House Price Responses to Monetary Policy Surprises: Evidence from the U.S. Listings Data*

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Abstract

Existing literature documents that house prices respond to monetary policy surprises with a significant delay, taking years to reach their peak response. We present new evidence of a much faster response. We exploit information contained in listings for the residential properties for sale in the United States between 2001 and 2019 from the CoreLogic Multiple Listing Service Dataset. Using high-frequency measures of monetary policy shocks, we document that a one-standard-deviation contractionary monetary policy surprise lowers housing list prices by 0.2–0.3 percent within two weeks—a magnitude on par with the effect on stock prices. House prices respond stronger to the surprises to future rates as compared to the surprise changes in the federal funds rate. Sale prices are mostly pre-determined by list prices and do not independently respond to monetary policy surprises.

Keywords: House prices. Monetary policy. Transmission of monetary policy. List and sales prices.

JEL codes: E52, R21, R31.

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1 Introduction

Existing literature finds that house price growth is fairly smooth and appears to be insensitive to changes in real or nominal interest rates (Case and Shiller, 1989). This sluggishness is at odds with an almost immediate response of financial assets (Bernanke and Kuttner, 2005). Indeed, Case and Shiller conclude: "There is thus a profitable trading rule for persons who are free to time the purchase of their homes." The sluggishness of house price growth is broadly attributed to inefficiencies in real estate markets due to search costs, transaction and carrying costs, tax considerations, or non-rational behavior and animal spirits. It should then come as no surprise that academics and central bankers are constantly trying to understand the insensitivity of house prices to macroeconomic conditions and monetary policy. Summarizing the existing literature, Kuttner (2013) and Williams (2015) conclude that house prices react significantly to changes in interest rates, but these effects materialize only gradually, after two years or so.

In this paper, we use high-frequency data on house listings and present new evidence of a much faster response of house prices to monetary policy shocks. The main difference of our approach from the existing literature is that we distinguish between the price at which a house comes on the market, list price, and the price at which the house is sold, sale price. We demonstrate that list prices play a central role in housing market adjustment to monetary policy announcements.

Our housing price data come from the CoreLogic Multiple Listing Service Dataset, which contains property-level data on listings and sales of residential properties in the United States. Our sample covers the majority of residential property listings across 43 states and D.C. (about 11 thousand zip codes) from 2001 to 2019, at a daily frequency. We study the responses of house list prices, sale prices, and time-on-market to monetary policy surprises using the local projections method of Jordà (2005). We use measures of monetary surprises associated with Federal Open Market Committee (FOMC) announcements and identified by high-frequency methods in Nakamura and Steinsson (2018), Swanson (2021), and Bauer and Swanson (2022). We use Swanson (2021)'s shocks to distinguish between the unexpected changes in short-term interest rates and surprises to the future interest rates.

We find that house list prices move quickly after monetary policy announcements, responding mainly to surprises about future interest rates. A one-standard-deviation contractionary surprise to future interest rates—measured by Swanson (2021)'s factors for forward guidance (FG) and large-scale asset purchases (LSAP)—lowers list prices by roughly 0.2 and 0.3 percent, respectively, within 2–3 weeks after the shock. These responses are similar in magnitude to responses of financial assets on the day of an FOMC announcement, derived using the same surprise measures. House prices are less sensitive to the surprise changes in the federal funds rate. We obtain similar results to the measures of monetary surprises constructed by Nakamura and Steinsson (2018) and Bauer and Swanson (2022) that utilize information on a range of short-term interest rates up to one year maturity.

We find that the response of house prices to monetary surprises passes through list prices as opposed to sale prices. The estimated response of sale prices for houses listed after the announcement closely follows the response of list prices, whereas for houses listed before the announcement list and sale prices are insensitive to the surprise. Hence, the adjustment of list prices to the monetary surprise is almost entirely passed through to sale prices, suggesting list prices play a more important role in allocating house market transactions than previously thought.

We estimate that in response to a surprise increase in future interest rates, captured by Swanson's FG or LSAP factors, the 30-year fixed-rate mortgage rates increase by around 4 basis points within a month. Rising mortgage rates cool housing demand, especially from financially constrained households, and may explain the decline in house prices that we estimate. While the house financing channel has been extensively studied in the literature, the speed of adjustment that we uncover is a novel finding.[1] We calculate the implied semi-elasticity of list price responses to mortgage rates of around 5, which is in the range of *long-run* semi-elasticities, between 3 and 8, in the literature surveyed by Kuttner (2013).

To obtain a more direct evidence of the house financing mechanism, we conduct two additional exercises. In the first exercise, we estimate the response of list prices to an exogenous variation in mortgage rates using an instrumental variable approach with three Swanson (2021) factors as instruments. Our results suggest that an exogenous 1 percentage point increase in mortgage rates lowers list prices by around 3 percent within one month. This direct estimate of the semi-elasticity is statistically and economically significant, although somewhat lower than the semi-elasticity implied by the responses to monetary shocks. It is consistent with Adelino, Schoar and Severino (2012) who use exogenous variation in credit conditions to estimate the semi-elasticity of house prices to interest rates.

In the second exercise, we study heterogeneity of the house price response to monetary policy surprises along four dimensions: household income, house value, density of local

[1] See Section 3 for the literature that studies the role of housing financing in the transmission of interest rate changes to housing markets.

bank branches, and local housing supply elasticity. We find that list prices in zip codes with lower household incomes or lower house values are more sensitive to the slope surprises than prices in high-income or high-value zip codes. Intuitively, zip codes with lower incomes or house values have a higher fraction of financially constrained buyers and sellers, who are more sensitive to interest rate changes. Finally, we find house price responses do not vary with local housing supply elasticity, even after controlling for local household income level and bank branch density.

Our focus on list prices contrasts with most of the existing literature that relies on indexes based on sales prices. Studies using closing-date indexes find the negative relationship between nominal interest rates and house prices only at very long horizons (Kuttner, 2013; Williams, 2015). A smaller number of studies examine the behavior of list prices, usually in the context of a search and matching framework where list prices direct buyers' search efforts by providing a signal of house quality or seller's preference.[2] There is some evidence that house list prices exhibit stronger contemporaneous correlations with equity prices and macroeconomic news shocks than conventional closing-date prices (Anenberg and Laufer, 2017). Nonetheless, evidence of a direct link between house prices and interest rates at short horizons has been elusive. Therefore, our main contribution is to demonstrate that house prices respond to monetary shocks much sooner than previously found. Prompt responses of house prices are, therefore, much more akin to the responses of prices of financial assets than prices of consumer goods and services. We show that house market outcomes after a monetary surprise are associated primarily with adjustment of list prices rather than closing-date prices.

The rest of the paper is structured as follows. Section 2 describes our property-level data and construction of house price indexes. Section 3 provides a theoretical primer on the impact of monetary policy shocks on house prices, discusses the monetary policy shocks series, and describes the estimation method. Section 4 presents the main results. Section 5 provides evidence for and the discussion of the mechanisms underlying the results. Section 6 provides the robustness results. Section 7 concludes and lays out avenues for future research.

[2] Early references include Wheaton (1990), Horowitz (1992), Yavas and Yang (1995). Studies that evaluate such models against the house price data include Head, Lloyd-Ellis and Sun (2014); Guren (2018); Anenberg (2016); Anenberg and Ringo (Forthcoming).

2 Data and Construction of House Price Indexes

2.1 Property-level house price data

We use data on home listings and sales from the CoreLogic Multiple Listing Service Dataset. The dataset contains detailed information on the universe of all housing units, commercial properties, rentals and land plots listed for sale or for rent on the Multiple Listing Service (MLS) platforms across 43 states and the District of Columbia in the United States.[3] The starting date for the MLS data varies by state, with many states reporting consistently since the late 1980s. To include as many locations as possible, we use data from 2001 to 2019 in our analysis.

The unit of observation in the data is a property listing, which comes at a daily frequency. For each listed property, we observe the list price, the date at which this property was listed, and all the characteristics of the property that the owner and the real estate broker used to describe the listing (e.g., address, property type, number of bedrooms and bathrooms, year built, living area, etc.). We also observe the dates at which the property was sold and the sale price. Our raw dataset contains about 115 million listings.

We focus on the listings of single-family homes and apartments and exclude listings of rentals, land parcels, mobile homes or commercial properties. We restrict our sample to properties that have a non-missing and non-zero list price and drop listings that have no information on the city where the property is located. We drop listings for properties that do not sell during our sample period so that we have both list and sale price observations for each listing. Applying these filters reduces observations by about 17%. Listings of rentals, land parcels, mobile homes and commercial properties account for the majority of observations that are filtered out. To ensure that our results are not driven by outliers, we winsorize the top and bottom 0.1% of list and sale prices by zip code in our sample. We also drop properties that are listed or sold in location-weeks that have fewer than 5 observations per location-week. Lastly, we drop zip codes that have fewer than 50 observations across all weeks in our sample.

Table 1 provides some summary statistics. Our cleaned data comprise a total of 92,064,327 listings, on average over 760 weeks per zip code, and 121,101 listings per week across zip codes. The data represent 10,958 zip codes across 43 U.S. states and D.C., with a median of 8 new listings per week per zip code. A typical house is listed at around 228 thousand 2010 dollars, and after spending a median of 110 days on the market is sold at a slightly lower price of 212 thousand dollars. There is substantial heterogeneity in prices

[3] Some states have very few observations in the Corelogic data or are not reported at all: Alaska, Maine, North Dakota, South Dakota, Utah, Vermont, Wyoming.

and time-on-market both within and across zip codes.

	Median	Inter-quartile range	
		within zip	across zip
List price, in 2010 dollars	228,965	127,749	188,771
Sale price, in 2010 dollars	212,492	94,608	172,784
Time-on-market, days	110	93	46
Number of listings per week	8		7
Total zip codes = 10,958			
Total listings/week = 121,101			
Total listings = 92,064,327			

Table 1. Summary Statistics For House Listings Across U.S. Zip Codes

Notes: Table provides statistics at zip code level from 2001 to 2019. List price and sale price (by list date) are provided in 2010 dollars. Column "Median" provides a weighted median across zip-level medians, where the weights are the number of listings in a zip code in a given week. Column "Inter-quartuile range, within zip" provides weighted median across zip-level inter-quartile ranges. Column "Inter-quartile range, across zip" provides weighted inter-quartile range across zip-level medians.

2.2 Construction of weekly house price indexes by zip code

In the first stage of our empirical analysis, we estimate real house price indexes by zip code using hedonic regressions. We deflate house prices using the Consumer Price Index for All Urban Consumers that is based on all items in the United States.

