House Prices, Increasing Returns, and the Effects of Government Spending Shocks*

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Abstract

We show that U.S. house prices increase in response to a fiscal expansion. In contrast, conventional dynamic general equilibrium models predict a drop in house prices, due to the simultaneous increase in the present-value tax burden. We devise a model combining endogenous entry and exit with taste for variety. This generates increasing returns in aggregate production, inducing total factor productivity to increase in the face of a fiscal expansion, thus overcoming the negative wealth effect. As a result, the response of house prices flips sign. The connection between house prices, business formation, and productivity finds strong support in the data.

Keywords: Fiscal Policy, House Prices, Endogenous Entry, Taste for Variety.

JEL classifications: E13, E20, E32, E62.

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1 Introduction

The Great Recession accentuated the importance of the housing market in shaping the macroeconomy. Since then, examining the response of house prices to a variety of shocks, as well as unveiling their interplay with various macroeconomic aggregates, has taken center stage in academics' and practitioners' research agendas. Concurrently, the last decade has witnessed an increasing interest in the macroeconomic effects of fiscal policy, both in academic and policy circles. Yet, surprisingly few studies have examined the link between changes in government spending and house prices, empirically as well as theoretically.

The contribution of this paper is twofold. First, we devise a Bayesian Vector Autoregression (BVAR) model to document that an expansion in fiscal spending generates a persistent increase in house prices alongside a range of other macroeconomic variables (in this respect, see also Khan and Reza, 2017; Auerbach et al., 2019; Alpanda et al., 2021). We corroborate this finding using contract data from the Department of Defense (DoD), reporting that a positive change in federal government spending in a given city expands the price of housing relative to other cities. Thus, our second and main contribution consists of devising and validating a mechanism that produces a coherent transmission of shocks to fiscal spending on the variables involved in both our aggregate and regional evidence, with a special focus on house prices.

A major limitation of standard dynamic stochastic general equilibrium (DSGE) models is that, in sharp contrast with the available evidence, house prices contract in the face of a fiscal expansion. This makes our second contribution particularly significant. In fact, the problem affects any standard framework in which a Ricardian household participates in the housing market, either as the only type of household in the economy or in conjunction with other agents. Why is this the case? As originally highlighted by Barsky et al. (2007)—who focus on explaining the counterfactual negative comovement between durable and nondurable goods consumption in the face of a monetary policy shock—the problem lies in that housing is a long-lived durable and, as such, it features an approximately constant shadow value, from the perspective of a Ricardian household. Two key elements lie behind this property. First, the marginal utility of housing depends on the stock of housing, which is weakly affected by changes in its flow. Second, temporary shocks—such as those to government spending—exert little influence on the future marginal utility of housing.[1] Following an increase in government spending, the present value of lifetime after-tax income drops, thus raising the shadow value of lenders' income, and reducing their consumption. Since the shadow value of housing remains approximately constant, the relative price of housing must track the behavior of Ricardian households' nondurable consumption. As discussed by

[1]In this respect, housing preference shocks represent an exception, as they feature directly in the housing Euler equation, thus breaking the direct link between the house price and the marginal utility of consumption. See, e.g., Iacoviello and Neri (2010) or Liu et al. (2013).

Khan and Reza (2017), any conventional remedy proposed so far—such as restrictions to housing supply, nominal stickiness, deep habits, and complementarity between private and public consumption—proves to be inadequate at breaking the quasi-constancy property of Ricardian households' shadow value of housing, even when producing consumption crowding-in.

To overcome this structural limitation, we focus on the conditional behavior of Ricardian households' shadow value of income. To this end, we devise a flexible-price model embedding a lender-borrower relationship with two layers of production: A fully competitive final good sector, and a monopolistically competitive intermediate goods sector. Combining endogenous entry in the intermediate goods sector with a certain degree of 'taste for variety' generates increasing returns to scale in the aggregate. The resulting channel has the potential to overcome the negative wealth effect induced by an increase in fiscal spending (financed either through a tax hike or an increase in government debt), so that Ricardian households' shadow value of income drops. How is this possible? To address this question, it is instructive to examine how the labor market equilibrium is attained in a standard model with no firm entry. In this case, an expansion in government spending leads to an increase in labor supply, at given factor prices. Holding the number of intermediate goods producers fixed typically implies a fall in the real wage, which exacerbates the drop in the present value of disposable income. As a result, households' shadow value of income expands, thus driving the house price down.

With free entry, instead, enhanced profit opportunities determine an increase in the number of intermediate goods producers, compressing their price markups and expanding total factor productivity (TFP); the so-called *competition effect* (see, e.g., Lewis and Winkler, 2017).[2] We first examine the potential of this channel in a simplified version of the model that we solve analytically. Within this context, the competition effect in isolation may be sufficient to generate a positive response of private consumption and the house price: An expansion in fiscal spending generates a positive impetus to firm entry and, thus, to TFP and labor demand (see, e.g., Hall, 2009), potentially overcoming the rise in labor supply, and ultimately increasing the real wage. However, in line with the literature we discuss below, this can only be the case if the steady-state markup and/or the Frisch elasticity of labor supply are set at implausibly high values. Therefore, we complement this mechanism with taste for variety à la Benassy (1996). This implies that, as the number of intermediate goods producers within a given sector increases, the aggregate sectoral good expands for a given input of intermediate goods; the *variety effect*. In this case, we derive the minimal degree of taste for variety required to induce a joint increase in nondurable consumption and house prices, for given values of the markup and the Frisch elasticity.

