inconsistent at any horizon growing at a rate that is faster than the square root of the sample size, causing the confidence intervals to explode with length.

Given the difficulties of VAR in dealing with longer horizons, the empirical strategy in our paper relies on well-established local projection techniques to analyze medium- and longer-term dynamics. In this context, Jordà et al. (2023) argue that LP methods are able to produce accurate estimates of the impulse response function (IRF) at very long horizons.[11] Particularly, as long as the truncation lag in local projections is chosen to grow with the sample size at a particular rate, local projections estimate the impulse response consistently at any horizon (Xu 2023). The rate of decay in impulse responses is less strong and, in the context of the bias-variance trade-off, the efficiency loss of LPs relative to VARs diminishes with the number of lags included, effectively vanishing in the limit. A handful of papers analyze the impact of monetary policy on housing using LPs (see Hülsewig & Rottmann 2021, Adra & Menassa 2022, Miles & Zhu 2023, for instance), although this is mostly because these techniques are well suited to observing regime switching or country-specific responses.

In a recent paper, Gorea et al. (2024) employ LP methods to provide evidence on the causal relationship between interest rates and US housing prices using listings data. Their empirical strategy is, in part, the closest to our methodology. While Gorea et al. (2024) find that houinge prices respond to a contractionary monetary policy surprise within two weeks, their long-run analysis points to the existence of a relationship over extended horizons. However, their focus is confined to a short-term horizon of up to one year. In contrast, our paper focuses on time series dynamics and explicitly explores the longer-term impact of monetary policy on housing prices, using LA-LP techniques as per Montiel Olea & Plagborg-Møller (2021). By including additional lags as controls in the LP regression, the standard confidence intervals can be assumed to have correct asymptotic coverage, avoiding a possible efficiency loss relative to VARs, even at more distant horizons. This is empirically confirmed by Buda et al. (2023) and Aruoba & Drechsel (2024), among others. Consequently, in our paper, the robust inference of LA-LP enables the identification of new dynamics between monetary policy and housing markets. We are able to

down, among other things, when suffering from unit roots or when looking at long horizons.

[10] Using a Bayesian VAR (BVAR), characterized by a very high-order lag polynomial, they provide evidence that unexpected increases in government spending are capable of boosting economic activity twice over time. However, they also show that their choice of lag length brings the impulse responses obtained by the BVAR close to what would have been obtained with lag-augmented LPs.

[11] It is evident when comparing LP to conventional VAR procedures, the latter often delivers significantly smaller standard errors than LPs. However, this comes at the cost of fragile coverage (i.e., bias-variance trade-off; Li et al. 2022).

show that housing prices fall not once but twice after a contractionary policy shock, indicating a double-dip behavior of prices in the long run.

Finally, our paper contributes to the literature on shock identification, such as Romer & Romer (2004), Gürkaynak et al. (2005), Jarociński & Karadi (2020). To analyze to what extent the housing deposit channel adds to the double-dip dynamics of housing prices, this paper uses unanticipated interest rate surprises around narrow windows on FOMC meeting days, while examining changes in zero maturity deposits in the week following the surprise interest rate changes. The "poor man's" sign restriction approach of Jarociński & Karadi (2020) is applied. Double-dip movements are more pronounced for shocks that account for the housing deposit channel of monetary policy than for those that do not.

In the remainder of this paper, section 2 explains the fundamentals of LP techniques and introduces the lag-augmented LP process. Section 3 then empirically assesses the long-term effects of contractionary monetary policy. Section 4 provides robustness checks. Section 5 empirically analyses the housing deposit channel and assesses the monetary policy efficiency. Section 6 analyses the housing deposit channel theoretically within in a general equilibrium setup. Finally, section 7 concludes.

# 2. Methodology

## 2.1. VAR and local projection

Suppose a VAR process can be written as,

$$y_t = \sum_{j=1}^{\infty} A_j y_{t-j}; \quad j = 1, 2, \dots . \tag{1}$$

Under standard regularity assumptions, a VAR($p$) process provides consistent estimates of the autoregressive matrices $A_1, ..., A_p$, so that $\|\hat{A}_j - A_j\| \xrightarrow{p} 0$, with $p, T \to \infty$ as long as $p$ grows at rate $p^2/T \to 0$ (as per Lewis & Reinsel 1985). However, as shown by Jordà et al. (2023), in a finite sample, inconsistent estimates of the impulse response function arise if (1) the truncation lag is too short ($k < p$); and (2) if the truncation lag is appropriate, but the impulse response is calculated for periods that extend beyond the truncation lag. Furthermore, even if one decides to specify the VAR truncation lag, $k$, to be as large as the impulse response horizon (as long as $k^2/T \to 0$, as seen in Lewis & Reinsel 1985), issues of parametric load and overparameterization would still persist (Kuersteiner 2005). As a result, even though long lags are chosen, the task of analyzing gets more difficult or even impossible (Jordà et al. 2023).

To overcome these issues, Jordà (2005) proposes computing impulse responses using local projections instead of VARs. For the sake of clarity and consistency throughout the paper, consider the $AR(1)$ model for the data, $y_t$,

$$y_t = \rho y_{t-1} + u_t, \quad t = 1, 2, ..., T, \quad y_0 = 0. \tag{2}$$

The parameter of interest is a nonlinear transformation of $\rho$, namely, the impulse response coefficient at horizon $h \in \mathbb{N}$. This parameter is denoted by $\beta(\rho, h) \equiv \rho^h$. Furthermore, it is

assumed that $u_t$ is strictly stationary and satisfies $E(u_t|\{u_s\}_{s\neq t}) = 0$. The assumption requires the innovations to be mean independent relative to past and future innovations.

For conducting inference with regard to the impulse response $\beta(\rho, h)$, it is worth taking the benchmark (i.e., non-augmented) local projection approach of Jordà (2005) into consideration. A common motivation for this approach is that the $AR(1)$ model implies

$$y_{t+h} = \beta(\rho, h)y_t + \xi_t(\rho, h), \tag{3}$$

where the regression residual (or *multi-step forecast error)*,

$$\xi_t(\rho, h) \equiv \sum_{j=1}^{h} \rho^{h-j} u_{t+j}, \tag{4}$$

is generally serially correlated, even if the innovation $u_t$ is independent and identically distributed (i.i.d.).

As can be seen, (non-augmented) local projections are basically direct estimates of the impulse response (moving average) coefficients, which straightforward regress $y_{t+h}$ on $y_t$. Truncating the lag structure, even when $h > k$, has asymptotically vanishing effects on the consistency of the estimator. Recent work by Xu (2023) shows that in settings where the true lag order is unknown and possibly infinite, LPs are semiparametrically efficient as long as controlled lags are allowed to grow with the sample size. This means that the efficiency loss of LPs relative to VARs (i.e., bias-variance trade-off) diminishes the more lags one includes, and it effectively vanishes at the limit (Jordà 2023). In contrast, truncated VARs must be inverted in order to construct the impulse response. The impulse response therefore depends on the entire dynamic specification of the VAR. The cumulation of small sample inconsistencies over increasing horizons can accumulate and turn into non-negligible distortions of the impulse response, especially at long horizons (Jordà et al. 2023).

In this context, a Monte Carlo experiment by Jordà et al. (2023) compares cumulative response estimates of a simple AR model with three, six, nine, and twelve lags to LPs with only two lags. They find that an AR process with less than twelve lags produces cumulative effects that are relatively short-lived and far removed from the true impulse when analyzing long-run dynamics, completely missing the unwinding of the dynamics of periods 1-12 that follows in periods 13-24. Only the early stages of the impulse response are captured. In contrast, local projections, even when truncated at short lags, provide a fairly close estimate of the true responses, especially at long horizons.

However, as shown by Montiel Olea & Plagborg-Møller (2021), the validity of this non-augmented LP approach, as well as of standard VAR processes, is sensitive to the persistence of the data. Specifically, it may lead to a non-normal limiting distribution for the impulse response estimator when $\rho \approx 1$, since the regressor $y_t$ exhibits near-unit-root behavior in this case. Hence, inference based on normal critical values will not be valid uniformly over all values of $\rho \in [-1, 1]$, even for fixed forecast horizons $h$. Only when $\rho$ is safely within a stationary region, is the LP estimator considered to be asymptotically normal. Yet, standard LP inference generally requires the use of HAC/HAR standard errors to account for serial correlation in the residual $\xi_t(\rho, h)$, which could be challenging when dealing with the small sample sizes available

in macroeconomics.

### 2.2. Lag-augmented local projection

To robustify and simplify inference, by solving the issues of data persistence and the interest in long impulse response horizons, Montiel Olea & Plagborg-Møller (2021) propose a *lag-augmented* local projection, which adds $y_{t-1}$ as an additional control variable and provides asymptotically valid inference for both stationary and non-stationary data over a wide range of response horizons.