Let $\phi^L_{il,d}$ denote the real list price for property i in zip code l listed on date d, and let $w(d)$ denote the calendar week containing date d. For each location l we estimate:

$$\ln \phi^L_{il,d} = \chi_l + p^L_{l,w(d)} + \mathbb{X}_{il,d} + \omega_{l,w(d)} + \varepsilon_{il,d}, \tag{1}$$

where χ_l is a constant; $p^L_{l,w(d)}$ is the time fixed effect (at weekly frequency); $\mathbb{X}_{il,d}$ is a set of housing characteristics for property i that includes the construction year of the property, total number of bedrooms, classification of land-use (e.g., apartment, townhouse, single-family residence, etc.), and the size of the living area in square feet; $\omega_{l,w(d)}$ is the week-of-year effect to capture seasonality in the housing markets across the 52 weeks of a year; and $\varepsilon_{il,d}$ is the error term. The weekly log list price index for zip code l is given by the estimates of the weekly time effects, $p^L_{l,w(d)}$.

We estimate a similar regression to construct weekly log time-on-market by list date indexes, $tom_{l,w(d)}$.

For sale prices, we construct two indexes—one defined by list date and one defined by closing date. Let d denote the date of the listing, as before, and x denote the closing date. The sale-by-list-date index uses the date of the listing, i.e., the same date for which we construct the list price index. Denoting by $\phi^S_{il,d}$ the real sale price for property i in zip code l listed on date d, we estimate for each zip code l a specification similar to the one in (1):

$$\ln \phi^S_{il,d} = \chi_l + p^{SL}_{l,w(d)} + \mathbb{X}_{i,d} + \omega_{l,w(d)} + \varepsilon_{il,d}. \tag{2}$$

The weekly log sale-by-list-date price index for zip code l is given by the estimated location-week time effects, $p^{SL}_{l,w(d)}$.

Denote by $\phi^S_{il,x}$ the sale price for property i in zip code l sold on date x. We estimate for each location l:

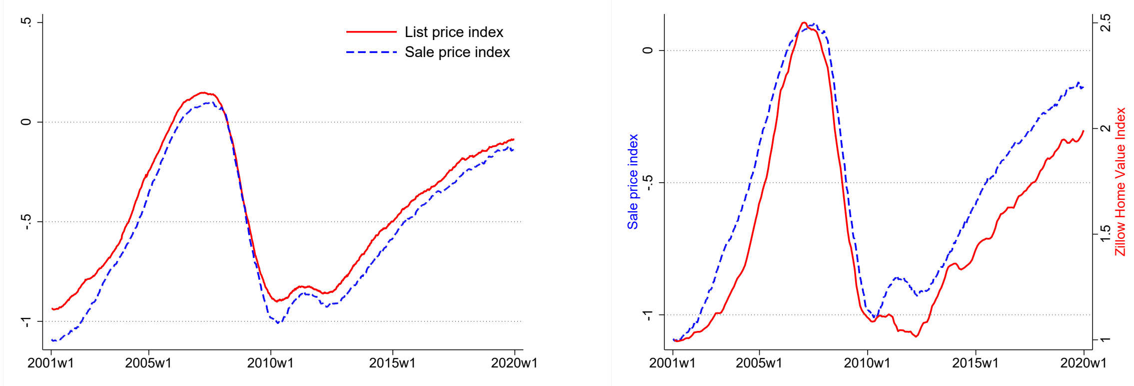
$$\ln \phi^S_{il,x} = \chi_l + p^{SS}_{l,w(x)} + \mathbb{X}_{i,x} + \omega_{l,w(x)} + \varepsilon_{il,x}, \tag{3}$$

The weekly log sale-by-sale-date price index for zip code l is given by the estimated location-week time effects, $p^{SS}_{l,w(x)}$.

The sale-by-list-date price index, $p^{SL}_{l,w(d)}$, is constructed using data on the same properties as the list price index $p^L_{l,w(d)}$. Therefore, it provides the average price at which houses listed in week w are eventually sold at some point later. The sale-by-sale-date price index $p^{SS}_{l,w(x)}$ provides the average sale price of all houses sold in week w for the set of houses listed at various points earlier. In our main analysis, we will examine whether the timing of the listing influences the response of sale prices to monetary policy surprises by comparing price indexes for properties sold after the shock versus properties *listed* after the shock.

To illustrate the constructed indexes, in Figure 1 we show the estimated list and sale-by-sale-date price indexes for the most populated zip code in the city of Los Angeles. We smooth the series by a 52-week moving average to facilitate comparisons. Panel (a) shows that both price indexes increased dramatically during the housing boom prior to 2007, declined sharply during the 2007–09 crisis and until about 2012, and gradually recovered thereafter. For the most part, the sale-by-sale-date price index lags the list price index. Panel (b) shows that the sale-by-sale-date price index aligns well with Zillow Home Value Index for this zip code.[4]

[4] Methodology for Zillow Home Value Index can be found at https://www.zillow.com/research/zhvi-methodology-2019-deep-26226/.



(a) List price and sales price indexes

(b) Sale price and Zillow Home Value Indexes

Figure 1. House Price Indexes for zip code 90011 (Los Angeles)

Notes: Panel (a) shows the list price index and the sale-by-sale-date price index estimated based on the hedonic regressions described in equations (1) and (3) for zip code 90011, the zip code with the highest number of inhabitants in Los Angeles. Both indexes are smoothed using a 52-week backward-looking moving average. Panel (b) shows the smoothed sale-by-sale-date price index and the Zillow Home Value Index for the same location.

3 Methodology of Estimating the Effect of Monetary Surprises on House Prices

In this section, we first provide a theoretical primer on the impact of monetary policy shocks on house prices. We then explain the measures of monetary policy surprises and describe our estimation method.

3.1 Theoretical primer

The basic mechanism linking unexpected changes in nominal interest rates to house prices can be explained using a standard business cycle model with households deriving utility from non-durable and housing consumption. For brevity, we only mention the key equations (details can be found, for example, in Barsky, House and Kimball (2007)).

When a central bank raises interest rates and the raise is unanticipated, the demand shifts from current consumption to future consumption. This can be illustrated via a standard Euler equation:

$$U_t^c = \beta(1 + i_t)E_t[U_{t+1}^c P_t^c / P_{t+1}^c],$$

where U_t^c is the marginal utility of non-durable consumption, P_t^c is its price, and i_t is the risk-free rate. The unexpected increase in the interest rate in period t is met with higher marginal utility (lower consumption) and lower prices in period t.

The effect on house prices depends on how interest rates affect the value of housing γ_t^h, given by the net present value of future real returns from housing services $\{r_{t+\tau}^h\}$:

$$\gamma_t^h = \beta E_t[\gamma_{t+1}^h(1 + r_{t+1}^h)].$$

Because houses are illiquid long-lived durables, their value is insensitive to temporary factors, such as central bank's interest rate changes, i.e., $\gamma_t^h \approx \gamma^h$. This means households have much less incentive to smooth house spending relative to spending on non-durables, which in turn has direct implication for how house prices respond to interest rate shocks.

Namely, households allocate their non-durable consumption and housing to equate the marginal values of an additional dollar spent so that $U_t^c / P_t^c = \gamma^h / P_t^h$. Since the nominal interest rate increase raises U_t^c / P_t^c, it also lowers house prices P_t^h. And since house spending is lumpier than non-durable consumption, house prices fall by more than non-durable prices. Indeed, using the Euler equation and constant value of housing, we can obtain the equation linking the risk-free rate and housing prices:

$$(1 + i_t) \approx \beta^{-1} E_t[P_{t+1}^h / P_t^h].$$

In response to an unexpected nominal interest rate hike house prices must fall, and their fall is larger than for non-durable consumption prices.

Because buying or selling a house often involves financing, interest rates also influence house prices via the cost of housing debt. Higher interest rates raise mortgage rates and reduce availability of credit, cooling housing demand, especially from financially constrained households. A large literature that studies the role of housing finance on the transmission of interest rate changes to housing markets includes Iacoviello (2005); Favara and Imbs (2015); Anenberg and Kung (2017); Garriga, Kydland and Šustek (2017); Greenwald (2018); Garriga, Manuelli and Peralta-Alva (2019); Bhutta and Ringo (2021); Berger, Milbradt, Tourre and Vavra (2021); Wong (2021); Eichenbaum, Rebelo and Wong (2022). Davis and Van Nieuwerburgh (2015) provide a review of the macroeconomic aspects of the housing finance literature.

Besides cooling the demand for housing services, higher current and future interest rates also reflect the risk premium associated with owning a house (Campbell, Davis, Gallin and Martin, 2009; Favilukis, Ludvigson and Van Nieuwerburgh, 2017). In addition, higher interest rates can increase the user cost of housing indirectly by raising expectation of house price depreciation (Glaeser, Gottlieb and Gyourko, 2013; Kuchler, Piazzesi and Stroebel, 2022) or changing property or income tax obligations across different homeowners (Poterba, 1984). A more general review of the literature on housing in macroeconomics is provided by Piazzesi and Schneider (2016).

3.2 Measures of monetary policy surprises

In the analysis, we use measures of high-frequency monetary policy surprises from the existing literature.[5] As described in Section 3.1, we are interested not only in the shocks to the level of the interest rate but also to the slope of the future rate changes.

In our baseline analysis, we use monetary policy surprises from Swanson (2021). Swanson applies a factor model to assets with maturities below one year and 2-, 5- and 10-year Treasury yields. He estimates top three factors, which explain 94% of the changes in these interest rate responses within a 30-minute window around scheduled FOMC announcements between 1991 and 2019.[6] The identifying assumption in such an approach is that the shock measures only reflect information contained in the FOMC announcement. Swanson shows that under additional restrictions the first factor can be related to unexpected changes in the federal funds rate, the second factor—to Fed's forward guidance, and the third factor—to Large-Scale Asset Purchases (LSAPs).

We broaden our set of measures of monetary policy shocks by including the series from Nakamura and Steinsson (2018) and Bauer and Swanson (2022). Nakamura and Steinsson (2018) measure the monetary policy shock as the first principle component of the unanticipated change in five short-term interest rates within a narrow 30-minute window around scheduled FOMC announcements. Because this measure uses interest rates at maturities within one year, it captures the effects of changes in both the current federal funds rate and expected future federal funds rates. The latter is influenced by the Fed's forward guidance and balance sheet policies. Nakamura and Steinsson (2018) show that the effects of their monetary policy shocks reflect, in addition to surprises to the path of interest rates, the effects of Fed announcements on private sector economic expectations. Nakamura and Steinsson (2018) construct the shocks from February 2000 to March 2014.

[5] High-frequency identification is based on the lumpy adjustments of bond prices within narrow window around scheduled Federal Open Market Committee (FOMC) announcements (Cook and Hahn, 1989; Kuttner, 2001; Cochrane and Piazzesi, 2002; Gürkaynak, Sack and Swanson, 2005). Recent studies using high-frequency identification of monetary shocks include Hanson and Stein (2015); Gertler and Karadi (2015); Nakamura and Steinsson (2018); Altavilla, Brugnolini, Gürkaynak, Motto and Ragusa (2019); Cieslak and Schrimpf (2019); Gurkaynak, Kara, Kısacıkoğlu and Lee (2021); Swanson (2021); Andrade and Ferroni (2021).