Combining the variety and the competition effects within our quantitative framework entails

[2] In this case, while output is a constant returns function to the primary factors of production—for a given measure of intermediates—an increase in the number of intermediates shifts the relationship between output and the production factors.

a robust increase in TFP and the wage rate, in response to a fiscal stimulus. Altogether, this leads to a substitution out of leisure and into consumption for both borrowers and—for the sake of generating a positive response of house prices—lenders. To discipline the model, we match the impulse responses from the model to the empirical ones from our BVAR, which features house prices, output, consumption, TFP, mortgage debt, and the real wage, along with federal government spending and tax revenues. In line with the quantitative model, all these variables increase following a fiscal expansion. In particular, the model accounts for between half and two thirds of the cumulative increase in house prices observed in the data throughout the first six years after the shock. In addition, our modeling strategy allows us to obtain independent estimates of the parameter controlling the taste for variety and the steady-state markup in the intermediate goods sector, as predicted by the analytical framework. The estimation scheme prefers a parameter combination with a moderate steady-state markup and a strong taste for variety, confirming that the variety effect is crucial in matching the data. Indeed, an estimated version of the model without taste for variety fails to match the empirical evidence.

The validatation of the propagation mechanism embedded in our model relies not only on the BVAR model employed as a quantitative benchmark, but also on regional data. We find that a positive change in federal government spending in a given city expands both the number of establishments and labor productivity, relative to other cities. Thus, we report that cities characterized by a sharper increase in house prices also display sharper hikes in business formation and labor productivity, in line with our narrative.[3]

Related literature The link between government spending shocks, net firm entry, and consumption crowding-in has previously been studied by Devereux et al. (1996) and Lewis and Winkler (2017), although none of them focus on house prices. In Devereux et al. (1996), firm entry generates increasing returns to specialization. This effect is closely related to the variety effect in our model, but it is distinct from the competition effect.[4] The specification in Lewis and Winkler (2017), instead, embeds a competition effect, but no variety effect. In both cases, the authors conclude that their baseline model requires unrealistically high values of the markup and/or the Frisch elasticity in order to generate a positive response of consumption, consistent with our analytical insights (in this respect, see also Bilbiie, 2011).

We contribute to a large literature on the macroeconomic effects of shocks to government spending, as extensively surveyed by Ramey (2016). The response of house prices to such shocks has received very little attention, with the exception of Khan and Reza (2017) and Alpanda et al.

[3]The empirical plausibility of our proposed mechanism is also supported by Epstein et al. (2022), who document a strong cross-country link between new firm creation and movements in house prices, though with no specific focus on government spending shocks.

[4]In our model, the variety effect is directly tied to the parameter measuring the taste for variety, whereas it is not separately parametrized in the model of Devereux et al. (1996).

(2021), who both report a positive effect based on structural VAR models for the U.S. economy. We present both aggregate and regional evidence of a positive response in house prices. As discussed above, our proposed mechanism relies on an increase in net firm entry after a government spending shock. This result is empirically supported by Lewis and Winkler (2017) in a structural VAR model using U.S. data. As already mentioned, these authors conclude that, for realistic parameter values, their baseline DSGE model is unable to generate an increase in consumption, unless government spending is assumed to be utility- or productivity-enhancing. In this respect, we have deemed other avenues to be more fruitful for our purposes, for two main reasons: First, Khan and Reza (2017) demonstrate that, although complementarity between private and public goods in consumer utility can bring about an increase in the consumption of Ricardian households, this does not imply a decline in their shadow value of income, which is necessary for a rise in house prices.[5] Second, the specification used in Lewis and Winkler (2017) assumes the *flow* of government spending to be productivity-enhancing (or, equivalently, that the public capital stock depreciates entirely each period). Assuming instead that what matters for production is the *stock* of public capital, and that this depreciates at a rate roughly similar to that of private capital (as traditionally done in the literature; see, e.g., Baxter and King, 1993; Leeper et al., 2010), we find that this mechanism only produces an increase in private consumption and the house price if the weight of public capital in the production function is prohibitively high.[6]

We choose instead to draw on an emerging literature combining endogenous firm entry with taste for variety. This builds in large part on Bilbiie et al. (2012), who show that incorporating these ingredients improves the empirical performance of standard RBC models in response to productivity shocks. We use a variant of the Constant Elasticity of Substitution (CES) function with generalized love of variety introduced by Benassy (1996). This function disentangles market power from love of variety, such that increasing returns to scale may imply a more marked reactiveness of the real wage to fiscal spending shocks, without requiring implausibly high markups and/or elasticities of labor supply. Several recent papers have also used this specification of the CES function to analyze the implications of endogenous entry and product variety for optimal fiscal policy (Chugh and Ghironi, 2011), optimal monetary policy (Bergin and Corsetti, 2008; Bilbiie et al., 2014), the monetary transmission mechanism (Lewis and Poilly, 2012), the international transmission of productivity shocks (Corsetti et al., 2007), the welfare costs of inefficient entry and variety (Bilbiie et al., 2019), and monetary neutrality (Bilbiie, 2021).