Let's define the covariate vector $x_t \equiv (y_t, y_{t-1})'$.[12] Given any horizon $h \in \mathbb{N}$, the lag-augmented LP estimator $\hat{\beta}(h)$ of $\beta(\rho, h)$ is specified by the coefficient on $y_t$ in a regression of $y_{t+h}$ on $y_t$ and $y_{t-1}$:

$$\begin{pmatrix} \hat{\beta}(h) \\ \hat{\gamma}(h) \end{pmatrix} \equiv \left( \sum_{t=1}^{T-h} x_t x_t' \right)^{-1} \sum_{t=1}^{T-h} x_t y_{t+h}, \quad x_t \equiv (y_t, y_{t-1})'. \tag{5}$$

Here $\hat{\beta}(h)$ is the impulse response estimator of interest, while $\hat{\gamma}(h)$ is a nuisance coefficient. The purpose of the lag augmentation is to make the effective regressors of interest stationary even when the data $y_t$ have a unit root. Note that one would obtain the same $\hat{\beta}(h)$ estimate, if regressing instead on $(u_t, y_{t-1})$, since $u_t$ is simply a linear transformation of $y_t$ and $y_{t-1}$, in that $u_t = y_t - \rho y_{t-1}$. Note further that the equations 3-4 imply

$$y_{t+h} = \beta(\rho, h) u_t + \beta(\rho, h+1) y_{t-1} + \xi_t(\rho, h). \tag{6}$$

If $u_t$ were observed, the above equation suggests regressing $y_{t+h}$ on $u_t$, while controlling for $y_{t-1}$. Intuitively, this will lead to an asymptotically normal estimator of $\beta(\rho, h)$ since the regressor of interest $u_t$ is stationary by assumption, and since the term that involves a possible non-stationary regressor, $y_{t-1}$, is controlled for. The coefficient $\hat{\beta}$ converges to the coefficient $\beta$. In other words, $u_t$ being a stationary variable, no matter the persistence (i.e, $\rho$) and no matter the horizon, one will always be regressing on a stationary regressor, leading to a uniformly normal limit distribution.[13]

Moreover, lag augmentation has the additional benefit of simplifying the computation of standard errors, as shown by Montiel Olea & Plagborg-Møller (2021). As equation 4 indicates, the residuals of a local projection generally have a moving average structure. Because they are dated $t$ to $t+j$, they do not affect the consistency of the local projection estimate $\hat{\beta}(\rho, h)$. However, the residual correlation affects the construction of standard errors (Jordà 2023). As already mentioned above, Montiel Olea & Plagborg-Møller (2021) show that, contrary to the solution offered by Jordà (2005) or Ramey (2016) of using Newey-West heteroscedasticity and

[12] This paper only provides a brief overview of the LA-LP theory in the context of a simple univariate $AR(1)$ model. The discussion merely intends to illustrate the main points and can easily be extended to a multivariate $VAR(p)$ model. However, as the formal uniformity result for a general $VAR(p)$ model would be too extensive in the context of our paper, and the application in an AR(1) is sufficient to capture the main idea of LA-LP, the reader is referred to Montiel Olea & Plagborg-Møller (2021) for further details and proof of the inference procedure for lag-augmented LPs in a $VAR(p)$ setting.

[13] This argument for why lag-augmented LP can be expected to have a uniformly normal limit distribution, even when $\rho \approx 1$, is analogous to the reasoning for using lag augmentation in AR inference, see Sims et al. (1990), Toda & Yamamoto (1995), Dolado & Lütkepohl (1996), Inoue & Kilian (2002,0).

autocorrelation consistent/robust estimators, HAC/HAR standard errors are not necessarily needed to conduct inference on *lag-augmented* LP. This is true despite the fact that the regression residual $\xi_t(\rho, h)$ might be serially correlated. Instead, Montiel Olea & Plagborg-Møller (2021) propose that it is sufficient to use usual heteroscedasticity-robust Eicker-Huber-White standard error of $\hat{\beta}(h)$, $\hat{s}(h)$ (see Appendix A).

Relying on the Eicker-Huber-White standard errors, $\hat{s}(h)$, one can define the nominal $100(1-\alpha)\%$ lag-augmented LP confidence interval for impulse response at horizon $h$ as,

$$\hat{C}(h,\alpha) \equiv \left[\hat{\beta}(h) - z_{1-\alpha/2}\hat{s}(h), \hat{\beta}(h) + z_{1-\alpha/2}\hat{s}(h)\right], \tag{7}$$

where $z_{1-\alpha/2}$ is the $(1-\alpha/2)$ quantile of the standard normal distribution. The main result shown by Montiel Olea & Plagborg-Møller (2021) is that the lag-augmented LP confidence interval above is uniformly valid regardless of the persistence of the data, i.e., whether or not the data have a unit root. Crucially for our paper, this result does not break down at moderately long horizons $h$, and the inference is valid even under the *worst-case choices* of parameter $\rho \in [-1, 1]$ and horizon $h \in [1, \bar{h}_T]$ (see Appendix B).

In summary, a simple way to obtain correct inference for the local projection is to add an additional lag as a regressor and then select a heteroscedasticity-robust estimator to compute the standard errors, as there is no longer a need to correct for serial correlation. For the empirical application of lag-augmented LP, Montiel Olea & Plagborg-Møller (2021) propose a parametric wild bootstrap procedure where data are simulated from a VAR and then local projections are fitted to the stimulated data to construct percentile t-confidence intervals. In the IRF analysis below, both standard delta methods and the parametric wild bootstrap method proposed by Montiel Olea & Plagborg-Møller (2021) to calculate confidence intervals are applied.

## 3. Empirical results

This section evaluates the effects of monetary policy on the housing market by adapting the VAR model described in Jarociński & Smets (2008) (JS, hereafter). JS assess the role of the housing market and monetary policy in the US business cycles from the second half of the 1980s to June 2007. Our paper uses the same variables and the same time period as JS.[14] The primary distinction lies in our adoption of an LA-LP approach to inference. Our paper explicitly accommodates valid IRFs at distant horizons, offering a more conservative outlook on the rate of decay. Simultaneously, we address the issue of excessive regressors in the model.

### 3.1. Data

Following JS, the propagation of a restrictive monetary policy shock in a system of nine US variables including real GDP, the GDP Deflator, commodity prices, the FFR, the money stock M2, as well as real consumption, real housing investment, real house prices, and the long-term

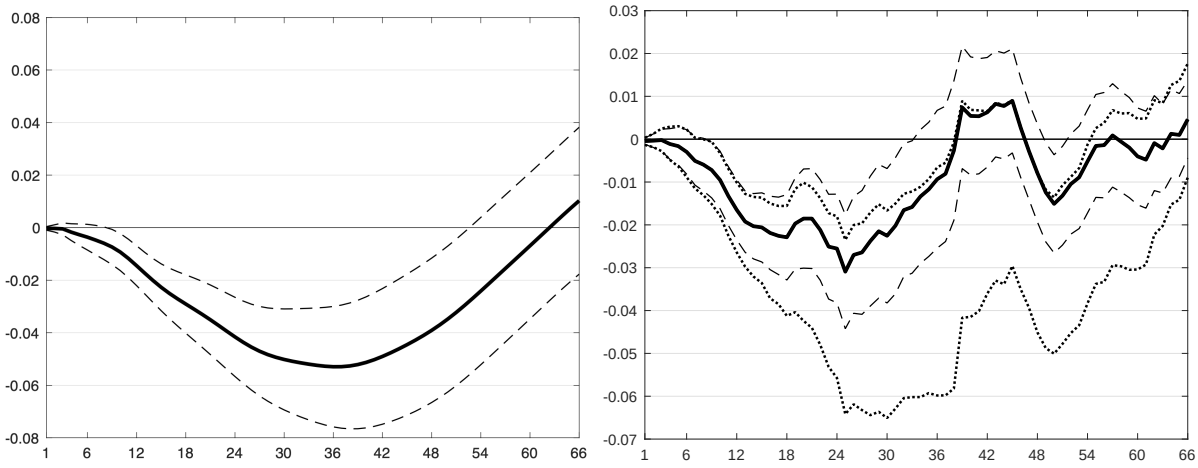
[14] Given the prolonged period of expansionary monetary policy following the financial crisis up to recent years, extending the sample to include a more recent period is not pertinent to the objective of this paper. Nonetheless, we have conducted the same analysis for the 2008-2020 period. Unsurprisingly, the results reveal a less pronounced dynamic regarding a double dip in housing prices, as contractionary monetary policy played an insignificant role after 2008 compared to earlier periods.

interest rate spread is studied. These variables are included at a monthly frequency, which enables a more detailed analysis of housing market developments.[15] Moreover, a higher number of observations is naturally obtained, permitting the extension of the impulse response horizon, while accounting for the short sample bias in LP estimation pointed out by Herbst & Johannsen (2024). All variables are in logs, with the exception of the interest rates and the housing investment share, implicitly allowing for long-run relationships (Calza et al. 2007). To measure housing prices, this paper uses the nationwide Case-Shiller house price index.[16] For additional information regarding the variables, see JS.

Monetary shocks in JS are identified recursively by zero restrictions. They are assumed to affect economic activity and prices with a lag, while having an immediate effect on the term spread and money stock. The original VAR is estimated using a sample running from January 1987 to June 2007.[17]

### 3.2. Baseline

Figure 3a plots the IRFs of housing prices in response to a monetary policy shock, as seen in JS. Housing prices begin to fall immediately after a contractionary monetary policy shock, bottoming out below baseline after approximately two and a half years.




(a) VAR (b) LA-LP

Figure 3: Impulse response functions of housing prices to monetary policy shock using: (a) VAR and (b) LA-LP. Time (horizontal axis) is in months. The solid black lines represent the IRFs. The dashed and dotted lines indicate one standard deviation delta-method and percentile-t bootstrap confidence interval, respectively.