[6] Swanson normalizes the federal funds rate factor to have a unit standard deviation from July 1991 to December 2008, the LSAP factor to have a unit standard deviation over the ZLB period, from January 2009 to October 2015, and the forward guidance factor to have a unit standard deviation from July 1991 to June 2019.

In our analysis, we use the series from Acosta and Saia (2022) who extend Nakamura and Steinsson's shocks through 2019.

Bauer and Swanson (2022), like Nakamura and Steinsson (2018), use a range of fed funds and Eurodollar futures contracts to construct high-frequency measures of monetary policy shocks. Bauer and Swanson refine their measure along two dimensions. They add information from the Fed Chair's speeches and testimonies and purge their measure of the component predicted by macroeconomic and financial data preceding the policy announcement.

The advantage of using high-frequency shock measures is that estimation of their effects requires relatively weak identifying assumptions, as long as the dependent variable is also measured at a relatively high frequency. Because we measure house prices at weekly frequency, we assume that monetary shocks are orthogonal to other variables that influence house prices in weeks after FOMC announcements.

3.3 Estimation method

Let $\mathrm{P}_{l,t}$ denote a house market index of list price, sale-by-list-date price, sale-by-sale-date price, and time-on-market in zip code l on weekly date t. Let S_t denote a measure of an identified monetary policy surprise in week t. In our baseline specification, S_t is a vector of Swanson (2021)'s three factors—FFR_t, FG_t, LSAP_t—which capture the surprise changes in the federal funds rate, forward guidance and large-scale asset purchases, respectively.

We employ Jordà (2005)'s local projections method to estimate the average effect of a monetary policy shock during FOMC week t on house market indexes $h = 0, ..., H$ weeks after the shock. For each horizon h, we estimate the following empirical specification:

$$\ln P_{l,t+h} - \ln P_{l,t-1} = \alpha^{(h)} + \beta^{(h)}\mathrm{S}_t + \sum_{q=1}^{52} \theta_q^{(h)} (\ln P_{l,t-q} - \ln P_{l,t-q-1}) + \chi_l^{(h)} + \varepsilon_{l,t}^{(h)}, \quad (4)$$

where $\ln P_{l,t+h} - \ln P_{l,t-1}$ is the change in a house market index over h weeks after the shock in week t. We use our estimates of $p^{L}_{l,w(d)}$, $p^{SL}_{l,w(d)}$, $p^{SS}_{l,w(x)}$ and $tom_{l,w(d)}$ as measures of $\ln P_{l,t}$. Controls are 52 lags of the weekly change in the index variable and zip code fixed effects, $\chi_l^{(h)}$. $\varepsilon_{l,t}^{(h)}$ is the error term, assumed to be heteroskedastic, independent across localities l, and serially correlated. Note that all coefficients are h-horizon-specific. We estimate equation (4) by fixed-effects panel regression method with Driscoll and Kraay (1998) standard errors. In the estimation, we weigh each observation by the number of listings in a given week in each zip code.

Our main coefficients of interest are $\beta^{(h)}$, $h = 0, ..., H$, which measure the response of a housing market index h weeks after the monetary policy surprise.[7] The identifying assumption is that the monetary shock in t is orthogonal to other variables that influence the change in the housing index over the h-week horizon after the shock.

The linear projections method of estimating impulse responses has several advantages over a vector autoregression or other time-series methods applied to aggregate house prices (Hamilton, 2008). First, we use existing series on monetary policy surprises estimated using high-frequency approach. The series are plausibly exogenous and, therefore, we can use them directly in our analysis bypassing endogeneity concerns. Second, since our house price data are high-frequency, we estimate the responses within weeks after a monetary shock, as opposed to months or quarters as is typically done with aggregate house price indexes.

4 House Price Responses to Monetary Policy Shocks

4.1 Baseline responses

Figure 2 shows the estimated responses of list prices to contractionary one-standard-deviation impulses for three monetary policy shocks estimated by Swanson (2021).[8] We find that the responses to factors associated with surprises to future interest rates—FG and LSAP factors—are negative and significant: house prices fall by 0.2 and 0.3 percent respectively within 2–3 weeks after the shock. By contrast, the responses to the surprise change in the federal funds rate—FFR factor—is positive but not statistically significant.

The magnitudes of the list price responses are comparable to the responses of financial assets on the day of announcement (see Tables 5 and 6 in Swanson (2021)). The key difference is that financial assets tend to respond strongly on the day of the announcement and their responses tend to be larger to the level surprise than to the forward guidance or LSAP surprises. For example, Swanson reports that the S&P500 stock index responds by –0.37, –0.14, and 0.03 percent to FFR, FG and LSAP shocks, respectively, and the first two responses are statistically significant. By contrast, house price responses take a couple of weeks to reach similar magnitudes, and the responses to slope factors are statistically significant, whereas the response to the level surprise is at best weakly significant. This evidence demonstrates that house prices respond to monetary shocks much faster than previously thought, and these responses are roughly on par with responses of financial

[7] In Section 6.2 we extend the estimation of responses to horizons prior to a policy announcement.

[8] The contractionary impulse is an increase in the FFR and FG factors and a decrease in the LSAP factor.

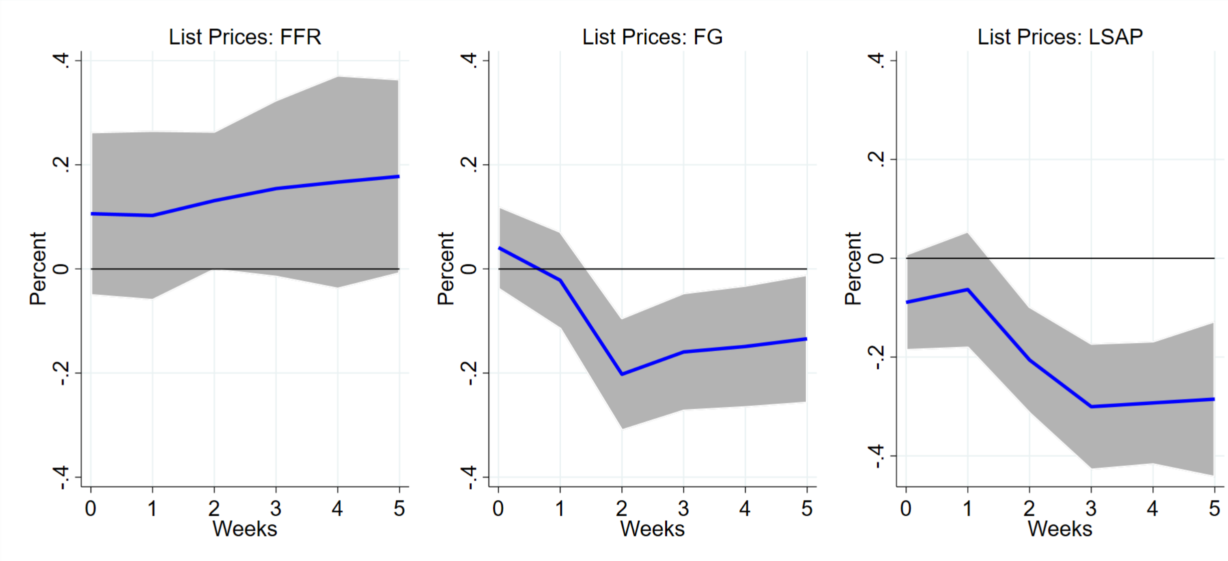


Figure 2. Responses of Housing List Prices to Contractionary Monetary Policy Surprises

Notes: The figure shows responses of the list price index to a one-standard-deviation increase in Swanson (2021)'s federal funds rate factor (left), forward guidance factor (middle), and (negative of) LSAP factor (right). Responses are estimated using specification in (4). Shaded areas represent the 90 percent confidence intervals based on Driscoll-Kraay standard errors.

assets on the day of announcement.

Somewhat counter-intuitive responses to the FFR factor reflect, in part, variation of house price dynamics over time. To see this, we estimate equation (4) separately for two sub-samples: from 2001 to 2008, and from 2009 to 2019. Figure 3 shows the responses to the FFR and FG factors. (We omit responses to the LSAP factor because the factor and its variation were negligible for the 2001–2008 period).

While the responses to FG factor are negative and significant for both sub-periods, the responses to FFR factor during the two periods differ. The response during the 2009–2019 is negative and significant. By contrast, the response during the 2001–2008 period is positive and not statistically significant.

Insensitivity of house prices to surprises in the short-term policy rate during 2001–2008 can be associated with the concurrent housing market boom. Credit market conditions were considerably slack during that period, counteracting the effect of rising interest rates by the Fed (Favilukis et al., 2017; Vojtech, Kay and Driscoll, 2018; Drechsler, Savov and Schnabl, 2022). Moreover, during a housing boom, house prices may be influenced by non-fundamental forces, such as expectations of house price appreciation, adding a considerable momentum to house price growth and desensitizing it to monetary policy surprises (Kuchler et al., 2022).

These results illustrate why it may have been difficult to detect high-frequency house

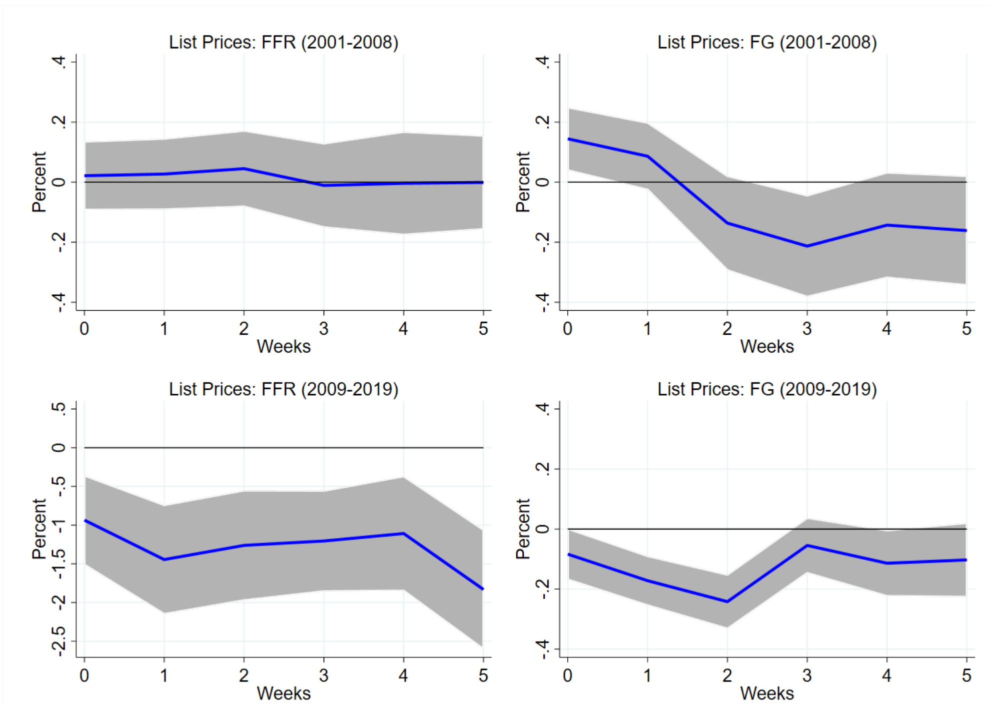
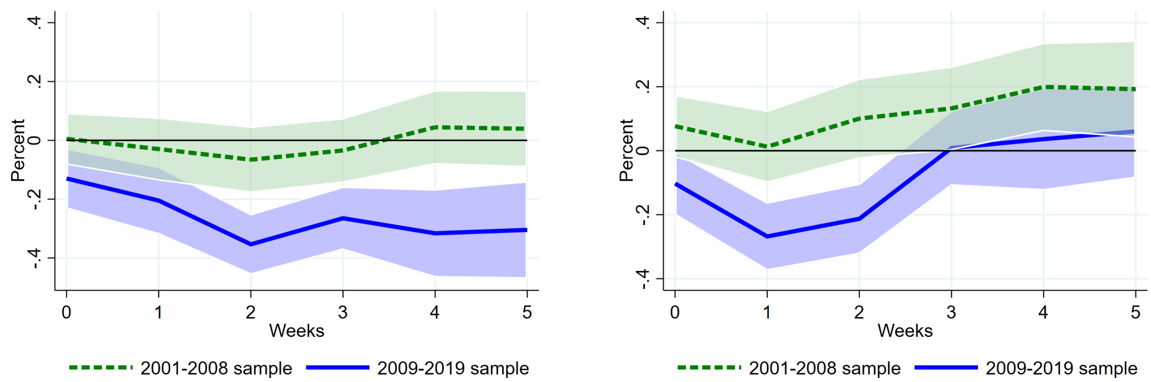