There is scant empirical evidence on plausible values for the parameter governing the extent of taste for variety (Chugh and Ghironi, 2011; Bilbiie et al., 2019). Lewis and Poilly (2012) estimate a DSGE model featuring love of variety using impulse-response matching to monetary policy shocks, and report that the love of variety parameter is poorly identified. While our estimate

[5] This is due to a counterfactual drop in the real wage.

[6] These results are available upon request.

of this parameter is somewhat higher than the values typically considered in the literature, we ascertain that our model produces a positive response of the house price for a wide range of parameter values. Nonetheless, our study echoes the call of Bilbiie et al. (2012) and Bilbiie (2021) for more empirical work on assessing the role of the taste for variety.

Finally, we contribute to a broader literature aiming to model house-price dynamics within DSGE models. A key implication of the insights of Barsky et al. (2007) is that, in any model in which a Ricardian household participates in the housing market, this agent effectively determines how house prices move. Several recent studies of house-price dynamics have circumvented this property by excluding this type of household from the housing market (see, e.g., Ferrero, 2015; Garriga et al., 2019, 2021), thus allowing house prices to be influenced by credit-constrained households.[7] By contrast, we confront this issue head-on, as our approach focuses on altering the dynamics of Ricardian households' shadow value of income, while retaining the property that these agents are responsible for pinning down the equilibrium response of the house price.

Structure The paper proceeds as follows. Section 2 reports empirical evidence based on aggregate and regional data on the response of U.S. house prices in the face of shocks to fiscal spending. In Section 3 we outline the details of the model to be employed in the quantitative analysis. Section 4 devises a stylized version of the quantitative model to provide an analytical inspection of the interplay between the competition and the variety effect in generating consumption crowding-in. Section 5 describes the calibration and estimation of the model. Section 6 discusses the qualitative and quantitative implications of our framework. Section 7 contains a validation exercise of the key transmission channel at work in the model, and Section 8 concludes.

2 Empirical evidence

In this section we provide empirical evidence to support the claim that increases in government spending have a positive effect on U.S. house prices. We first look at aggregate data, so as to establish a benchmark for the quantitative assessment of the framework we will introduce in Section 3. To this end, we devise a structural BVAR, following the tradition of most of the empirical literature on the aggregate effects of government spending shocks. While we are mainly interested in the aggregate effects of fiscal policy, we then seek to corroborate our findings by means of regional data. Thus, we study how a change in federal government spending in a given city—as compared with other cities—affects relative house price movements. We do so by following the approach of Nakamura and Steinsson (2014) and Auerbach et al. (2020b), where military procurement is used as a source of regional variation in spending. Both types of analysis focus

[7]Equivalently, Justiniano et al. (2019) obtain the same effect by assuming that Ricardian households own a fixed share of the aggregate housing stock.

on selected variables whose conditional behavior is key to validate the transmission mechanism embodied by our quantitative framework.