A somewhat different picture can be seen when looking at the IRFs produced by lag-augmented local projections in Figure 3b.[18] Following Jordà et al. (2023) local projection techniques provide close estimates of the true responses at long horizons,[19] even when truncated at a

[15] Monthly values of the quarterly Real GDP are retrieved through interpolation with the industrial production index. Monthly values of the quarterly GDP Deflator are retrieved trough interpolation with the Consumer Price Index and Producer Price Index. Interpolation follows the procedure described in Chow & Lin (1971).

[16] The use of alternative house price indices produces qualitatively similar results, see Appendix C.

[17] The VAR in levels of JS uses standard Minnesota priors. However, no specific priors were included to capture long-run behavior in order to examine dynamics at longer horizons.

[18] A lag length of $p = 5$ is chosen to capture relevant information. We check for robustness using a lag length of $p = 2$ based on the Schwarz information criterion for monthly data. However, as shown by Jordà et al. (2023), the Schwartz criterion tends to select lag lengths that are too small. Accordingly, we also consider lag lengths of $p = 3$, $p = 4$, and $p = 6$. The estimated IRFs do not change significantly when the lag length is changed.

[19] Following Xu (2023), together with the lag length of $p = 5$, we chose a horizon, $h$, which is small enough so

low lag length. In this context, the lag-augmented LPs in particular provide asymptotically valid inference over a wide range of response horizons (Montiel Olea & Plagborg-Møller 2021). Both standard delta methods (in Figure 3b, dashed lines) as well as the parametric wild bootstrap methods (in Figure 3b, dotted lines) are used to calculate confidence intervals. Similar to the VAR responses, real house prices immediately fall after impact. However, they revert to their pre-shock levels after roughly three years. Following this recovery period, a second recessionary wave at a 4 year horizon is evident. The second contraction is clearly visible and significant, while being less strong.

The above dynamics suggest a new mechanism that we call it, housing deposit channel of monetary policy, described as follows: an initial increase in the FFR triggers an immediate dip in housing prices as effective housing demand declines. In turn, a sudden increase in the policy rate by the central bank makes holding cash more expensive. Banks have the effective market power to raise the deposit spread. As a result, lender households pull back on their bank deposits, causing deposits to flow out of the banking system (for me details, see also Drechsler et al. 2017). This outflow finds its way into concentrated markets. With housing being the primary asset on household balance sheets in many advanced economies and offering considerable utility, these displaced funds often flow into the housing market. Lower housing prices create an incentive for lender households to reallocate their funds, financing purchases through the housing deposit channel. This shift plays a crucial role in driving the recovery of housing prices.

As prices recover, collateral constraints relax, and demand for mortgages rises. This increases banks' demand for deposits that had previously been withdrawn in significant amounts. However, the high demand for deposits after the initial recovery reduces the demand for housing in the markets and pushes housing prices down for a second time, although not as low as in the initial decline.

In summary, a contractionary monetary policy shock is empirically found to trigger a double-dip effect in housing prices: an initial decline immediately after the shock's impact, followed by a second decline in the long run after a preceding period of increase. Further analysis of the housing deposit channel and monetary policy efficiency is provided in section 5.

## 4. Robustness checks

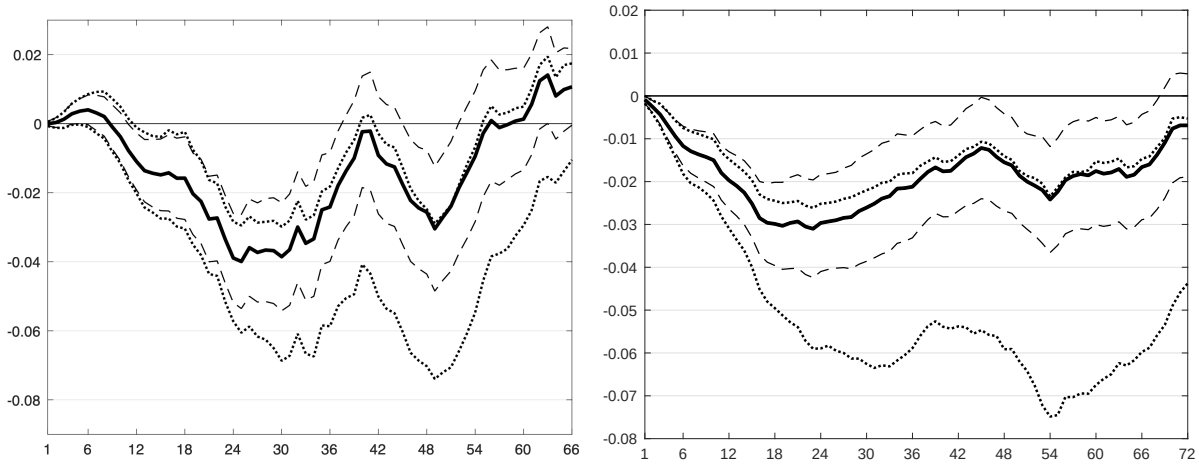
The following section provides a robustness check of the baseline analysis. Firstly, endogenous movements in the FFR are separated from exogenous monetary policy shocks, using Romer & Romer (2004) like shocks. The corresponding IRFs yield similar results to the baseline estimation. Secondly, the choice of alternative variables, as per the studies by Calza et al. (2007) and Calza et al. (2013), also reveal double-dip dynamics of housing prices in response to interest rates. Thirdly, shocks other than to interest rates do not lead to the same outcome. Although credit risk and uncertainty shocks also cause housing prices to decline in the short run, there is no evidence of a second dip in the longer run.

that $n \geq 3h + p - 3$. Here, we use a horizon of $h = 66$, binding $n \geq 200$, which is the case for used observations from 1987m1 to 2007m6 ($n = 246$).

### 4.1. Alternative monetary policy shocks

Since interest rate changes may be anticipated by markets, potentially distorting the effect of real interest rate hikes on prices, our paper follows Romer & Romer (2004) (RR, hereafter) in identifying monetary policy innovations that are cleansed of anticipatory effects related to economic conditions. RR first construct a historical measure of changes in the target FFR at each Federal Open Market Committee (FOMC) meeting from 1969 to 1996. Using the real-time Fed staff forecasts presented in the Greenbooks prior to each FOMC meeting, they construct a measure of monetary policy shocks from the component of policy changes at each meeting that is orthogonal to the Fed's information set. Following Coibion et al. (2017), the RR policy shock is extended to December 2008.

The IRFs shown in Figure 4a, are quite similar to those in Figure 3b. A drop is observed within two years, followed by a gradual recovery to almost pre-shock levels. As in the previous section, the LA-LP method reveals responses at long horizons. One discerns a second wave of decline after the initial shock, which leads to a further dip in housing prices.




(a) Romer & Romer (2004) (b) Calza et al. (2013)

Figure 4: Impulse response functions of housing prices to monetary policy shock with different specification: (a) Romer & Romer (2004) vs (b) Calza et al. (2013). Time (horizontal axis) is in months. The solid black lines represent the IRFs. The dashed and dotted lines indicate one standard deviation delta-method and percentile-t bootstrap confidence interval, respectively.

These analyses confirm the robustness of our conclusions, particularly with respect to the response of housing prices to monetary policy shocks. By using an alternative monetary policy shock, the results obtained in section 3.2 are effectively validated. This consistency across different shocks underscores the reliability and generalizability of our findings and provides additional confidence in the conclusions drawn from our study.

### 4.2. Alternative specification

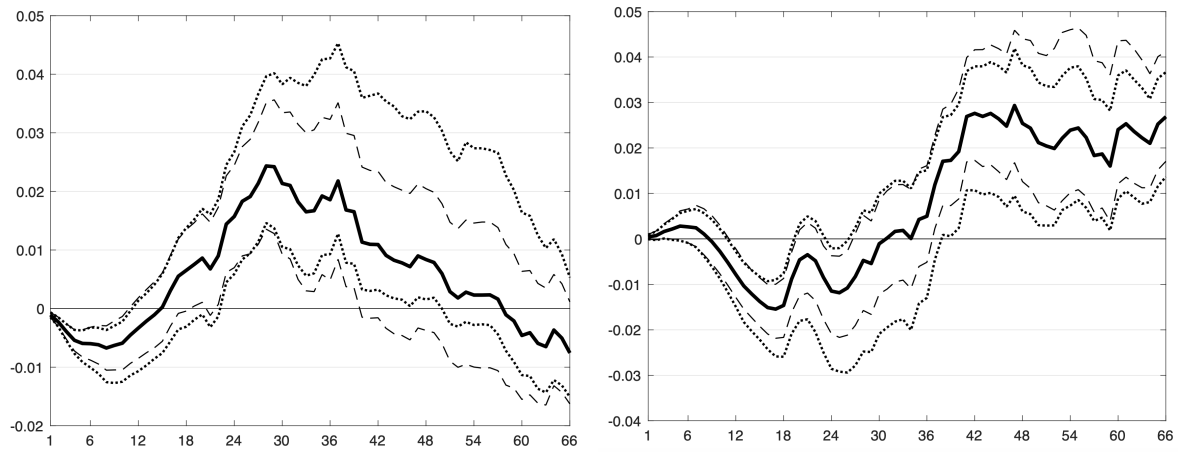
The aforementioned characteristics should also be evident when examining other specifications. For example, when assessing the propagation of monetary shocks in a set of endogenous variables used by Calza et al. (2013) and adopting their identification scheme, comparable conclusions emerge. Figure 4b shows the IRFs of real house prices generated by LA-LPs, including, in this order, private consumption, residential investment, the consumer price index, the real house

price index, a three-month interbank interest rate, and the real effective exchange rate.[20] It should be noted that, instead of using a shock to the FFR as a contractionary monetary policy change, Calza et al. (2013) apply an innovation to the three-month interbank interest rate. This may lead to slightly different results as transmission does not necessarily work the same way or is less pronounced when considering interest rates other than the FFR.[21]

Indeed, this becomes evident when examining the housing price responses. Although the IRFs show a significant post-shock decline followed by a steady recovery to pre-shock levels, the recovery period is less pronounced than in the other models. However, a second drop can be observed after the recovery period. As in the previous section, our lag-augmented LP technique reveals long-term responses that would be overlooked in other models.