Figure 3. Responses of Housing List Prices to Contractionary Monetary Policy Surprises, by Sub-Period

Notes: The figure shows responses of the list price index to a positive one-standard-deviation impulse to Swanson (2021)'s federal funds rate factor (left) and forward guidance factor (right). Responses are estimated using specification in (4) separately for two sub-samples: from 2001 to 2008 (top) and from 2009 to 2019 (bottom). Shaded areas represent the 90 percent confidence intervals based on Driscoll-Kraay standard errors.

price responses so far in the literature (Williams, 2015). Much of the 2001–2008 data are dominated by the housing boom which tends to blunt the effect of interest rates on house prices. Our long sample helps circumvent this issue. Furthermore, it might be difficult to measure a house price response using broader measures of monetary policy shocks, because house prices are more responsive to surprises in the slope of the yield curve and not the level. Recent advances in the high-frequency identification of monetary policy shocks have brought about measures that distinguish between current and future interest rate surprises.

Our results remain robust to using alternative monetary policy shock measures. We estimate specification (4) using two alternative measures—by Bauer and Swanson (2022) and Nakamura and Steinsson (2018). We standardize both shocks to our sample and split the responses by two sub-periods. Figure 4 shows that the results remain broadly the same. Because these measures are constructed using a range of fed funds and Eurodollar futures with up to one year maturity, Nakamura and Steinsson's responses are somewhat weaker than the responses to slope factors in our baseline results. Nonetheless, using the shock measure from Bauer and Swanson, we find the results similar to our benchmark results, e.g., we find insignificant response during the 2001–2008 period and negative response (around –0.2 percent) in the 2009–2019 period.



(a) Bauer-Swanson impulse

(b) Nakamura-Steinsson impulse

Figure 4. Responses of Housing List Prices to Contractionary Monetary Policy Surprises, Bauer and Swanson (2022) and Nakamura and Steinsson (2018) Shocks

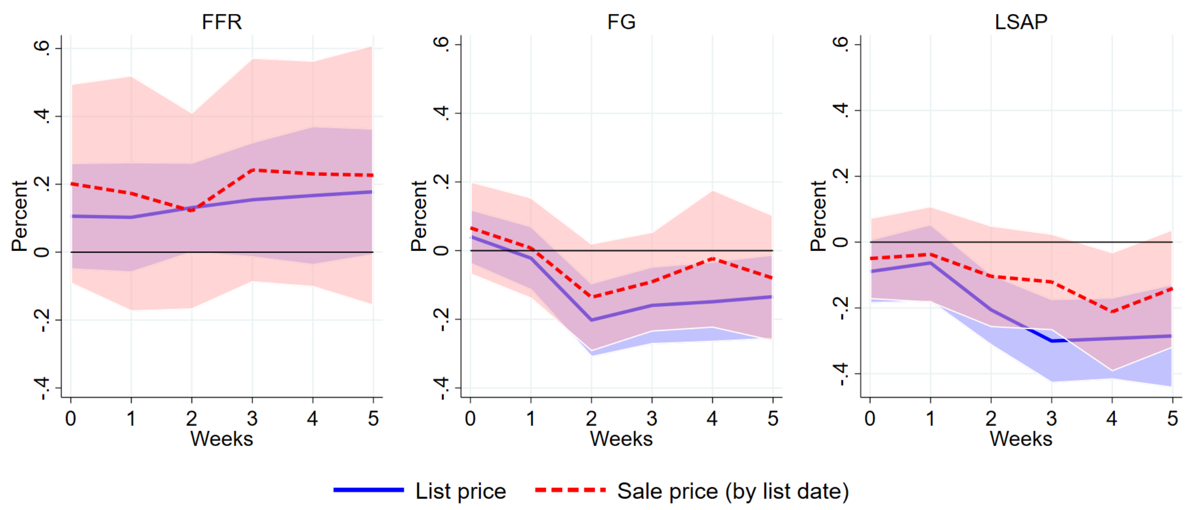
Notes: The figure shows list price responses to a positive one-standard-deviation shock by Bauer and Swanson (2022) and Nakamura and Steinsson (2018). Responses are estimated using specification in (4), separately for two sub-samples: from 2001 to 2008 (green dashed line) and from 2009 to 2019 (blue solid line). Shaded areas represent the 90 percent confidence intervals based on Driscoll-Kraay standard errors.

4.2 Sale prices and time-on-market

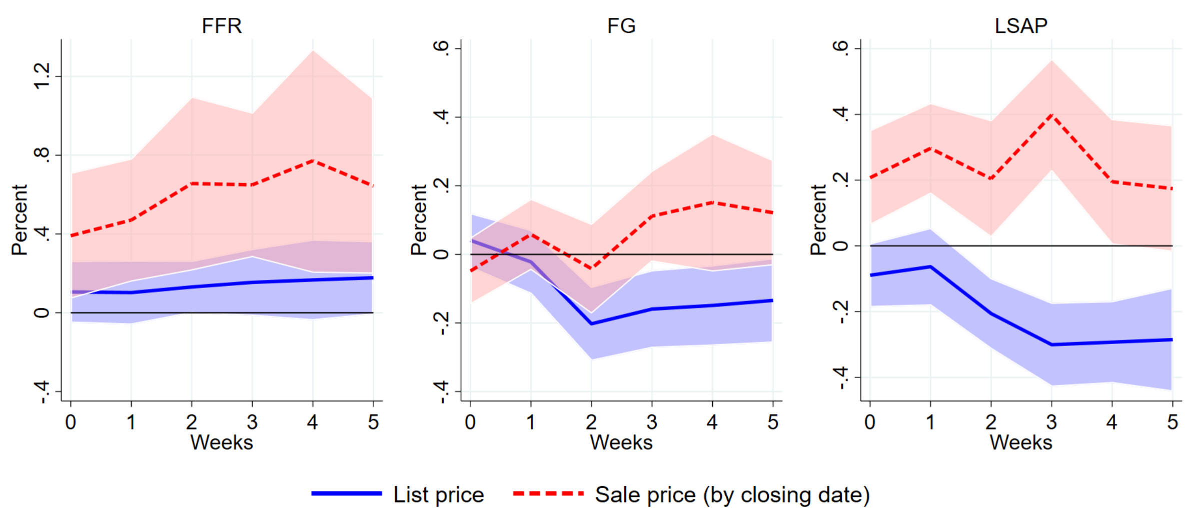
Our baseline results demonstrate that in response to monetary policy tightening surprises, home sellers lower their listing prices within a couple of weeks after monetary policy announcements. Do sale prices respond to monetary surprises over and beyond the response already built in the list prices? Do lower list prices following a monetary tightening significantly change the time houses remain on the market?

To answer these questions, we estimate the same specification as in (4), replacing the dependent variable with the a sale or time-on-market index. Figure 5 provides the responses of sale prices. The top panel shows the responses of sale-by-list-date prices and, for comparison, the responses of list prices from Figure 2, i.e., both indexes are constructed using the same set of housing properties. The responses of sale-by-list-date prices closely follow the responses of list prices, suggesting that the responses of sale prices mostly reflect earlier list price adjustments by sellers reacting to the monetary policy change.

The bottom panel in Figure 5 shows the responses of the sale-by-sale-date price index.



(a) List price vs Sale price (by list date).



(b) List price vs Sale price (by closing date).

Figure 5. Responses of Sale Prices.

Notes: The figure provides responses of the list price index from Figure 2 (blue line) and sale price indexes (red dashed line) to a positive one-standard-deviation impulse to Swanson (2021)'s fed funds factor (left), forward guidance factor (middle), and (negative of) LSAP factor (right). Responses are estimated using specification (4) sale price index by list date (top) and by closing date (bottom). Shaded areas represent the 90 percent confidence intervals based on Driscoll-Kraay standard errors.

The responses of this sale price index to slope factors are close to zero for FG factor, and they are positive for FFR and LSAP factors. The sale-by-sale-date index includes properties listed at different times: some are sold quickly and others may have sat on the market for a while. Since on average houses take almost four months to sell (as we report in

Section 2.2), responses of the sale-by-sale-date index reflect a bulk of houses that were listed before the policy announcement. Their list prices and subsequent sale prices tend to be higher than prices for houses listed after the contractionary surprise. Such a mix of listings in the sale-by-sale-date index creates an upward bias in its response to a contractionary monetary surprise and may obscure the estimation of short-run house price responses using conventional closing-date price indexes.

Our findings are consistent with results in Anenberg and Laufer (2017) who use listings data for nine major U.S. metro areas from 2008 to 2012 to construct a Case-Schiller-style list price index. They show that their list price index implies a stronger contemporaneous correlation with equity prices and macroeconomic news shocks, whereas the conventional closing-dated index does not. Our evidence indicates that Anenberg and Laufer's findings can be explained by a tighter link of list prices to monetary surprises.