2.1 Aggregate evidence

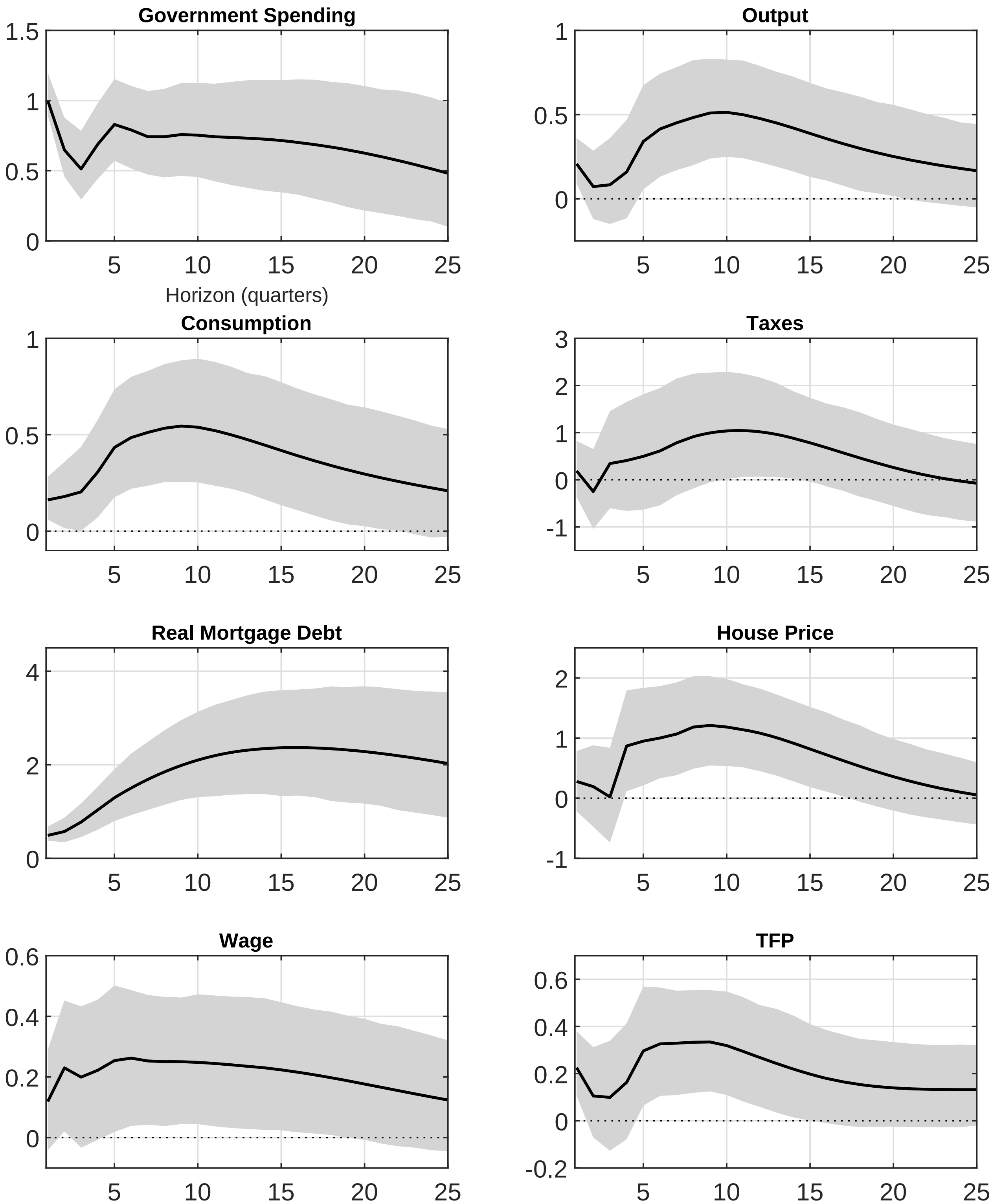
We set up a BVAR model for the U.S. economy. In order to identify truly unexpected government spending shocks that do not suffer from potential anticipation effects, we rely on the survey-based forecast errors of the growth rate of government spending (denoted by FE_t) computed by Auerbach and Gorodnichenko (2012). We order this variable first in the BVAR model, thus identifying an unanticipated government spending shock as an innovation to the forecast error of the growth rate of government spending. The BVAR further includes the following variables: Real government consumption and investment (G_t), real GDP (Y_t), real private consumption (C_t), real net tax revenues (T_t), real mortgage debt (B_t), the real house price (Q_t), the real wage (W_t), and total factor productivity (TFP_t). We use the Median Sales Price of Houses Sold, which is constructed by the U.S. Census Bureau, and deflate it using the GDP deflator.[8] The data span from 1966:Q4 to 2019:Q4, with the start of the sample dictated by the availability of FE_t from Auerbach and Gorodnichenko (2012), and its end chosen so as to avoid the turbulence of Covid-19. We use the update of the forecast error series of Auerbach and Gorodnichenko (2012) until 2019 constructed by Jørgensen and Ravn (2022). The BVAR includes 4 lags and a constant, and we detrend all variables except FE_t using a linear and a quadratic trend. We use a standard Minnesota prior, with the tightness of the prior determined by the data using the method of Giannone et al. (2015). Additional details regarding the BVAR and the data are provided in Appendix A.1.[9]

In Figure 1 we present the impulse responses from the BVAR to a positive government spending shock normalized to 1 percent, along with 68 percent credible sets. Following such a shock, we observe a persistent and hump-shaped increase in the house price. This finding is consistent with the evidence reported by Khan and Reza (2017) and Alpanda et al. (2021). To get a sense of the magnitude of the effect, we can divide the responses by the sample average of the ratio of government spending to GDP (which equals 0.237). This results into a house-price response that peaks at around 5 percent, two years after an increase in government spending equal to 1 percent of GDP. Furthermore, the positive responses observed for output, consumption, and the wage

[8] Other popular house price indices, such as the Case-Shiller National Home Price Index or the All-Transactions House Price Index, are only available for shorter samples. Since government spending shocks have been found to be much smaller and less persistent since around 1980 (e.g. Bilbiie et al., 2008), we opt for retaining a long data sample.

[9] We have chosen not to include a measure of business formation in the BVAR model, although this dimension will play an important role in our theoretical analysis. The data series for firm entry used by Lewis and Winkler (2017) ends in 1995 and, as discussed by these authors, more recent data series are not fully comparable. We prioritize a longer sample period, and therefore exclude this variable. Instead, we consider it in our regional analysis described below.

Figure 1: Estimated effects of a government spending shock from the BVAR model



Notes: The figure shows the effects of a shock to government spending estimated in the BVAR model (solid black line), with the grey areas representing the 68 percent credible sets (i.e., the 16th and 84th percentiles of the posterior distribution based on 1000 draws.)

rate are in line with most of the existing literature, e.g. Galí et al. (2007), while the increase in TFP is consistent with recent studies by d'Alessandro et al. (2019) and Jørgensen and Ravn (2022), among others. The increase in mortgage debt corroborates recent evidence reported by Auerbach et al. (2020a) and Bayer et al. (2020), who both find that government spending has a stimulative effect on credit markets.

2.2 Regional evidence

The analysis on regional data relies on (yearly) Department of Defense (DoD) contract data from the website USAspending.gov, covering the 2001-2019 time window. This website contains information on individual prime contracts signed between companies and the DoD, which we aggregate up to the Metropolitan Statistical Area (MSA) level, to get a variable for all DoD contracts obligated annually to each MSA. We refer to this variable as DoD spending. Additional information on the data and the aggregation procedure is described in Appendix A.2.1. To measure local house prices, we use the Freddie Mac House Price Index, while we normalize DoD spending by local economic activity using GDP from the Bureau of Economic Analysis (BEA). The final panel dataset covers 380 MSAs from 2001 through 2019, at the annual frequency.