### 4.3. Do other shocks produce similar IRFs?

Finally, our paper considers whether other shocks could also lead to double-dip dynamics similar to those seen in the previous sections. Indeed, the literature finds that shocks to uncertainty or financial risk, among others, can cause housing prices to fall (see, for instance El-Montasser et al. 2016, Gertler & Gilchrist 2019, Balcilar et al. 2021, Becard & Gauthier 2023). However, the transmission channels are different, and it is questionable whether these types of innovations also lead to a second decline, as in the case of responses to a monetary policy shock.




(a) Uncertainty shock (b) Credit spread shock

Figure 5: Impulse response functions of housing prices to uncertainty and credit spread shock: (a) uncertainty shock and (b) credit spread shock. Time (horizontal axis) is in months. The solid black lines represent the IRFs. The dashed and dotted lines indicate one standard deviation delta-method and percentile-t bootstrap confidence interval, respectively.

To measure the tightness of financial market conditions, our papers relies on corporate bond spreads. Specifically, the excess bond premium (EBP) is included in the data set[22] from Gilchrist & Zakrajšek (2012). This is an estimate of the extra compensation that bond investors demand for bearing the credit risk of US non-financial corporations. According to Gilchrist & Zakrajšek (2012), the EBP has predictive power for future economic activity, with shocks to the

[20] All variables are in logs except for the interest rates. A lag length of two is chosen by Calza et al. (2013), which also seems appropriate for a model with monthly frequency data, following the Schwarz information criterion.

[21] We also employ the variables used in Calza et al. (2007) as robustness, whereby residential investment is excluded from the aforementioned data set. Similar IRFs are obtained.

[22] Data provided by the US Federal Board System (Giovanni et al. 2016).

excess bond premium that are orthogonal to the current state of the economy leading to declines in economic activity and asset prices. An increase in the EBP appears to reflect a reduction in the risk-bearing capacity of the financial sector, leading to a contraction in the supply of credit, a deterioration in macroeconomic conditions and a collapse in housing prices.

Similarly, uncertainty events can affect the dynamics of the housing market. For example, under uncertainty, consumers tend to reduce their consumption and investors tend to delay their investments (Bloom 2009) due to precautionary motives, thereby depressing both the demand and supply in the housing market. Higher risk premium for financing projects in the presence of uncertainty increases the cost of debt financing (Christiano et al. 2014). To measure uncertainty, the Economic Policy Uncertainty (EPU) index constructed by Baker et al. (2016) is used. The EPU is an index based on newspaper articles about policy uncertainty. It counts the number of newspaper articles containing the terms uncertain or uncertainty, economic or economy, and one or more policy-related terms.

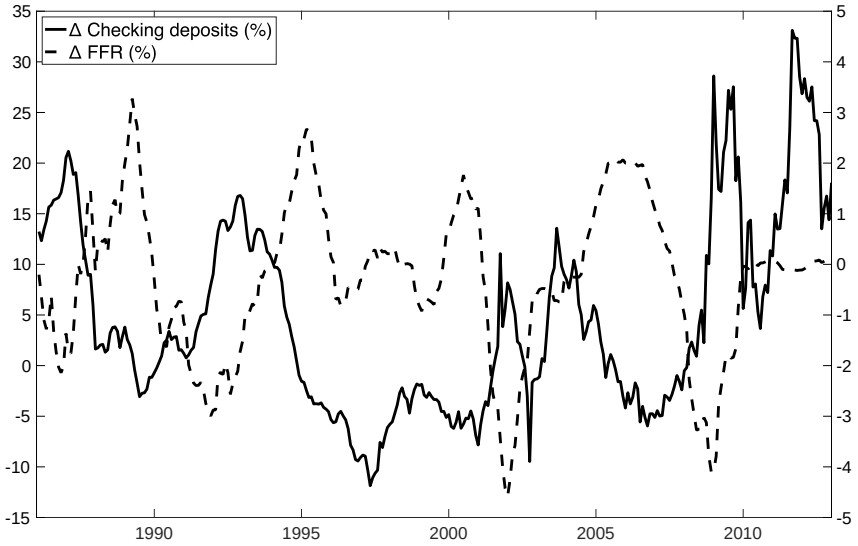
Figures 5a and 5b show the IRFs of real house prices to an uncertainty (EPU) and credit spread (EBP) shock, respectively, using LA-LPs with control variables from JS. As suggested by classical economic theory, both impulse responses show a negative reaction after the respective shock, before recovering in the long run. In contrast to a monetary policy shock, there is no double-dip in housing prices.

## 5. Housing deposit channel, monetary policy efficiency and inflation targeting

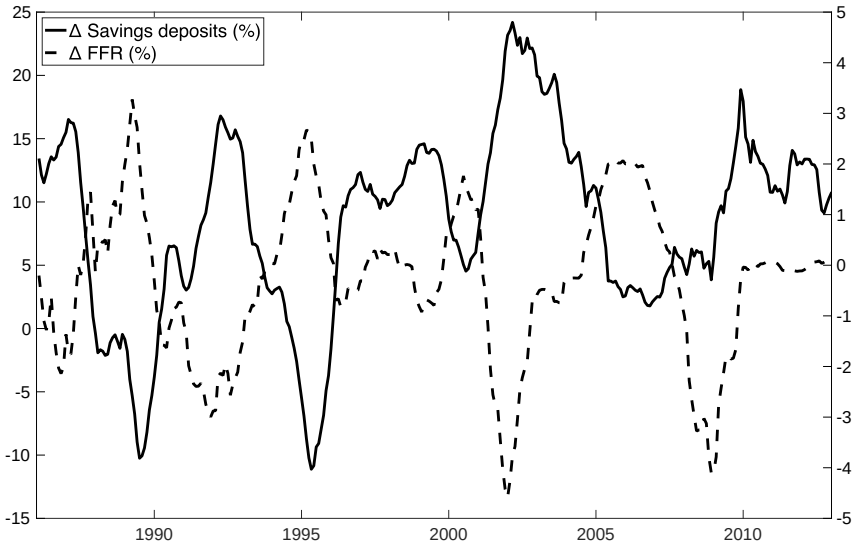
This section assesses the role of the housing deposit channel in transmitting monetary policy. The process starts via the deposit channel. According to Drechsler et al. (2017), banks have a certain degree of market power that allows them to increase the deposit spread, the difference between the FFR and the deposit rate, following an increase in the FFR. Households face a higher cost of holding deposits as opposed to cash or other investments and consequently withdraw deposits from the bank.

In this context, Drechsler et al. (2017) find that deposit spreads and zero maturity deposits immediately respond to rate changes in the FFR. Figure 6 plots the year-over-year change in the FFR against the percentage growth rate in the aggregate amounts of checking deposits (Figure 6a) and savings deposits (Figure 6b). These empirical insights reveal clear relationships. Figure 6a depicts a significant negative correlation of -35% between the growth rate of checking deposits and fluctuations in the FFR. Figure 6b reinforces the findings, revealing an even more pronounced correlation of -59% between changes in the FFR and the growth rates of liquid savings deposits.

Consequently, the reaction of zero maturity deposits to an interest rate change can help distinguish between shocks transmitted through a deposit channel and those that are not. To distinguish monetary policy shocks that operate through a housing deposit channel of monetary policy from those that do not, our paper uses unanticipated interest rate surprises around narrow windows on FOMC meeting days as a variable for a monetary policy shock, as per Romer & Romer (2004) and Coibion et al. (2017). We then check for changes in deposits in the week




(a) Checking deposits




(b) Savings deposits

Figure 6: Deposit and interest rate growth: (a) checking deposits and (b) savings deposits. Left vertical axis indicates year-over-year growth of deposits in %. Right vertical axis indicates year-over-year growth of federal funds rate in %. Data is provided by the Federal Reserve System.

following the surprise rate changes.[23] Interest rate changes that affect the economy through the deposit channel are identified using the "poor man's" sign restriction approach of Jarociński & Karadi (2020): Only interest rate surprises that are accompanied by opposite sign deposit changes in the week following FOMC meetings are considered.

The responses of housing prices to these shocks are depicted in Figure 9. The paper finds a similar dynamic as in depicted in Figure 4a and as will be theoretically shown in section 6. An immediate dip in housing prices after impact is followed by a recovery period to pre-shock levels, and finally a second dip. This proves that monetary policy is transmitted to the housing market through a deposit channel. In addition, one observes that the two dips are of greater magnitude than those observed after a "normal" interest rate shock in the previous section by

[23] Changes in other zero maturity deposits, such as savings deposits, were also considered. They generate similar IRFs.