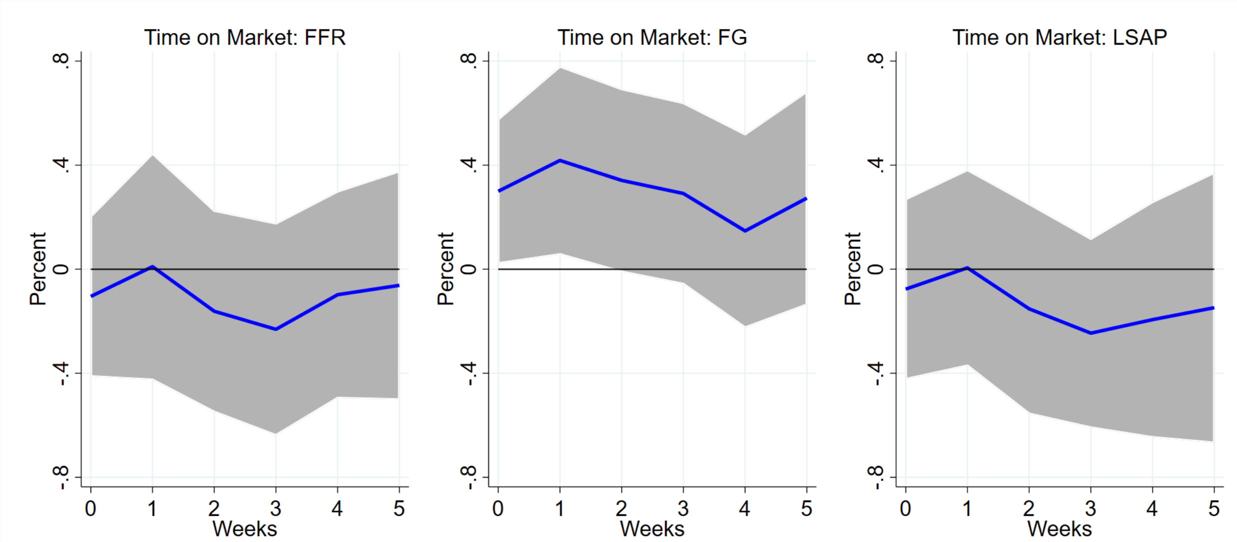


Figure 6. Responses of Time-on-Market to Contractionary Monetary Policy Surprises

Notes: The figure shows responses of the time-on-market index to a positive one-standard-deviation impulse to Swanson (2021)'s fed funds factor (left), forward guidance factor (middle), and (negative of) LSAP factor (right). Responses are estimated using specification (4). Shaded areas represent the 90 percent confidence intervals based on Driscoll-Kraay standard errors.

The responses of the time-on-market index, shown in Figure 6, are not statistically different from zero for the FFR and LSAP factors, and weakly positive for the FG factor. By construction, this index captures the response of the time-on-market for properties listed after the announcement date. Such a response reflects two opposing forces. A contractionary monetary policy shock would cool housing demand and lengthen time-on-market. On the other hand, lower list price helps houses to sell and shorten time-on-market. Anenberg and Ringo (Forthcoming) explain such a trade-off faced by the sellers after an interest rate change in the model with search frictions, long-term fixed-rate mortgages and housing construction. A combination of strong price responses and weak time-on-market responses reflects insensitivity of the shadow value of housing to monetary surprises, as we explained in Section 3.1.

5 Mechanism Behind House Price Responses

In this section, we demonstrate that surprises to the expected future interest rates affect house prices by raising mortgage rates. Higher mortgage rates reduce availability of credit, cooling housing demand, especially from financially constrained households.

5.1 The response of mortgage rates to monetary surprises

Figure 7 shows the estimated responses to monetary policy surprises of the mean weekly 30-year fixed-rate mortgage rates in the United states. The weekly mortgage rate data are from the Primary Mortgage Market Survey.[9] In response to a surprise about future interest rates, captured by the FG or LSAP factors, mortgage rates increase by around 4 basis points within a month. By contrast, the response of mortgage rates to the level surprise is close to zero. Based on data through 2006, Hamilton (2008) demonstrates that changes in information about the level and slope of the federal funds rate are positively correlated with 30-year mortgage rates, with slope effects 2.6 times stronger than level effects. Hamilton argues that the mortgage rate response materializes as soon as markets realize the changes in the path of the federal funds rate. Our results in Figure 7 suggest that it takes only a few weeks for monetary surprises to be reflected in mortgage rates.

Fast responses of mortgage rates to monetary surprises align with our baseline results where list prices react strongly to surprises about future interest rates and are insensitive to current rate shocks. They suggest monetary surprises affect list prices primarily via mortgage rates. The finding that increases in mortgage rates tend to lower house prices via house financing channel is well documented in the literature. Rising mortgage rates make house financing more expensive by tightening debt-to-income or loan-to-value constraints and lead to lower demand for housing (Anenberg and Kung, 2017; Favilukis et al., 2017; Greenwald, 2018; Garriga et al., 2017; Bhutta and Ringo, 2021). In anticipation of the imminent fall in demand, home sellers lower their list prices. Furthermore, sellers

[9] See `https://www.freddiemac.com/pmms/pmms_archives`. The Survey rates represent rates charged between Monday and Wednesday of the corresponding week. We estimate the responses using linear projections on Swanson (2021) factors and four lags of the weekly change in the dependent variable.

may wish to sell their houses sooner than later while mortgage rates are still rising. Garriga et al. (2019) show that prices respond not only to changes in interest rates, but also to changes in expected future financial conditions.

Our results are consistent with the house financing channel not only qualitatively, but also quantitatively. The conditional semi-elasticity of list price responses to mortgage rates implied by the results in Figures 2 and 7 is around 5 for FG factor and 7.5 for LSAP factor. These estimates are within the range of long-run semi-elasticities, between 3 and 8, from the literature surveyed in Kuttner (2013). Our results, therefore, suggest the impact of the mortgage-financing mechanism on house price materializes much faster than previously found, within weeks after a policy announcement.

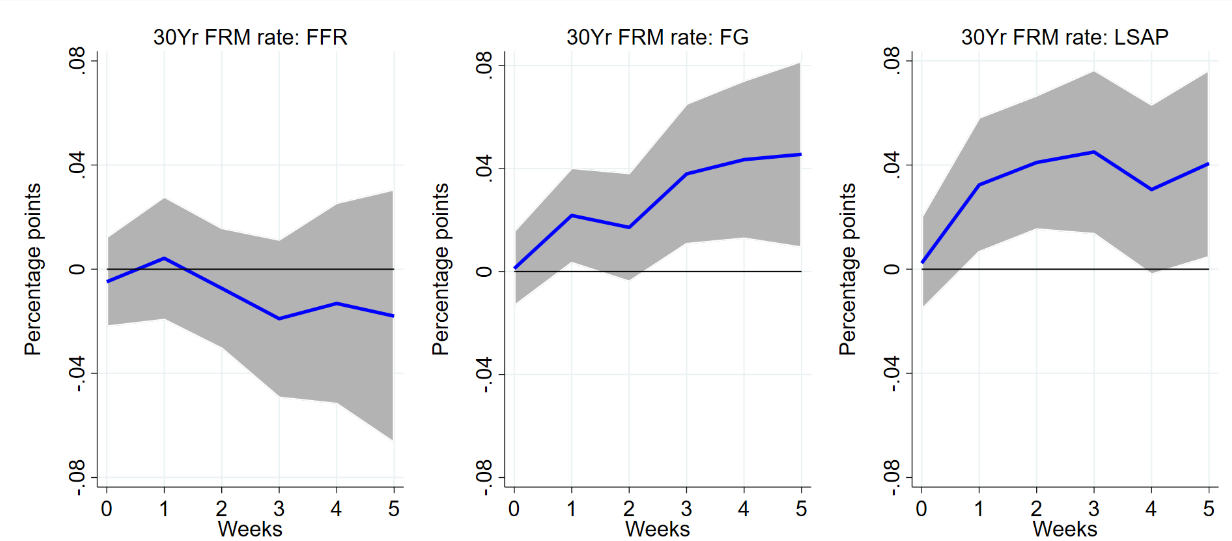


Figure 7. Responses of Mortgage Rates to Contractionary Monetary Policy Surprises

Notes: The figure shows responses of the mean weekly 30-year fixed-rate mortgage rates to a positive one-standard-deviation impulse to Swanson (2021)'s federal funds rate factor (left), forward guidance factor (middle), and (negative of) LSAP factor (right). Mortgage rate data are from the Primary Mortgage Market Survey, `https://www.freddiemac.com/pmms/pmms_archives`. The responses are estimated using linear projections on Swanson factors and four lags of the weekly change in the dependent variable. Shaded areas represent the 90 percent confidence intervals based on robust standard errors.

5.2 The semi-elasticity of list prices with respect to mortgage rates

To obtain a more direct evidence of the mechanism, we estimate the response of list prices to an exogenous variation in mortgage rates using an instrumental-variable approach. In the second stage, we estimate a specification similar to the baseline linear projections (4):

$$\ln P_{l,t+h} - \ln P_{l,t-1} = \alpha^{(h)} + \beta^{(h)} \text{FRM}_t^u + \sum_{q=1}^{52} \theta_q^{(h)} (\ln P_{l,t-q} - \ln P_{l,t-q-1}) + \chi_l^{(h)} + \varepsilon_{l,t}^{(h)}, \quad (5)$$

where $\ln P_{l,t+h} - \ln P_{l,t-1}$ is the cumulative change in log list price index over h weeks after the shock in week t, and FRM_t^u is an exogenous variation in 30-year FRM rates in week t. For each horizon h, we estimate (5) using two stage least squares with fixed effects using three Swanson (2021) factors as instruments. Under our identifying assumptions, variation in mortgage rates is due to monetary surprises, and it is orthogonal to the error $\varepsilon_{l,t}^{(h)}$. Estimated coefficients $\beta^{(h)}$ provide responses of the log list price index to an unanticipated 1 percentage point (ppt) increase in 30-year FRM rates.

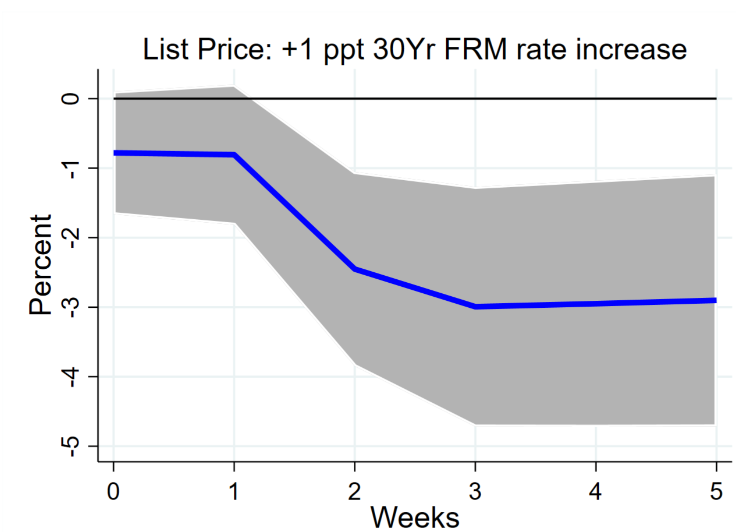


Figure 8. Responses of List Prices to an Exogenous +1 ppt Change in Mortgage Rates

Notes: The figure shows responses of the log list price index to an exogenous +1 ppt increase in the mean weekly 30-year fixed-rate mortgage rates. Mortgage rate data are from the Primary Mortgage Market Survey, `https://www.freddiemac.com/pmms/pmms_archives`. The responses are estimated by two stage least squares with fixed effects using specification (5) with Swanson (2021) factors as instruments. Shaded areas represent the 90 percent confidence intervals based on Driscoll-Kraay standard errors.