We estimate the following regression of house price growth in MSA i over h years on the initial change in (normalized) DoD spending over one year:

$$\frac{Q_{i,t+h} - Q_{i,t}}{Q_{i,t}} = \alpha_{i,h} + \eta_{t+h} + \beta_h \frac{G_{i,t+1} - G_{i,t}}{Y_{i,t}} + \gamma_h X_{i,t} + \varepsilon_{i,t+h}, \tag{2.1}$$

where $Q_{i,t}$ denotes the house price index, $G_{i,t}$ is DoD spending, $X_{i,t}$ is a vector of controls, and $Y_{i,t}$ is GDP (all at the MSA level). The MSA fixed effect, $\alpha_{i,h}$, controls for MSA-specific trends in house prices, while the time fixed effect, η_{t+h}, controls for common, national variation in house prices.[10] All variables are measured in nominal terms, although we obtain similar results upon transforming $Q_{i,t}$, $G_{i,t}$, and $Y_{i,t}$ to real terms using the MSA-level GDP deflator.[11]

The coefficient of interest is β_h, which measures the growth in house prices from t to $t+h$ relative to other MSAs, as a result of a 1 percent increase in DoD spending from period t to

[10] The MSA-level normalized change in DoD spending is winsorized at the 1 percent level by year, since the series contains outliers (the maximum is around 10 times larger than the 99th percentile). Non-winsorized estimates are somewhat smaller in magnitude but qualitatively similar, as shown in Appendix A.2.5.

[11] We use nominal values, since there are no official statistics that accurately measure cross-regional differences in prices. Although the BEA produces MSA-level GDP deflators, they are constructed by applying national price indices to current dollar values of MSA-level GDP at the industry level (Bureau of Economic Analysis, 2015). Hence, these statistics do not capture cross-regional differences in prices, but differences in industry composition. Hazell et al. (2022) circumvent this imputation issue by constructing regional price indices using BLS micro data. However, these indices are only constructed for 34 states.

$t+1$, and relative to period-t GDP.[12] However, the OLS estimate of β_h is likely to be biased, since military contracts tend to flow disproportionately more to areas that experience relatively bad economic outcomes, due to political factors influencing the allocation of contracts (Nakamura and Steinsson, 2014).

We deal with the potential bias by resorting to a Bartik (1991) instrument: The change in national DoD spending interacted with the MSA's average share of DoD spending to local GDP over the sample period. This instrument identifies the effect of spending on house prices by relating changes in the MSAs' DoD spending to their persistent and differential exposure to changes in national military spending. That is, when the federal government expands military spending, some MSAs tend to receive more DoD contracts than others, because they are systematically more exposed to changes in military spending. This systematic component of changes in local DoD spending is isolated by the instrument. In this respect, Figure A.2 in Appendix A.2 plots the period-by-period first-stage Kleibergen-Papp F-statistics. The F-statistics are in the range 26-63, which is above the cluster-robust threshold for weak instruments of 23.1 provided by Montiel Olea and Pflueger (2013).

The identifying assumption behind this approach is that, conditional on controls, there are no confounding factors affecting local house price growth—not just contemporaneously, but also at different leads and lags—that are correlated with the MSAs' exposure to changes in military spending over the cross section, as well as with changes in national military spending in the time-series dimension:[13]

$$E\left[\varepsilon_{i,t+h+j} \times \left(\bar{G}_i \frac{G_{t+1}^{nat} - G_t^{nat}}{Y_{i,t}}\right) | X_{i,t}\right] = 0 \quad \text{for } j \in \{\dots, -2, -1, 0, 1, 2, \dots\}, \tag{2.2}$$

where $\bar{G}_i = \frac{1}{T}\sum_t \frac{G_{i,t}}{Y_{i,t}}$ is MSA i's average share of DoD spending to local GDP over the sample period, and $\frac{G_{t+1}^{nat} - G_t^{nat}}{Y_{i,t}}$ is the change in national DoD spending from period t to $t+1$, normalized by local GDP in period t.

The average DoD spending share in a given MSA is likely to be an equilibrium object determined by factors such as industry composition, or by having a military base nearby. For this reason, it is worth stressing that the exogeneity condition is formulated in terms of *changes* in the outcome, rather than *levels*. Even if the level of local house prices was codetermined with

[12]The effect of government spending over h periods is captured by β_h, and results from both the effect of changes in spending from period t to $t+1$, as well as from the subsequent flows in spending induced by the initial shock. When estimating spending multipliers, it is common to use the cumulative change in spending over the response horizon, instead of the initial change in spending—as we do in our model—since this allows for direct estimation of cumulative multipliers. The reason for using changes in the initial level of spending is that this makes our estimates comparable to the impulse responses from the BVAR that will serve as a basis for the calibration of the model.