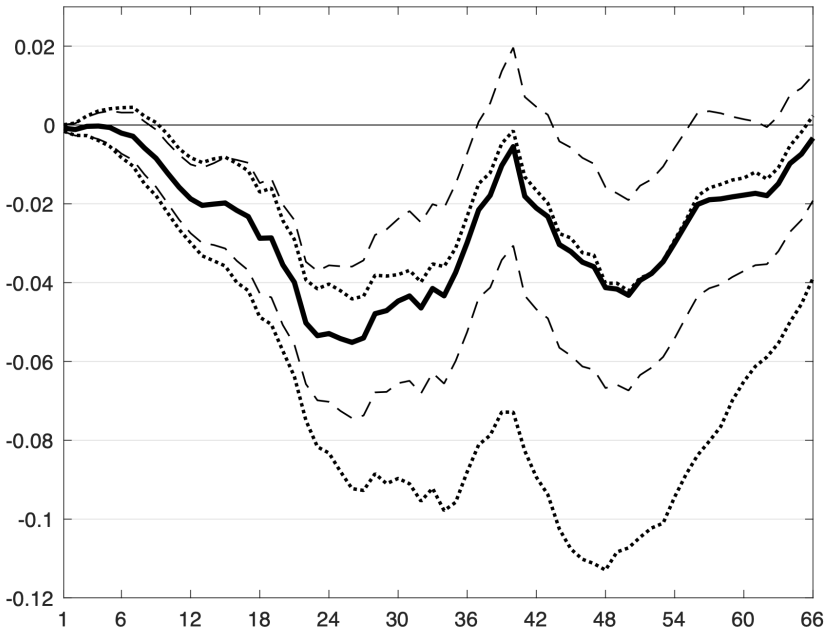



Figure 7: Impulse response function of housing prices to monetary policy shocks transmitted through housing deposit channel. Time (horizontal axis) is in months. The solid black line represents the IRF. The dashed and dotted lines indicate one standard deviation delta-method and percentile-t bootstrap confidence interval, respectively.

comparing Figures 3b and 4a.[24]

These findings have significant implications for the effectiveness of monetary policy, as they contribute to the discussion on the role of housing prices in central banks' inflation targeting (see Bernanke & Gertler 2001, Cecchetti et al. 2003, for instance). While inflation-targeting central banks explicitly aim to control inflation, most do not target housing prices or house price inflation (Ingves 2007). However, as Iacoviello & Minetti (2003) explain, an increase in housing prices raises the collateral value of properties, thereby enhancing the borrowing capacity of debtors. Consequently, rising housing prices can signal future inflationary pressures and should be considered in monetary policy decisions, sparking a debate on the inclusion of housing prices in the Taylor rule (for a detailed literature review see Käfer 2014).

This divergence in the literature on defining the role of housing prices in inflation targeting has led to the emergence of a variety of perspectives, particularly after the recent housing price hikes in Canada and the US.[25] On the one hand, a substantial body of research supports the "flexible inflation-targeting" approach. The central argument is that the costs associated with policies that explicitly take housing prices into consideration, such as output losses and increased unemployment, outweigh the potential benefits of reducing the probabilities of future financial crises (see, for instance, Svensson 2017). In this context, Benati (2021) utilizes VAR techniques and finds that the impact of monetary shocks on real house prices is approximately three times greater than on real GDP in the US and Canada. This suggests that policies aimed at counteracting housing price fluctuations would lead to significant GDP losses.

[24]We also analyzed shocks that do not pass through the deposit channel, i.e., changes in the federal funds rate followed by changes in deposits with the same sign. The IRFs show a quite different pattern, as can be seen in Appendix D.

[25]See for example, Bhar et al. (2024)

In addition, for example, Duca et al. (2021) outline the factors that drive housing prices and their impact on economic activity. The study highlights both theoretical and empirical research that links housing markets with credit markets, housing price expectations, financial stability and the broader economy. These findings underscore the importance of incorporating housing price dynamics into the framework of monetary policy in order to achieve a more holistic approach to economic stability and growth.

In light of these newly emerging developments on the housing market, our paper provides new evidence to help develop a consensus on the implementation of inflation-targeting monetary policy. For instance, most inflation-targeting central banks have opted for forecast horizons of two or three years. Benati (2021), for example, provides evidence suggesting that the trade-offs between housing prices and GDP are largely a short-term phenomenon. However, our paper demonstrates that interest rate changes have a long-lasting impact on housing prices, both immediately and over medium and longer horizons.

Our findings indicate that monetary policy should not focus solely on the immediate effects of house and asset prices on inflation and the real economy. Instead, it is crucial to consider the longer-term consequences of these dynamics. The medium- and long-term behavior of housing prices can have important implications for the formulation of monetary policy. Consequently, a more comprehensive approach that takes into account the lasting effects of interest rate adjustments on housing markets is necessary in order to enhance the effectiveness and stability of inflation-targeting strategies.

On the other hand, there is support for an alternative monetary policy approach known as the "leaning against the wind" theory. This theory advocates policies that systematically react to signs of "excessive" credit growth or "disequilibria" in asset prices by slightly raising interest rates above the levels required solely to control inflation and real economic activity (Bank for International Settlements 2016). Proponents argue that such a strategy could mitigate the build-up of financial imbalances and reduce the risk of future economic downturns by allowing asset prices to be taken into account in monetary policy (i.e., extending the Taylor rule to include housing prices).

Our results show that a policy of "leaning against the wind", which involves further increases in interest rates to cool the housing market, can be counterproductive. The empirical model suggests that housing prices recover in the medium term after a contractionary monetary policy shock. However, it also suggests the existence of a second, albeit smaller, dip in housing prices over longer horizons due to the lingering effects of the initial interest rate shock. Therefore, policymakers should not be overly concerned with short-term fluctuations in housing inflation, as housing prices are likely to decline again in the long run after the transmission of interest rate hikes through the housing deposit channel. This finding underscores the importance of considering the long-term dynamics of the housing market when formulating monetary policy, rather than simply responding to immediate changes in housing prices.

## 6. The theory of housing deposit channel

This section presents a general equilibrium model to elucidate the dynamics of the housing deposit channel. The model provides more insight into the intricacies that influence housing

prices. This gives support to the housing deposit channel as analyzed empirically in the previous sections. The model includes two types of households in the same style as Iacoviello (2015), in order to distinguish between different reactions to interest rate hikes. Superscripts $l$ and $b$ denote lender (patient) and borrower (impatient) households, respectively. Each type of household is represented by a unit measure of infinitely-lived agents. The model further includes a central bank and a financial sector. As the focus is on the households' behavior, the central bank and financial sector are simplified. The model is nominal, and all prices in the economy are relative, with the price of consumption serving as the numeraire.

### 6.1. Lender household

The household's problem is

$$\mathbb{E}_t \sum_{j=0}^{\infty} \beta_l^j \left[ \log c_{t+j}^l + \chi_l^h \log h_{t+j}^l + \chi_l^d \log d_{t+j} \right]$$

$$s.t$$

$$c_t^l + d_t + p_t^h (h_t^l - (1-\delta_h) h_{t-1}^l) = w^l + r_{t-1}^d \frac{d_{t-1}}{\pi_t} \tag{8}$$

where $t$ represents time. $\beta_l < 1$ is the lender household discount factor. At every period $t$, the lender household engages in the following activities: consuming non-durable goods $c_t^l$, depositing in financial institutions $d_t$, buying houses $h_t^l$ at the price $p_t^h$ . The household receives $w^l$ as endowment wage (for simplicity's sake, the wage is assumed fixed), $r_{t-1}^d$ as the interest rate on the bank deposits of the last period. $\chi_l^h$ and $\chi_l^d$ are the coefficients which represent the relative importance of housing and liquidity in the utility function, respectively. In this model, bank deposits are considered as liquid assets and they provide liquidity to the household as per Drechsler et al. (2017).

The first-order conditions with respect to consumption, deposit and housing are

$$\lambda_t^l = \frac{1}{c_t^l} \rightarrow \Lambda_{t+1}^l = \beta_l \mathbb{E}_t \frac{\lambda_{t+1}^l}{\lambda_t^l} \tag{9}$$

$$1 = \frac{\chi_l^d}{\lambda_t^l} d_t^{-1} + \Lambda_{t+1}^l r_t^d / \pi_{t+1} \tag{10}$$

$$p_t^h = \frac{\chi_h}{\lambda_t^l} (h_t^l)^{-1} + \Lambda_{t+1}^l p_{t+1}^h (1-\delta_h) \tag{11}$$

where $\lambda_t^h$ is the Lagrange multiplier.

### 6.2. Borrower households

The problem of impatient households is

$$\begin{aligned}
&\mathbb{E}_t \sum_{j=0}^{\infty} \beta_b^j \left[\log c_{t+j}^b + \chi_b^h \log h_{t+j}^b]\right] \\
&s.t. \\
&c_t^b + p_t^h(h_t^b - (1-\delta_h)h_{t-1}^b) + r_{t-1}^m \frac{m_{t-1}}{\pi_t} \leq w^b + m_t \\
&m_t \leq \theta p_t^h h_t^b
\end{aligned} \tag{12}$$

$\beta_b < 1$ is the discount factor of borrower households ($\beta_b < \beta_l$). Borrower households at time $t$ consume $c_t^b$, accumulate housing $h_t^b$, request mortgages $m_t$ at the rate $r_t^m$ from the financial sector and receive total income $w^b$ (for simplicity, the wage is assumed fixed).

The last equation in the borrower problem (12) is the borrowing constraint and standard in the literature, for example in Justiniano et al. (2015), Alpanda & Zubairy (2016), Ghiaie (2020) and Ghiaie (2023), among others. This constraint restricts the borrower household mortgage to a fraction of its housing value. $\theta$ is the Loan-to-Value ratio (LTV) in the mortgage market.