Figure 8 shows that a 1 ppt exogenous increase in mortgage rates lowers list prices by around 3 percent within one month.[10] Such a response implies a somewhat lower semi-elasticity of 3, at the low end of the range of 3 to 8 in the literature. Lower estimate is consistent with the range reported in Adelino et al. (2012) who use exogenous variation in credit conditions to estimate the elasticity of house prices to interest rates. Altogether,

[10] The first-stage F-statistic is far above the threshold value for all horizons (Stock, Yogo and Wright, 2002), rejecting the null of weak instruments. p-values for Hansen's J-statistic are well above 10% for all horizons, implying that the model is correctly specified. R^2 is 0.39 or higher for all horizons.

the estimated response is economically and statistically significant, corroborating the important role of mortgage rates for the quick transmission of monetary shocks to house prices.

5.3 Cross-section evidence

Rich variation of macroeconomic and house market outcomes across locations provides another avenue for evidence on the mechanisms behind price responses. We examine how the responses differ across zip codes along four dimensions: household income, house value, density of local bank branches, and housing supply elasticity. We summarize the data and results here, relegating the details to the Appendix.

Household income and house values across zip codes are obtained from the U.S. Census Bureau, 2016-2020 American Community Survey. House income is the median household income in the past 12 months, in 2020 inflation-adjusted dollars. House values are median dollar values. Total population is obtained at zip code level from the U.S. Census Bureau, 2010 Census.

Branch density is the number of bank branches per 1000 people. The number of bank branches includes branches of state-chartered banks (obtained from The Federal Deposit Insurance Corporation) and credit unions (from the National Credit Union Administration).

House supply elasticities are obtained from Baum-Snow and Han (2021) who estimate elasticities using repeat sales price index and the fraction of the U.S. Census tract developed from 2001 or 2011. We use their quadratic finite mixture model estimates for supply elasticities of residential housing units for 2011.

Table 2 provides summary statistics for the distribution of household income, house value, branch density and housing supply elasticity across zip codes. All variables exhibit significant dispersion across almost 11 thousand zip codes.

Statistic	p25	Median	p75	# zip codes
Median income, 2020 dollars	56,025	72,402	93,108	10,778
Median value, dollars	182,600	270,700	413,100	10,767
Num branches/1000 people	0.14	0.23	0.36	9,181
Housing supply elasticity	0.09	0.17	0.30	8,929

Table 2. Statistics for auxiliary variables, by U.S. zip codes.

Notes: Statistics are weighted by the number of listings."p25" and "p75" refer to 25th and 75th percentiles.

To assess how price responses to monetary shocks differ across zip codes, we first re-estimate baseline specification (4) separately for zip codes in the top and bottom quartiles of income, value, branch density and house supply elasticity. We later complement these results with estimates that use these controls jointly.

List prices in zip codes with lower household incomes and house values are more sensitive to slope surprises than those in high-income and high-value zip codes. For example, Figure 9 provides the responses for zip codes in the top and bottom income quartiles. This evidence is consistent with the financing mechanism driving house price responses. Zip codes with lower household income or lower house values have a higher fraction of financially constrained buyers and sellers, multiplying the impact of monetary shocks on house prices. Adelino et al. (2012) estimate local elasticity of house prices to interest rates using exogenous variation in conforming loan limit to instrument for lower cost of house financing. They report stronger elasticities for zip codes in the lowest income quartile and zip codes with low income growth. As we report in the Appendix, differences in responses across house value quartiles are somewhat smaller than differences across income quartiles, suggesting a more influential role of payment-to-income constraints than loan-to-value constraints (Garriga et al., 2017; Greenwald, 2018). List price responses do not vary with housing supply elasticity. As we show in the Appendix, these responses are not sensitive to including income, branch density and supply elasticity jointly in the regressions.[11]

Lastly, list price responses appear less sensitive to surprises in the level of federal funds rate in zip codes with fewer bank branches, but more sensitive to slope surprises (see Appendix). The response to the level surprise is consistent with the literature that argues that in less concentrated markets the pass-through of monetary policy to lending and mortgage rates is lower (Scharfstein and Sunderam, 2016; Drechsler, Savov and Schnabl, 2017). The responses to slope surprises seem to go in the opposite direction. Wang, Whited, Wu and Xiao (2022) argue that when interest rates are low, interaction of market power with bank capital regulation may reverse the pass-through.

[11] In the Appendix, we report that income and house values are strongly correlated across zip codes (correlation 0.68), and house elasticities are negatively correlated with incomes (correlation –0.09) and with house values (correlation –0.30). Branch density is uncorrelated with income or house values.

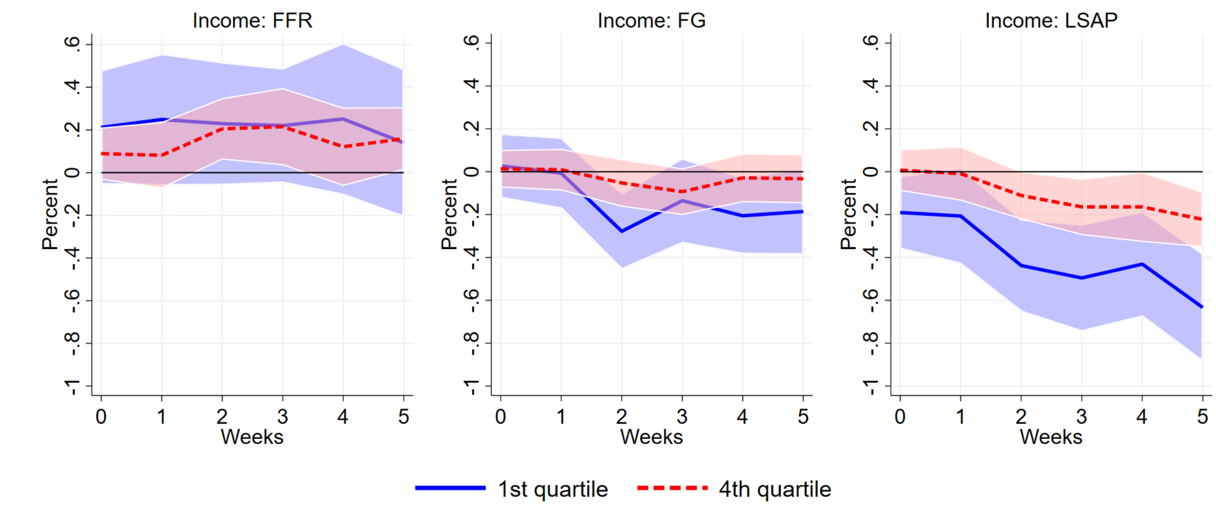


Figure 9. Responses of List Prices for Zip Codes in the Top and Bottom Quartiles of the Distribution of Median Household Income in 2016–2020.

Notes: The figure shows responses of the list price index to a one-standard-deviation increase in Swanson (2021)'s federal funds rate factor (left), forward guidance factor (middle), and (negative of) LSAP factor (right). Responses are estimated using specification (4) for top and bottom quartiles of distribution of median household income across zip codes. Shaded areas represent the 90 percent confidence intervals based on Driscoll-Kraay standard errors.

6 Additional Results

6.1 Long-horizon responses

In the baseline, we kept the impulse response horizon to within 5 weeks after the surprise to contain the event window to one FOMC announcement and avoid subsequent FOMC events from influencing the estimates. Figure 10 shows the estimated responses for the baseline regression but over longer horizons of up to 12 weeks after the surprise. The main results remain the same. One noteworthy difference is an even stronger fall in house prices after the LSAP surprise, reaching below –0.5 percent 7 weeks after the shock.

6.2 Price behavior before the surprise

We also extend the estimated responses to horizons prior to a policy announcement. The null hypothesis is that list prices do not respond to surprises associated with forthcoming announcements. We, therefore, estimate specification (4) for horizons $h = -2, -3,$ $-4, -5$, where for $h = -2$ the dependent variable $\ln P_{l,t-2} - \ln P_{l,t-1}$ is the cumulative change in log list price index between weeks 1 and 2 prior to the announcement in week t, and so on for $h = -3, -4, -5$. Lagged weekly changes on the right-hand side start with

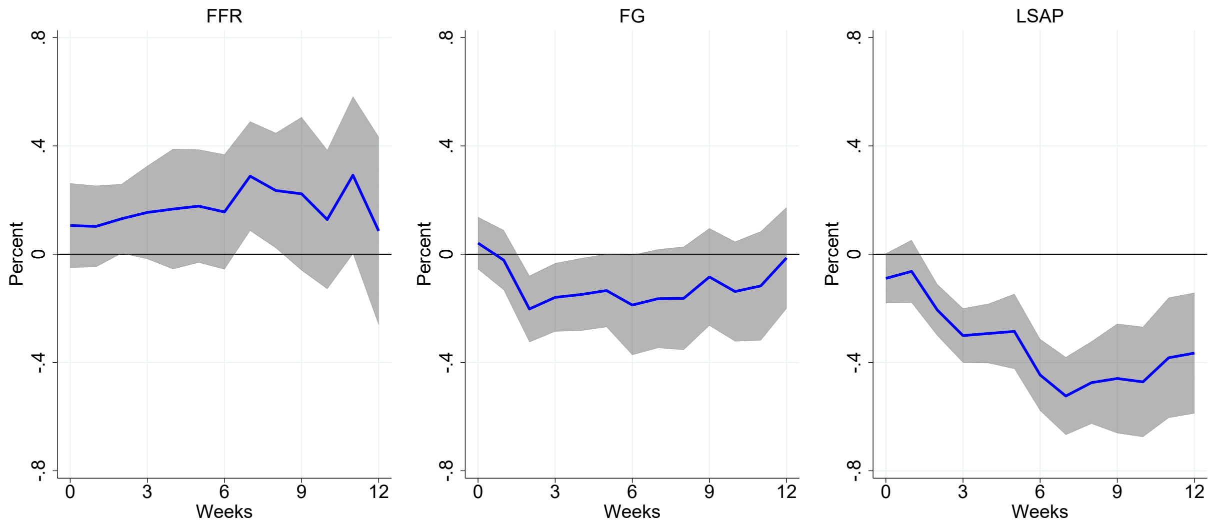


Figure 10. Responses of List Prices to Contractionary Monetary Policy Surprises Over 12 Weeks after the Shock

Notes: The figure provides responses of the list price index to a positive one-standard-deviation impulse to Swanson (2021)'s federal funds rate factor, forward guidance factor, and (negative of) LSAP factor. Responses are estimated using specification (4). Shaded areas represent the 90 percent confidence intervals based on Driscoll-Kraay standard errors.

lag 6: $\sum_{q=6}^{52} \theta_q^{(h)} (\ln P_{l,t-q} - \ln P_{l,t-q-1})$.