[13]The lead-lag exogeneity condition is stronger than conventional IV contemporaneous exogeneity conditions, since the dependent variable depends on past and future shocks that must be orthogonal to the contemporaneous instrument (Stock and Watson, 2018).

the local DoD spending share, equation (2.2) would still hold. However, a potential concern is that the MSAs' exposure to military spending is related to their exposure to the national business cycle—for instance through different industry or housing market composition—which drives the differential house price response across MSAs through correlation between national DoD spending and the business cycle. More formally, this would imply that the exogeneity condition (2.2) does not hold, because the error term $\varepsilon_{i,t+h+j}$ carries a $\gamma_i \xi_{t+h+j}$ structure, where γ_i is correlated with $\bar{G}_i$ over the cross section, and ξ_{t+h+j} is correlated with $\frac{G^{nat}_{t+1}-G^{nat}_{t}}{Y_{i,t}}$ in the time-series dimension. We address some of these concerns in Appendix A.2.5, through a number of robustness checks, as summarized in the next subsection.

An additional identifying assumption is necessary because of the dynamic effects of government spending: Lagged and leading shocks to local government spending should also be unrelated to the contemporaneous instrument (Stock and Watson, 2018). This assumption does not hold, since the instrument itself is autocorrelated. In that case, the estimate of β_h will not only pick up the contemporaneous effect from the shock, but also the effects from past shocks, so that β_h cannot be interpreted as the effect of an unanticipated shock to DoD spending. For this reason, we follow Ramey and Zubairy (2018) by including two lags of the instrument, $\bar{G}_i \frac{G^{nat}_{t+1}-G^{nat}_{t}}{Y_{i,t}}$, as well as two lags of the one-year normalized change in local spending, $\frac{G_{i,t+1}-G_{i,t}}{Y_{i,t}}$. Table A.3 illustrates that conditioning on autoregressive dynamics in the instrument and local spending leads to a smaller estimate of the house price response, thus capturing past changes in shocks to local spending. Finally, we add two lags of the one-year growth in the dependent variable to pick up momentum in house prices. As we show in Appendix A.2.3, house prices are slightly— yet, significantly—lower in the two years prior to a spending change, which could reflect anticipation effects.[14] However, controlling for this pre-trend—as we do in our baseline regression—makes little difference to our results besides reducing standard errors.

A word of caution is needed on the use of MSA-level data. By construction, the regional framework we employ does not allow one to capture aggregate general equilibrium effects, being these subsumed by the time dummies. This impairs our capacity to compare the regional estimates to the aggregate results obtained above. The spirit of our regional analysis is therefore mainly corroborative with respect to our aggregate BVAR evidence.

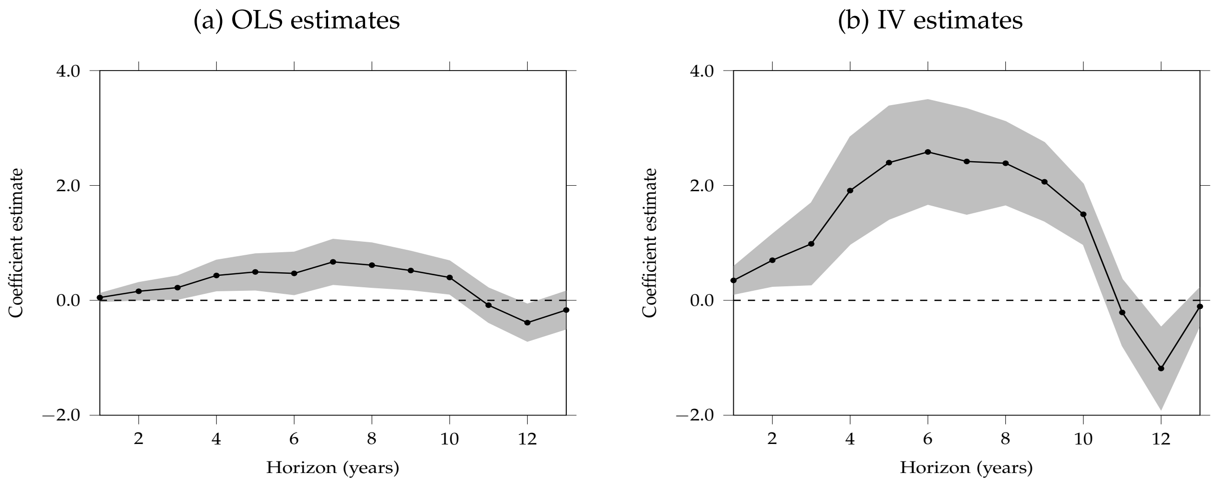
2.2.1 The response of house prices to a fiscal spending shock

We estimate h separate (2.1) regressions, for $h = 1, 2, \ldots, 13$ horizons, and report the estimates in Figure 2. The OLS and IV estimates are shown in panels (a) and (b), respectively. Standard errors are heteroskedasticity-robust and clustered at the MSA level, so as to account for within-

[14] As Auerbach et al. (2019) argue, the instrument does not filter out all anticipation effects since anticipated changes in national military spending are not removed by the instrument.

MSA correlation of the error term. Clustering by MSA allows for non-parametric time series dependence in the errors by taking advantage of the cross-sectional dimension of the data.[15] Grey areas indicate 95 percent confidence bands based on the point estimate standard errors.

Figure 2: Regional house price responses to military spending



Notes: The figure shows the estimates of β_h from regression (2.1) based on an annual panel of 380 MSAs covering the period 2001-2019. The OLS and IV estimates for the house price response are plotted in panels (a) and (b), respectively. The regressions include as controls two lags of the one-year growth in house prices, two lags of the instrument, and two lags of the one-year change in local spending normalized by GDP. Grey areas indicate the 95 percent confidence bands constructed using heteroskedasticity-robust standard errors clustered by MSA.