The first-order conditions with respect to consumption, borrower households and mortgage, respectively, are

$$\lambda_t^b = \frac{1}{c_t^b} \rightarrow \Lambda_{t+1}^b = \beta_b \mathbb{E}_t \frac{\lambda_{t+1}^b}{\lambda_t^b} \tag{13}$$

$$p_t^h - \frac{\lambda_t^m}{\lambda_t^b}\theta p_t^h = \frac{\chi_b^h}{\lambda_t^b}(h_t^b)^{-1} + \Lambda_{t+1}^b(1-\delta_h)p_{t+1}^h \tag{14}$$

$$1 - \frac{\lambda_t^m}{\lambda_t^b} = \Lambda_{t+1}^b \frac{r_t^m}{\pi_{t+1}} \tag{15}$$

where $\lambda_t^b$ is the Lagrange multiplier of the budget constraint and $\lambda_t^m$ is the Lagrange multiplier of the borrowing constraint at time $t$.

### 6.3. Central bank and financial sector

As the aim of this section is to explore the housing deposit channel in more depth, for simplicity's sake, this paper assumes both $r_t^d$ and $r_t^m$ are exogenous. $r_t^m$ is set by the central bank and $r_t^d$ by the financial sector. Following Drechsler et al. (2017), the financial sector sees the policy rate $r^m$ and then sets their deposit rate $r^d$ based on their market power where $r^m = r^d + s$ and $s > 0$. $s$ is the spread and depends on the market power of the banking sector. We refer the readers to Drechsler et al. (2017) for more details on $s$ and the bank market power.

### 6.4. Housing price equilibrium

From equation 10 and 11, the steady state[26] of the housing price emerges from the lender side:

$$p_l^h = \frac{\chi_l^h}{\lambda^l(1-(1-\delta_h)(1-\frac{\chi_l^d}{\lambda^l}d^{-1})\frac{1}{r^d})h^l}. \tag{16}$$

Combining equation 14 and 15, and without loss of generality assuming $\theta = 1$, and setting $\delta_h = 0$ for the sake of simplicity, give the steady state of the housing price from the borrower side:

$$p_b^h = \frac{\chi_b^h}{\lambda^b\beta_b(r^m-1)h^b}. \tag{17}$$

Following Iacoviello (2015), the housing supply is assumed constant and normalized to one i.e, $h_t^l + h_t^b = 1$ to focus on the demand side.[27] Given this normalization and the fact that at the equilibrium $p_l^h = p_b^h$, solving equation 16 and 17 for $h^b$ implies

$$h^b = \frac{\frac{\chi_b^h}{\lambda^b(\beta_b r^m-\beta_b)}}{\frac{\chi_l^h}{\lambda^l(1-\beta_l)} + \frac{\chi_b^h}{\lambda^b(\beta_b r^m-\beta_b)}}. \tag{18}$$

Substituting 18 in 17 reveals the housing price dynamics at the equilibrium:

$$p_E^h = \underbrace{\frac{\chi_l^h}{\lambda^l(1-\beta_l)}}_{\text{lender side}} + \underbrace{\frac{\chi_b^h}{\lambda^b(\beta_b r^m-\beta_b)}}_{\text{borrower side}} \tag{19}$$

The right-hand side of equation 19 represents the lender side (first term) and borrower side (second term). The following subsection describes the consequent housing price dynamics which emerge from a contractionary monetary policy.

### 6.5. What drives the housing price double-dip?

In order to respond to this question, it is necessary to direct our attention to equation 19. Both lenders and borrowers exert an influence on housing prices. the housing price double-dip is explained in three phases.

*First dip.* We initially focus on the borrower side. A change in the interest rate $r^m$ has both direct and indirect effects on borrowers. When the interest rate rises, borrower households face higher costs, leading to a direct decrease in the demand for housing. In addition, there is an indirect negative impact through wealth effects, resulting in a decrease in the borrower's consumption $c^b$ (note, $\lambda^b = \frac{1}{c^b}$). As a result, the borrower's side of equation 19 is always a decreasing function of the interest rate $r^m$.

Studies such as Morris & Palumbo (2001) and Krueger et al. (2024), among others, find a strong positive dynamic between wealth effects and household consumption, showing how

[26] We assume $\pi_t = \frac{p_t}{p_{t-1}}$ therefore at the steady state $\pi = 1$. This value does not change the results of our analysis. Detail calculations are provided in appendix E

[27] In Appendix F, we relax this assumption and solve this equation with flexible housing supply. The results indicate that it is appropriate to focus solely on the price movements resulting from housing demand since accounting for a variable housing supply does not alter the conclusions of this paper.

changes in wealth, particularly housing wealth, influence consumption patterns. This relationship has been further explored in the context of monetary policy by Campbell & Cocco (2007) who find that housing wealth has a significant impact on consumption, particularly for credit-constrained households (in our model, borrowers). In addition, Coibion et al. (2017) and Mumtaz & Theophilopoulou (2017) study the distributional effects of monetary policy in the US and the UK, and find significant consumption inequality across different household quantiles. They show that borrower households reduce their consumption following an increase in the policy rate.

*Housing deposit channel.* We now focus on the lender side of equation 19. On the lender side, the response to interest rate changes is more nuanced. While the borrower side is always a negative function of the interest rate, the lender side depends on the pass-through of interest rates from banks to depositors.

As $r^m$ increases, banks decide to increase their deposit spread $s$ and keep the deposit rate $r^d$ low, following the housing deposit channel of monetary policy explained in the empirical analysis of this paper.[28] Therefore, the opportunity costs of holding deposits for lender households increase, leading the utility for deposits, $\chi_l^d$, to decline (see Polo 2021). Simultaneously, since housing prices are low, following the consumption behavior of borrower households, the utility, $\chi_l^h$, and consequently demand for housing of new investors increases, incentivizing lenders to shift their funds into the housing market.[29]

In both scenarios, the first term of equation 19 drives up the housing price. Therefore, if lenders' effect is greater than borrowers' negative wealth effects, such that $\frac{\chi_l^h}{\lambda^l(1-\beta_l)} > \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}$, the housing price will converge to pre-shock levels. This shift can lead to a partial recovery in housing prices after the initial dip caused by the interest rate hike, as observed in the empirical analysis in section 5.

*Second dip.* In response to deposit outflows, banks attempt to attract more deposits to meet mortgage demand and regulatory obligations[30] by raising $r^d$. Additionally, the earlier surge in deposit spreads, which had triggered significant deposit outflows, now intensifies the liquidity constraints of lender households, increasing $\chi_l^d$. These two dynamics which put downward pressure on housing prices, see equation 16, mark the onset of the second dip in housing prices. In this situation, bank deposits become attractive again which prompts lender households to shift their preference toward rebuilding deposits. As a result, housing demand weakens, leading to a decline in prices. However, this second decline is generally less significant than the initial drop, as financial conditions stabilize and the market adapts to the changing liquidity environment.

Moreover, the impact of other factors, such as regulatory changes, financial innovation, and shifts in market expectations, can further influence this dynamic. For example, the introduction of tighter lending standards (see $\theta$ in equation 12) can amplify the negative impact of interest rate increases on borrower demand, while innovations in mortgage products may mitigate this effect by making borrowing more accessible even in a high-interest-rate environment.

Our theoretical model successfully captures the housing price double dip and closely aligns with the empirical patterns observed in our previous sections. The bottom line of this section is

[28] For more details, see section 3.2 and Drechsler et al. (2017).

[29] Case & Shiller (2003) and Mikhed & Zemčík (2009) discuss a similar situation where the sharp increase in housing prices after a period of low prices is due to stock market investors switching to the housing market as a result of poor stock market conditions. An increase in demand for housing created by these investors pushes housing prices up.

[30] For example the capital requirement ratio (CRR) in Basel III and other obligations.

that the key driver of housing price dynamics following a monetary policy shock is the financial sector's pass-through of interest rates to depositors, which activates the housing deposit channel.

## 7. Conclusion

This paper argues that monetary policy influences the housing market well beyond the forecast horizon of inflation-targeting central banks. By using lag-augmented local projection techniques, our findings demonstrate that interest rate hikes not only cause an immediate decline in housing prices but also trigger a subsequent rise, followed by a second drop at later stages, a phenomenon termed the housing price double-dip. Moreover, through an identification strategy examining weekly deposit data for monetary policy surprises, we uncover a new transmission mechanism: the housing deposit channel. This channel emerges as the key factor driving the housing price double-dip.

These empirical findings are supported by a straightforward general equilibrium model, which distinguishes between the reactions of borrower and lender households after an interest rate increase. While borrower demand and consumption are consistently negative functions of interest rates, leading to an initial drop in housing prices, the reaction of lender households to interest rate changes is more complex and can drive prices upward before causing a subsequent decline. This alignment with empirical evidence strengthens our model's ability to capture the mechanisms behind the observed double-dip pattern.

Although the results of our paper illustrate a distinct behavior of housing prices and propose a novel transmission channel of monetary policy, further studies could be conducted to extend its scope. One avenue for future research is to consider the varying exposure of individual banks and their customers, similar to Drechsler et al. (2022), or in different regions of the US, similar to Albuquerque et al. (2024). Moreover, given the recent tightening cycle in the US, Canada, and the euro area, it would be interesting to examine whether similar effects on housing prices emerge in post-COVID-19 monetary policy, in line with Burgert et al. (2024). However, extending the empirical approach to a post-COVID-19 period seems more challenging, given the potential for biased estimates due to the sequence of extreme observations and the limited number of observations available.