It turns out zero pre-announcement responses are rejected in this case (see Appendix). The reason for this is that Swanson factors are correlated with lagged house market indexes. This correlation may reflect adjustments by the financial markets to economic news omitted in the construction of high-frequency measures of monetary surprises (Bauer and Swanson, 2021, 2022). Therefore, we purge Swanson factors of information in lagged house market indexes by regressing them on weekly changes in log list price index over the year preceding monetary policy announcement. We then repeat the estimation of pre-announcement responses using filtered factors as monetary shocks.

Figure 11 shows that all prior responses are now not statistically different from zero, as expected. Furthermore, the responses in the weeks following the FG and LSAP surprises are close to our baseline estimates, suggesting that filtering of Swanson factors does not change our main results.

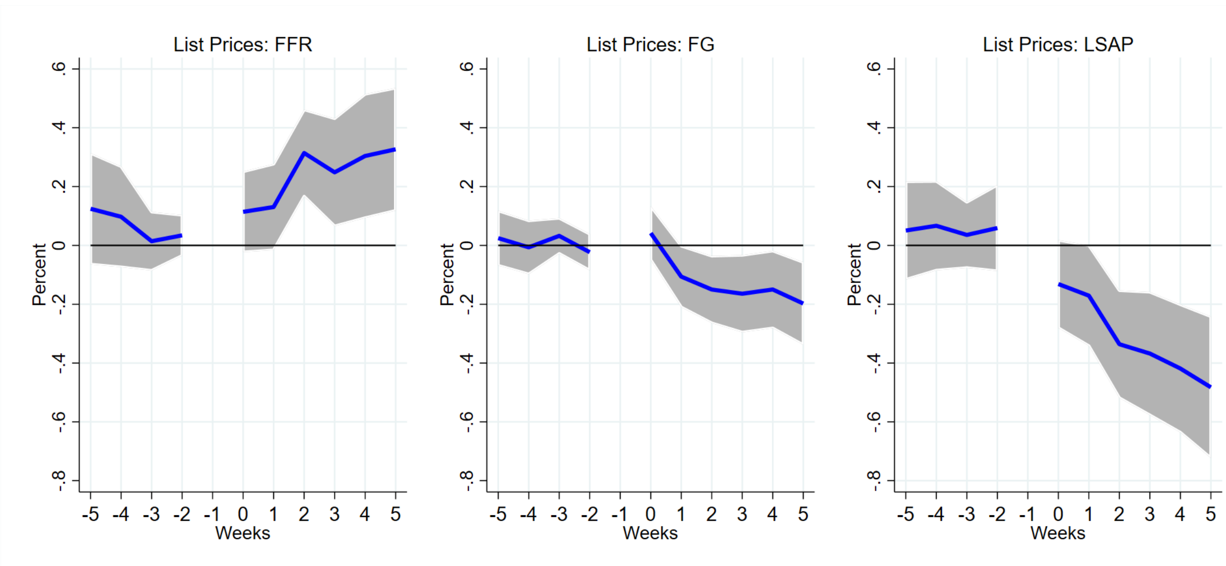


Figure 11. Responses of List Prices Prior to a Policy Announcement.

Notes: The figure extends the responses of the list price index to weeks prior to monetary policy announcement. we purge Swanson factors of information in lagged house market indexes by regressing them on weekly changes in log list price index over the quarter preceding monetary policy announcement. We then repeat the estimation of prior responses using filtered factors as monetary shocks. Shaded areas represent the 90 percent confidence intervals based on Driscoll-Kraay standard errors.

7 Conclusions

We present new evidence that challenges the conventional wisdom that house prices respond to monetary policy shocks with a considerable lag. In fact, prices of listed houses respond to a surprise within a couple a weeks, reaching magnitudes on par with responses of financial assets to policy announcements.

Future research can provide further evidence on the mechanisms behind prompt house price responses. Our results suggest sellers' list price decisions play a crucial role in allocating house market outcomes, whereas sale prices mostly follow the adjustments of list prices. We show evidence that house prices react to changes in mortgage rates in the wake of the monetary shock, indicating that activities related to house financing play a role in house market adjustment to a policy announcements. Indeed, we show that house prices in locations with lower incomes and house values tend to be more responsive.

The second natural continuation of our findings is their implications for policy. Since housing is a major component of households' consumption and wealth, sensitivity of house prices to monetary shocks must have important implications for the effectiveness of monetary policy to stabilize consumption, income, wealth, prices, and their respective distributions across population.

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A Cross-section evidence

We merge price data by zip codes with cross-section data for household income, house value, density of local bank branches, and housing supply elasticity.

A.1 Data

Household income is obtained at zip5 level from the U.S. Census Bureau, 2016-2020 American Community Survey (ACS) 5-Year Estimates, series B19013_001E "MEDIAN HOUSEHOLD INCOME IN THE PAST 12 MONTHS (IN 2020 INFLATION-ADJUSTED DOLLARS)".

House values are obtained at zip5 level from the U.S. Census Bureau, 2016-2020 American Community Survey (ACS) 5-Year Estimates, series B25077_001E "MEDIAN VALUE (DOLLARS)".

Total population is obtained at zip5 level from the U.S. Census Bureau, 2010 Census.

Number of bank branches includes branches of state-chartered banks that are not members of the Federal Reserve System and State-chartered savings associations and federal and state-chartered credit unions. The number of state bank branches is obtained at zip5 from The Federal Deposit Insurance Corporation (FDIC) website. The number of credit union branches is obtained at zip5 level from The National Credit Union Administration. We define **branch density** by the number of branches per 1000 people.

House supply elasticities are obtained from Baum-Snow and Han (2021) who estimate elasticities using repeat sales price index and fraction tract developed from 2001 or 2011. We use their quadratic finite mixture model (FMM) estimates for supply elasticities of housing units for 2011,"gamma11b_unit_FMM". The elasticities are estimated at Census tract level. The tract to zip crosswalk is downloaded from HUD's Office of Policy Development and Research website at `https://www.huduser.gov/portal/datasets/usps_crosswalk.html#codebook`.

A.2 Evidence

To assess how price responses to monetary shocks differ across zip codes, we re-estimate baseline specification (4) separately for zip codes in the top/bottom quartiles of income, value, branch density and house supply elasticity. Figure 9 in the main text compares the responses for top/bottom income quartiles. Figure A.1 provides the results for other variables.

	Median income	Median value	Branch density
Median value	**0.68**		
	0.00		
Branch density	**0.01**	**0.00**	
	0.22	*0.91*	
Supply elasticity	**-0.09**	**-0.30**	**-0.02**
	0.00	*0.00*	*0.0086*

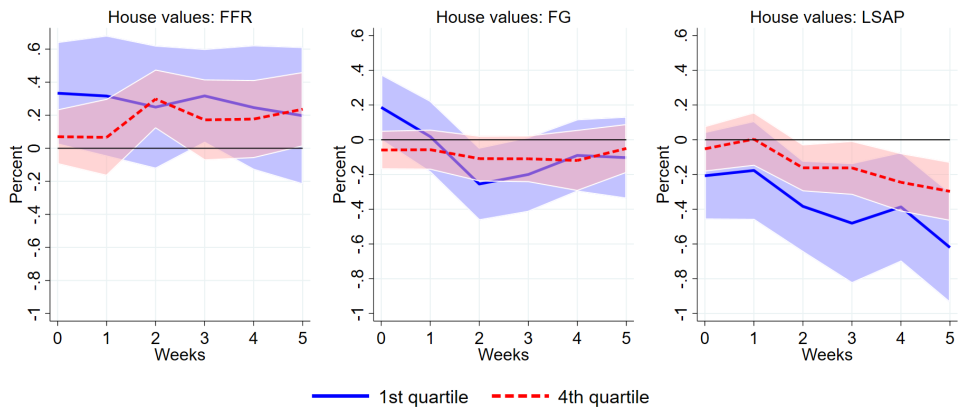
Table A.1. Correlations across U.S. zip codes.

Notes: Statistics are unweighted. p-values are provided in italics.

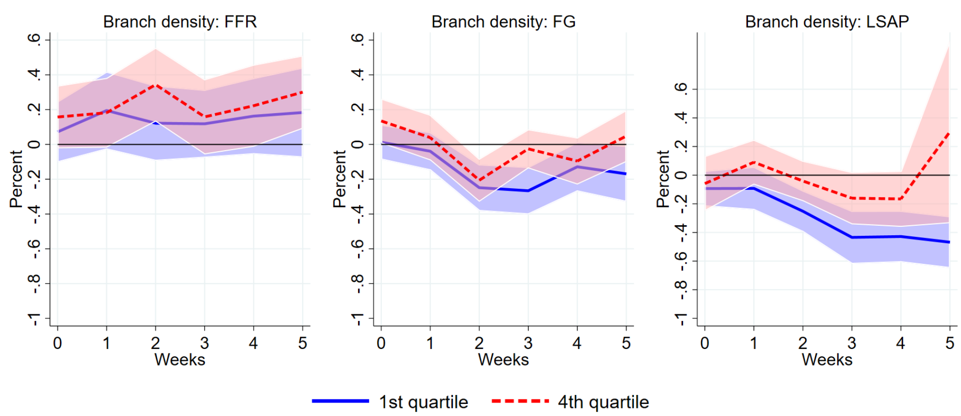
Response differentials across zip codes reported here and in the main text are not influenced by mutual correlations reported in Table A.1. To this end, we standardize log income, density and elasticity variables across zip codes by subtracting weighted median and dividing by weighted inter-quartile range. For example, standardized log income for zip code l is $\widehat{INC}_l = \frac{\ln INC_l - median(\ln INC_l)}{iqr(\ln INC_l)}$, and we apply similar definitions branch density $\widehat{BRA}_l$, and housing supply elasticity $\widehat{ELA}_l$. We then run baseline regression (4) with 9 additional interaction terms with three Swanson factors:

$$\begin{aligned}
\ln P_{l,t+h} - \ln P_{l,t-1} = \alpha^{(h)} &+ \beta^{(h)}_{FFR} \cdot FFR_t + \beta^{(h)}_{FG} \cdot FG_t + \beta^{(h)}_{LSAP} \cdot LSAP_t \\
&+ \beta^{(h)}_{INC-FFR} \cdot \widehat{INC}_l \cdot FFR_t + \beta^{(h)}_{INC-FG} \cdot \widehat{INC}_l \cdot FG_t + \beta^{(h)}_{INC-LSAP} \cdot \widehat{INC}_l \cdot LSAP_t \\
&+ \beta^{(h)}_{BRA-FFR} \cdot \widehat{BRA}_l \cdot FFR_t + \beta^{(h)}_{BRA-FG} \cdot \widehat{BRA}_l \cdot FG_t + \beta^{(h)}_{BRA-LSAP} \cdot \widehat{ELA}_l \cdot LSAP_t \\
&+ \beta^{(h)}_{ELA-FFR} \cdot \widehat{ELA}_l \cdot FFR_t + \beta^{(h)}_{ELA-FG} \cdot \widehat{ELA}_l \cdot FG_t + \beta^{(h)}_{ELA-LSAP} \cdot \widehat{ELA}_l \cdot LSAP_t \\
&+ \sum_{q=1}^{52} \theta^{(h)}_q (\ln P_{l,t-q} - \ln P_{l,t-q-1}) + \chi^{(h)}_l + \varepsilon^{(h)}_{l,t}
\end{aligned} \tag{6}$$