According to the IV estimates, the response of house prices to an expansion in government spending follows a hump-shaped pattern, peaking after six years, thus reverting back to trend. In terms of magnitude, we appreciate a relative increase in house prices of 2.6 percent after six years, as a result of an increase in spending of 1 percent of GDP in the first year. The OLS estimates also follow a hump-shaped pattern, but are biased toward zero. The difference between the IV and OLS estimates is sizeable, with non-overlapping confidence bands until the 11-year horizon. Formally, period-by-period endogeneity tests firmly reject the null of the IV estimate being equal to the OLS estimate until $h = 11$, as shown in Figure A.4 in Appendix A.2.

These estimates are robust to a number of alternative specifications of (2.1), as reported in Appendix A.2.5. Specifically, we present results from regressions with real variables, alternative normalizations of DoD spending changes, a proxy for DoD outlays instead of obligations, and controls for differential house price movements associated with potential confounding factors, such as local industry composition and exposure to regional business cycles. All of these robustness checks indicate results that are analogous to those in the baseline analysis. We also confirm

[15]Using clustered standard errors tends to be more conservative than relying on a heteroskedasticity-and-autocorrelation (HAC) robust estimator since the latter typically imposes a parametric autoregressive structure on the regression errors (Jordà et al., 2015).

our findings to be invariant to using the beginning-of-sample share of DoD spending to local GDP, instead of the average share over the sample. In addition, we examine the robustness of our results to outliers and different sets of controls in the baseline regression. There are some key takeaways from the robustness exercises, which support a causal interpretation of the IV estimates. We show that our results are robust to controlling for three well-known measures of house-price cycle exposure interacted with year dummies.[16] This addresses the concern that our results could reflect a differential exposure to the 2000s boom-bust cycle in U.S. house prices. Moreover, differences in industry composition do not drive our results since controlling for two-digit industry employment shares multiplied by year dummies does not alter the estimates.

3 The model

We devise a real business cycle economy populated by two types of households, differentiated by their discount factors: Impatient households have a lower discount factor than patient households, and can borrow up to a share of the present value of their housing stock. This implies that patient households act as lenders. Both household types work, consume nondurables and accumulate housing. Patient households also accumulate capital that is rented to firms producing intermediate goods. The inclusion of impatient households is based on our desire to explain movements in mortgage debt alongside those in house prices, as these two variables are closely related in the data.

Production of nondurables and investment goods occurs in a two-layer production sector, in the vein of Rotemberg and Woodford (1992), Jaimovich (2007), and Jaimovich and Floetotto (2008), among others. The first production layer consists of a continuum of sectors of measure one. Each sector contains a finite number of firms producing differentiated sector-specific goods using capital and labor as inputs, while firms enter and exit the sectors until a zero-profit condition is satisfied. The differentiated goods are bundled to produce an aggregate sectoral good to be used as an input in the second production layer. That layer consists of a representative firm combining the continuum of aggregate sectoral goods to produce a final good to be sold to households and the government.

3.1 Households

The economy is populated by two groups of households, each consisting of a continuum of unit mass. Both household types derive utility from nondurable consumption, C_t^j, housing, H_t^j, and the fraction of time devoted to labor, N_j^t, where $j \in \{b, l\}$ indexes impatient and patient household-specific variables, respectively. Each type of household maximizes the following life-time utility

[16] The three measures of house-price cycle exposure are the Wharton Regulation Index, the Saiz (2010) instrument, and the Bartik-like instrument for sensitivity to regional house price movements by Guren et al. (2021).

function:
$$E_0 \left\{ \sum_{t=0}^{\infty} \left(\beta^j\right)^t \left[\frac{\left(C_t^j - h^j C_{t-1}^j\right)^{1-\sigma_c}}{1-\sigma_c} + \Upsilon^j \frac{\left(H_t^j\right)^{1-\sigma_h}}{1-\sigma_h} - \Psi^j \frac{\left(N_t^j\right)^{1+\psi}}{1+\psi} \right] \right\}, \tag{3.1}$$

where $\beta^l > \beta^b$ are the discount factors. This difference in impatience implies that patient households will act as lenders to impatient households. In addition, $\sigma_c \geq 0$ and $\sigma_h \geq 0$ are the coefficients of relative risk aversion for consumption and housing, respectively, ψ is the inverse Frisch elasticity, and $h^j \in [0;1[$ measures the degree of internal habit formation in consumption, while $\Upsilon^j > 0$ and $\Psi^j > 0$ are the utility weights on housing and labor, respectively.

Impatient households choose consumption, housing, labor, and borrowing subject to their budget constraint and a loan-to-value constraint:

$$C_t^b + q_t H_t^b + R_{t-1} B_{t-1}^b = w_t^b N_t^b + B_t^b + q_t H_{t-1}^b - \tau_t^b, \tag{3.2}$$

$$B_t^b \leq \gamma B_{t-1}^b + (1-\gamma) m \frac{E_t \left\{ q_{t+1} H_t^b \right\}}{R_t}, \tag{3.3}$$

where q_t is the price of housing in units of consumption, B_t^b is the stock of real debt held at the end of period t, R_t is the gross real interest rate on debt between period t and $t+1$, w_t^b is the real wage of impatient households, and τ_t^b is a lump-sum tax.