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# Appendix A Standard errors

The heteroscedasticity-robust Eicker-Huber-White standard error of $\hat{\beta}(h)$, as proposed by Montiel Olea & Plagborg-Møller (2021), is given by,

$$\hat{s}(h) \equiv \frac{\left(\sum_{t=1}^{T-h} \hat{\xi}_t(h)^2 \hat{u}_t(h)^2\right)^{1/2}}{\sum_{t=1}^{T-h} \hat{u}_t(h)^2}, \tag{A1}$$

where the lag-augmented LP residuals are defined as follows:

$$\hat{\xi}_t(h) \equiv y_{t+h} - \hat{\beta}(h) y_t - \hat{\gamma}(h) y_{t-1}, \quad t = 1, 2, ..., T-h, \tag{A2}$$

and the realized regressor of interest being

$$\hat{u}_t(h) \equiv y_t - \hat{\rho} y_{t-1}, \quad t = 1, 2, ..., T-h, \quad \hat{\rho}(h) \equiv \frac{\sum_{t=1}^{T-h} y_t y_{t-1}}{\sum_{t=1}^{T-h} y_{t-1}^2}. \tag{A3}$$

The reason why the approach of Montiel Olea & Plagborg-Møller (2021) works is, that under the assumption made on $u_t$, the term $\xi_t(\rho, h) u_t$ is serially uncorrelated even if $\xi_t(\rho, h)$ itself is serially correlated. Notice that this result crucially relies on (i) lag-augmenting the local projections and (ii) the strengthening in the assumption of $u_t$ being strictly stationary with $E(u_t | \{u_s\}_{s \neq t}) = 0$ of the usual martingale difference assumption on $u_t$.[31] As mentioned by Jordà (2023), in practical settings, it is important to ensure that enough lags are included to ensure that this condition is met.

# Appendix B LA-LP inference

Consider any upper bound $\bar{h}_T$ on the horizon which satisfies $\bar{h}_T / T \to 0$. This implies that,

$$\inf_{\rho \in [-1,1]} \inf_{1 \leq h \leq \bar{h}_t} P_\rho \left( \beta(\rho, h) \in \hat{C}(h, \alpha) \right) \to 1 - \alpha \quad as \; T \to \infty, \tag{B1}$$

where $P_\rho$ denotes the distribution of the data, $y_t$, under the $AR(1)$ model in equation 4 with parameter $\rho$. This gives the probability that the confidence interval covers the impulse response for that specific parameter, $\rho$, and at any horizon, $h$.

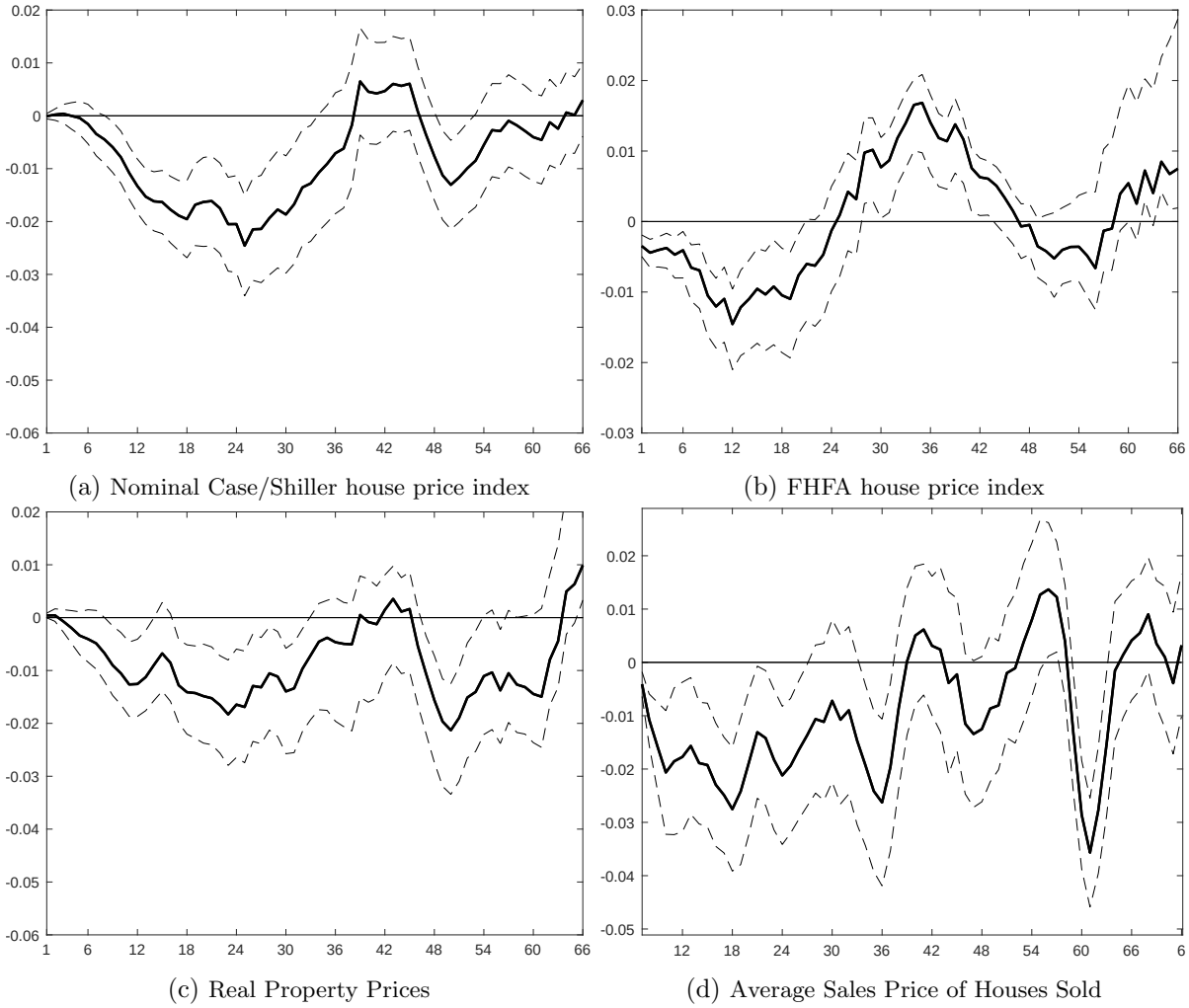
Equation B1 states that, for sufficiently large sample sizes, LP inference is valid even under the *worst-case choices* of parameter $\rho \in [-1, 1]$ and horizon $h \in [1, \bar{h}_T]$. Such *uniform* validity is a much stronger result than *point-wise* validity for fixed $\rho$ and $h$. In fact, if one restricts attention to only the stationary region $\rho \in [-1 + \alpha, 1 - \alpha], \alpha \in (0, 1)$, then the statement in equation B1 is true with the upper bound $h = (1 - \alpha)\bar{h}_T$ on the horizon. That is, if one knows the time series is not close to a unit root, then local projection inference is valid even at long

[31] The strengthening still allows for conditional heteroscedasticity and other plausible features of economic shocks, see Montiel Olea & Plagborg-Møller (2021).

horizons $h$ that are non-negligible fractions of the sample size $T$.

## Appendix C IRF: Alternative house price variables

Figures 8a-d show the responses of alternatives to the real house price index used in the paper to a monetary policy shock. The nominal nationwide Case-Shiller house price index (Figure 8a) as well as the US house price index data taken from the US Federal Housing Finance Agency (FHFA) (Figure 8b) produce IRFs showing a clear double-dip behavior. Same is true for the responses of real residential property prices (Figure 8c) and the average sales price for houses sold in the US (Figure 8d). Monthly values of the quarterly real residential property prices and average sales price for houses sold are obtained by interpolation with the Case/Shiller real house price index, similar as seen in Rahal (2016). The interpolation follows the Chow & Lin (1971) procedure.



(a) Nominal Case/Shiller house price index

(b) FHFA house price index

(c) Real Property Prices

(d) Average Sales Price of Houses Sold

Figure 8: Impulse response functions of housing prices to monetary policy shock. Time (horizontal axis) is in months. The solid black lines represent the IRFs. The dashed lines indicate one standard deviation confidence interval.

# Appendix D IRF: Poor man's approach - same sign

Figure 9 shows the response of housing prices to an interest rate shock, where only interest rate surprises that are accompanied by the same sign deposit changes in the week following FOMC meetings are considered.