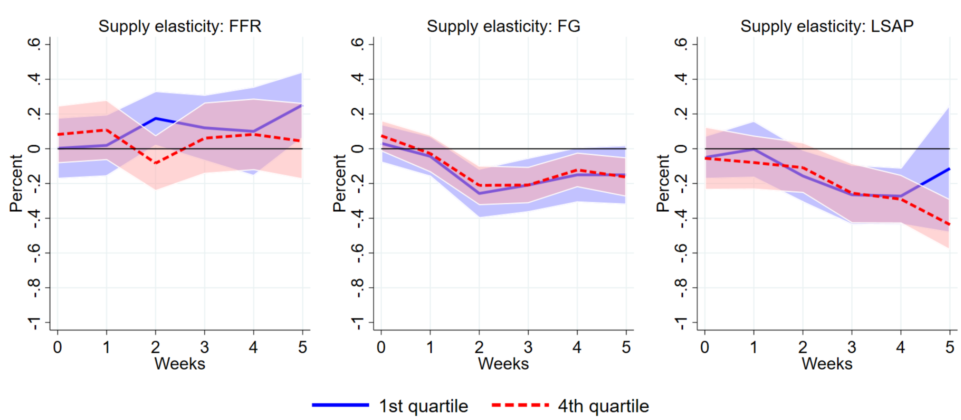
where $\beta^{(h)}_{INC-FFR}, \beta^{(h)}_{INC-FG}, \beta^{(h)}_{INC-LSAP}, ...$ estimated coefficients for 9 interactions of log income, branch density, supply elasticity with three Swanson factors at horizon h. We did not include interactions with house value as it is highly correlated with household income. The estimated coefficients, depicted on Figure A.2 yield results similar to those reported for each of these variables separately.



(a) Responses by house value



(b) Responses by the number of branches per 1000 people



(c) Responses by house supply elasticity

Figure A.1. Responses across zip codes.

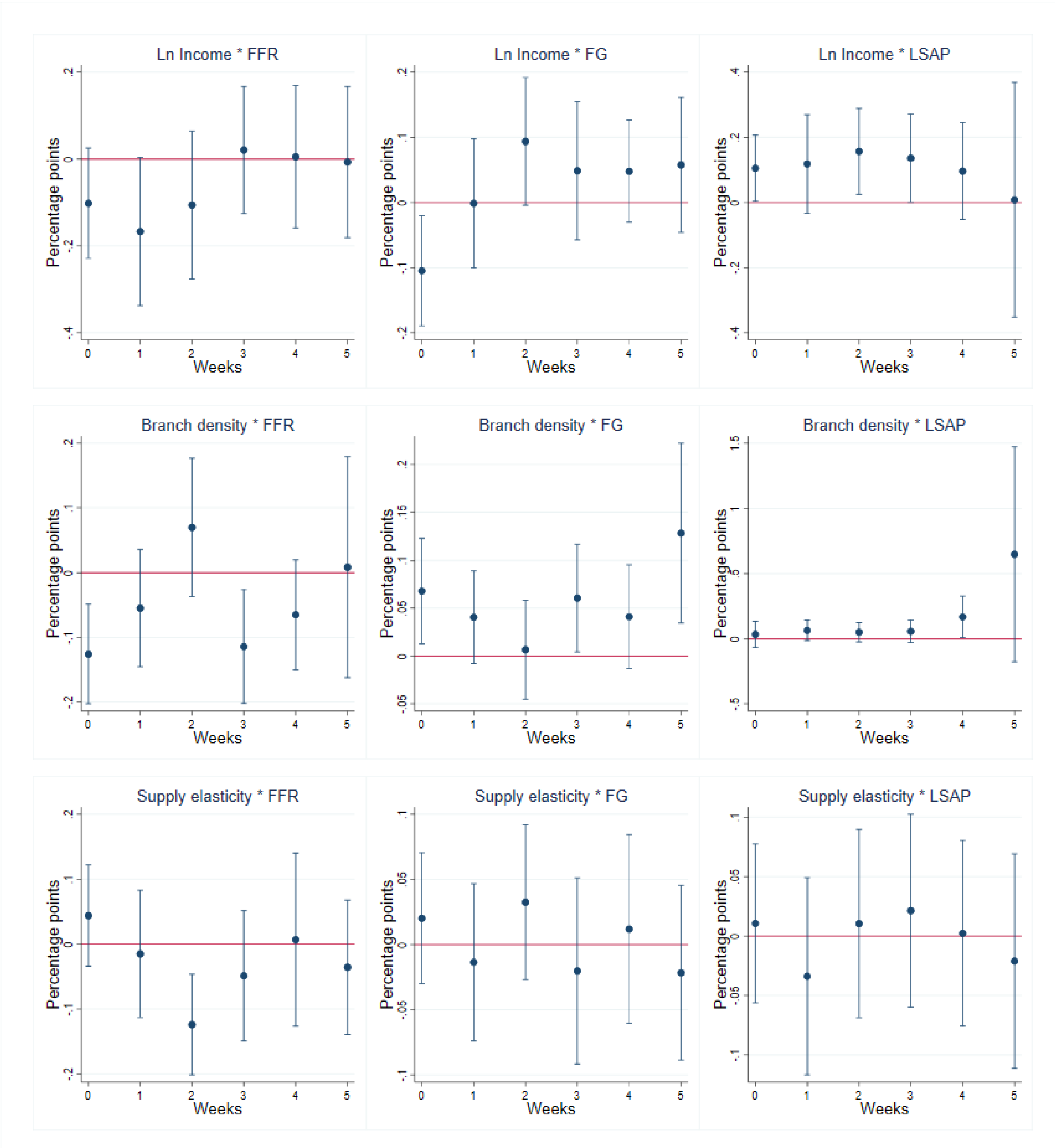


Figure A.2. Estimated coefficients for interaction terms.

Notes: The figure shows estimate coefficients for interaction terms in regression (6) estimating responses of the list price index to a one-standard-deviation increase in Swanson (2021)'s federal funds rate factor (left), forward guidance factor (middle), and (negative of) LSAP factor (right). Interaction terms are for log income (top), branch density (middle), housing supply elasticity (bottom). Shaded areas represent the 90 percent confidence intervals based on Driscoll-Kraay standard errors.

B Pre-announcement responses

We extend the estimated responses to horizons prior to a policy announcement. The null hypothesis is that list prices do not respond to surprises associated with forthcoming announcements. To this end, we estimate baseline specification (4) for horizons $h = -2, -3, -4, -5$, where for $h = -2$ the dependent variable $\ln P_{l,t-2} - \ln P_{l,t-1}$ is the cumulative change in log list price index between weeks 2 and 1 prior to the announcement in week t, and so on for $h = -3, -4, -5$. Lagged weekly changes on the right-hand side start with lag 6: $\sum_{q=6}^{52} \theta_q^{(h)} (\ln P_{l,t-q} - \ln P_{l,t-q-1})$.

Figure B.1 shows that zero pre-announcement responses are rejected. The reason for this is that Swanson factors are correlated with lagged house market indexes. In the main text we purge Swanson factors of information in lagged house market indexes by regressing them on weekly changes in log list price index over the year preceding monetary policy announcement. We then repeat the estimation of pre-announcement responses using filtered factors as monetary shocks. After filtering, all prior responses are now not statistically different from zero, and the responses in the weeks following the FG and LSAP surprises remain close to our baseline estimates.

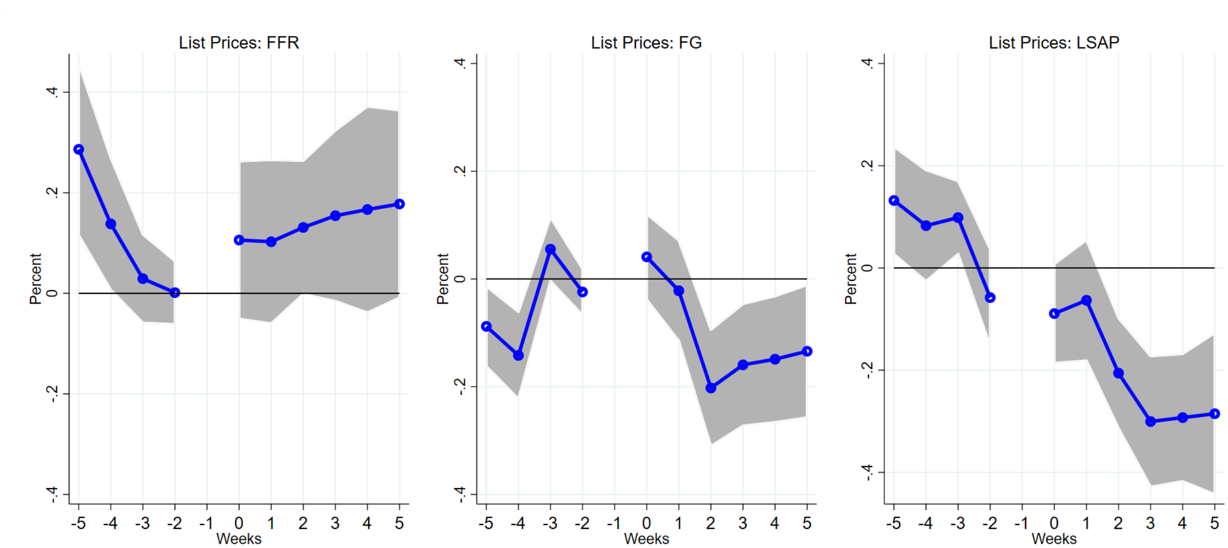


Figure B.1. Responses of Housing List Prices to Contractionary Monetary Policy Surprises

Notes: The figure shows responses of the list price index to a one-standard-deviation increase in Swanson (2021)'s federal funds rate factor (left), forward guidance factor (middle), and (negative of) LSAP factor (right). Responses are estimated using specification (4). Shaded areas represent the 90 percent confidence intervals based on Driscoll-Kraay standard errors.