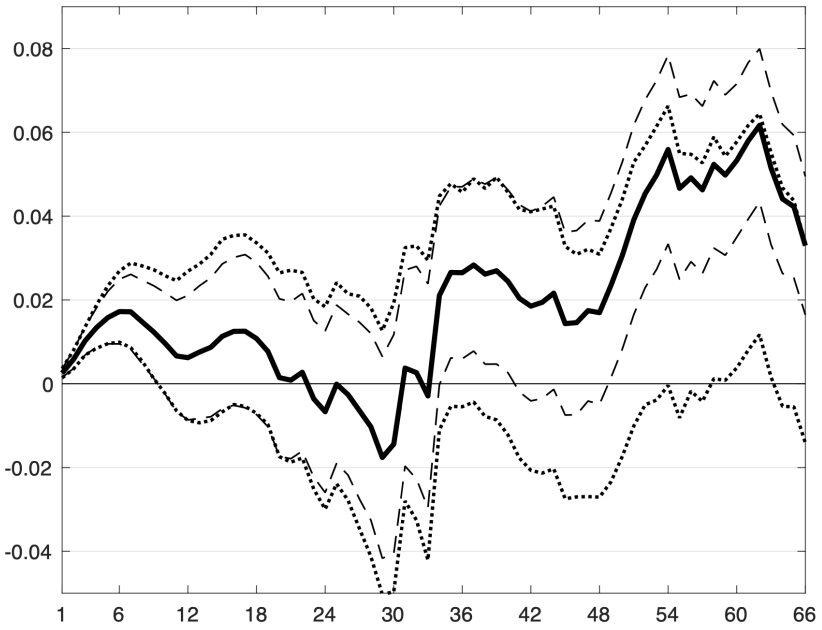



Figure 9: Impulse response function of housing prices to monetary policy shocks not transmitted through housing deposit channel. Time (horizontal axis) is in months. The solid black line represents the IRF. The dashed and dotted lines indicate one standard deviation delta-method and percentile-t bootstrap confidence interval, respectively.

# Appendix E Housing price equilibrium (step-by-step)

## E.1 Housing price (lender side)

The housing price $p^h$, emerging from the lender side, in equation 16 is derived from equation 10 and 11. Equation 10 can be solved for $\Lambda^l_{t+1}$, so that,

$$\Lambda^l_{t+1} = (1 - \frac{\chi^d_l}{\lambda^l_t} d_t^{-1}) \frac{\pi_{t+1}}{r^d_t}. \tag{E1}$$

Substituting $\Lambda^l_{t+1}$ in equation 11 implies at the steady state:

$$p^h_t = \frac{\chi_h}{\lambda^l} (h^l)^{-1} + \left[ (1 - \frac{\chi^d_l}{\lambda^l} d^{-1}) \frac{\pi}{r^d} \right] p^h (1 - \delta_h). \tag{E2}$$

Subtracting the term $\{ \left[ (1 - \frac{\chi^d_l}{\lambda^l} d^{-1}) \frac{\pi}{r^d} \right] p^h (1 - \delta_h) \}$ from the right-hand-side and factoring out $p^h$, one obtains:

$$\frac{\chi^h_l}{\lambda^l} (h^l)^{-1} = p^h (1 - (1 - \delta_h)(1 - \frac{\chi^d_l}{\lambda^l} d^{-1}) \frac{1}{r^d}). \tag{E3}$$

By solving for $p^h$, it emerges the housing price from the lender side, as seen in equation 16:

$$p^h = \frac{\chi_l^h}{\lambda^l(1-(1-\delta_h)(1-\frac{\chi_l^d}{\lambda^l}d^{-1})\frac{1}{r^d})h^l}. \tag{E4}$$

### E.2 Housing price (borrower side)

The housing price $p^h$, emerging from the borrower side, in equation 17 is derived from equation 14 and , assuming $\theta = 1$. First, equation 15 is solved for $\frac{\lambda_t^m}{\lambda_t^b}$, so that,

$$\frac{\lambda_t^m}{\lambda_t^b} = 1 - \Lambda_{t+1}^b \frac{r_t^m}{\pi_{t+1}}. \tag{E5}$$

Substituting then $\frac{\lambda_t^m}{\lambda_t^b}$ in 14 implies at steady state:

$$p^h - \left[1 - \Lambda^b r^m\right] p^h = \frac{\chi_b^h}{\lambda^b}(h^b)^{-1} + \Lambda^b(1-\delta_h)p^h. \tag{E6}$$

Factoring out $p^h$ at the left-hand-side implies,

$$p^h \Lambda^b r^m = \frac{\chi_b^h}{\lambda_t^b}(h_t^b)^{-1} + \Lambda^b(1-\delta_h)p^h; \tag{E7}$$

$$\frac{\chi_b^h}{\lambda^b}(h^b)^{-1} = p^h \Lambda^b r^m - \Lambda^b(1-\delta_h)p^h; \tag{E8}$$

$$\frac{\chi_b^h}{\lambda^b}(h^b)^{-1} = p^h \beta_b(r^m - (1-\delta_h)). \tag{E9}$$

Solving for $p^h$ gives the housing price emerging from the borrower side, as seen in equation 17:

$$p^h = \frac{\chi_b^h}{\lambda^b \beta_b(r^m - (1-\delta_h))h^b}. \tag{E10}$$

### E.3 Housing price (equilibrium)

Combining equation 16 and 17, since, in the equilibrium, $p_l^h = p_b^h$, implies:

$$\frac{\chi_b^h}{\lambda^b(\beta_b(r^m-1)h^b} = \frac{\chi_l^h}{\lambda^l(1-(1-\delta_h)(1-\frac{\chi_l^d}{\lambda^l}d^{-1})\frac{1}{r^d})h^l}. \tag{E11}$$

Following equation E1 and the first-order condition 13 of the lender side, $\beta_l = \Lambda^l = (1-\frac{\chi_l^d}{\lambda^l}d^{-1})\frac{1}{r^d}$. Together with setting $\delta_h = 0$ (for the sake of simplicity) and factoring in $\beta_b$, this implies:

$$\frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)h^b} = \frac{\chi_l^h}{\lambda^l(1-\beta_l)h^l}. \tag{E12}$$

Following Iacoviello (2015), the housing supply is assumed constant to one, so that $h_t^l + h_t^b =$

1. Consequently, by definition, $h^l = 1 - h^b$, which implies,

$$\frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)h^b} = \frac{\chi_l^h}{\lambda^l(1-\beta_l)(1-h^b)}. \tag{E13}$$

When solving for $h^b$, one multiplies by $(1 - h^b)$, so that,

$$\frac{(1-h^b)}{h^b}\frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)} = \frac{\chi_l^h}{\lambda^l(1-\beta_l)}; \tag{E14}$$

$$(\frac{1}{h^b} - 1)\frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)} = \frac{\chi_l^h}{\lambda^l(1-\beta_l)}; \tag{E15}$$

$$\frac{1}{h^b}\frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)} - \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)} = \frac{\chi_l^h}{\lambda^l(1-\beta_l).} \tag{E16}$$

From the latter equation one can easily find $h^b$ (as seen in equation 18):

$$h^b = \frac{\frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}}{\frac{\chi_l^h}{\lambda^l(1-\beta_l)} + \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}}. \tag{E17}$$

Substituting $h^b$ in equation 17 reveals the housing price dynamics at the equilibrium:

$$p^h = \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}\frac{\frac{\chi_l^h}{\lambda^l(1-\beta_l)} + \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}}{\frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}}; \tag{E18}$$

$$p_E^h = \underbrace{\frac{\chi_l^h}{\lambda^l(1-\beta_l)}}_{\text{lender side}} + \underbrace{\frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}}_{\text{borrower side}}. \tag{E19}$$

# Appendix F Housing price (equilibrium) - with a non-constant housing supply

Let's suppose the housing supply is not constant, but rather a function of the interest rate, set by the central bank: $h^l + h^b = f(r^m)$. Raising interest rates can significantly impact housing construction by increasing the overall cost of construction projects. Higher interest rates make borrowing more expensive, leading to reduced investment in new housing developments. This, in turn, can exacerbate existing supply chain issues, driving up prices for building materials as demand decreases but costs rise. As construction becomes more expensive, fewer projects may be initiated, further tightening housing supply. Consequently, housing supply can be viewed as a function of interest rates, as fluctuations in rates directly influence the feasibility and cost of construction.

Equation 38 changes as follows:

$$\frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)h^b} = \frac{\chi_l^h}{\lambda^l(1-\beta_l)(f(r^m) - h^b)}. \tag{F1}$$

When solving for $h^b$, one multiplies by $1 - f(r^m)$, so that,

$$\frac{(f(r^m) - h^b))}{h^b} \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)} = \frac{\chi_l^h}{\lambda^l(1 - \beta_l)}; \tag{F2}$$

$$(\frac{f(r^m)}{h^b} - 1) \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)} = \frac{\chi_l^h}{\lambda^l(1 - \beta_l)}; \tag{F3}$$

$$\frac{f(r^m)}{h^b} \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)} - \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)} = \frac{\chi_l^h}{\lambda^l(1 - \beta_l).} \tag{F4}$$

From the latter equation one can easily find $h^b$ (as seen in equation 18):

$$h^b = f(r^m) \frac{\frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}}{\frac{\chi_l^h}{\lambda^l(1-\beta_l)} + \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}}. \tag{F5}$$

Substituting $h^b$ in equation 17 reveals the housing price dynamics at the equilibrium:

$$p^h = \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)} \frac{\frac{\chi_l^h}{\lambda^l(1-\beta_l)} + \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}}{f(r^m) \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}}; \tag{F6}$$

$$p_E^h = \frac{1}{f(r^m)} (\frac{\chi_l^h}{\lambda^l(1 - \beta_l)} + \frac{\chi_b^h}{\lambda^b(\beta_b r^m - \beta_b)}). \tag{F7}$$

Since $f(r^m)$ is always a positive function, increasing $f(r^m)$ only reduces the speed, not the direction of the price dynamics found in Section 6. Moreover, since $r^m$ is bounded, the assumption that $\frac{1}{f(r^m)}$, and thus housing supply, is constant, following Iacoviello (2015), is valid. Consequently, it is correct to focus only on the price movements resulting from the demand for housing, as we've done so far, since accounting for a non-constant housing supply does not significantly change the conclusions of this paper.