The borrowing constraint in equation (3.3) states that impatient households can borrow up to a fraction $m \in [0;1]$ of the present value of their housing stock at the beginning of the next period, as in Kiyotaki and Moore (1997). Following Guerrieri and Iacoviello (2017), we allow for inertia in the dynamics of mortgage debt, as measured by $\gamma \in [0;1[$, to account for the slow moving nature of the stock of debt in the data. We assume that shocks to the economy are sufficiently small that the borrowing constraint invariantly holds with equality in the neighborhood of the steady state.

Patient households choose consumption, housing, labor, capital, investment, and bond holdings subject to their budget constraint:

$$C_t^l + q_t H_t^l + I_t + B_t^l + B_t^g = w_t^l N_t^l + q_t H_{t-1}^l + R_{t-1} B_{t-1}^l + R_{t-1} B_{t-1}^g + r_t^k K_{t-1} - \tau_t^l, \tag{3.4}$$

where I_t is investment in capital, B_t^l are one-period bonds at the end of period t, B_t^g denotes one-period government bonds (which, for simplicity, are assumed to earn the same risk-free interest rate as private bonds), w_t^l is the real wage of patient households, r_t^k is the real rental rate of capital, and τ_t^l is a lump-sum tax. We assume that capital rented to the firms evolves according to the following law of motion:

$$K_t = K_{t-1}(1-\delta) + I_t (1 - \Phi_t), \tag{3.5}$$

where $\delta \in [0;1]$ denotes the depreciation rate, and $\Phi_t = \frac{\phi}{2}\left(\frac{I_t}{K_{t-1}} - \delta\right)^2 \frac{K_{t-1}}{I_t}$ is convex costs of capital adjustment, with $\phi > 0$. Appendix B reports the first-order conditions for the borrower and other agents in the model economy.

3.2 Production

Production occurs in two stages. A first layer of intermediate goods firms produces distinct intermediate goods using capital rented from the patient households and labor supplied by both household types. There exists a continuum of sectors indexed by $j \in [0;1]$, with each of these sectors consisting of $F_t(j)$ intermediate goods firms. These firms sell their goods to a representative final good firm in a monopolistically competitive market subject to free entry. Second, the final good firm transforms the intermediate goods into aggregate sectoral goods, $\{Q_t(j)\}_{j=0}^{1}$, which in turn are aggregated into a final good, Y_t, that is sold to households and the government in a perfectly competitive market.

3.2.1 Final good production

The final good, Y_t, is produced by a representative firm using a CES production function that aggregates a continuum of measure one of aggregate sectoral goods:

$$Y_t = \left[\int_0^1 Q_t(j)^{\omega} dj\right]^{\frac{1}{\omega}}, \qquad \omega \in]0;1[. \tag{3.6}$$

Each intermediate good sector consists of $F_t(j) > 1$ firms producing differentiated goods that are aggregated into a sectoral good using the following aggregation function proposed by Benassy (1996):

$$Q_t(j) = F_t(j)^{\tau + \frac{\rho-1}{\rho}} \left[\sum_{i=1}^{F_t(j)} m_t(j,i)^{\rho}\right]^{\frac{1}{\rho}} \qquad \rho \in]0;1[, \tag{3.7}$$

where $m_t(j,i)$ is the output of firm i in sector j. The production function in equation (3.7) is a generalization of the Dixit and Stiglitz (1977) CES aggregation function that disentangles the variety effect from the elasticity of substitution across inputs, $1/(1-\rho)$.[17] The variety effect is measured by $\tau \geq 0$, and implies that, as the number of intermediate firms within a sector increases, the sectoral aggregate good increases for a given input of intermediate goods. If $\tau = -(\rho-1)/\rho$, the function reduces to the Dixit-Stiglitz function, in which the variety effect is tied to the elasticity of substitution, while $\tau = 0$ implies that the variety effect is nil.[18]

The final good firm's demand for each sectoral aggregate good, $Q_t(j)$, is given by the following

[17] A similar function was already studied in a working-paper version of Dixit and Stiglitz (1977); see Dixit and Stiglitz (1975).

[18] Alternatively, the variety effect can be modeled by assuming that consumers derive utility directly from an increase in the number of intermediate goods, as in Lewis and Poilly (2012), Bilbiie et al. (2012), or Bilbiie et al. (2019). In this alternative interpretation, however, one would need to adjust for the variety effect when taking the model to the data, as the welfare-consistent price index in such a model includes the variety effect, while the CPI constructed by the BLS does not.

standard demand function:

$$Q_t(j) = \left(\frac{p_t(j)}{P_t}\right)^{\frac{1}{\omega-1}} Y_t, \tag{3.8}$$

where $p_t(j)$ is the price index for the sector j aggregate good, and $P_t = \left[\int_0^1 p_t(j)^{\frac{\omega}{\omega-1}} dj\right]^{\frac{\omega-1}{\omega}}$ is the aggregate price index.

In turn, the demand for good $m_t(j,i)$ follows from solving the final good firm's cost minimization problem and is given by

$$m_t(j,i) = \left(\frac{p_t(j,i)}{p_t(j)}\right)^{\frac{1}{\rho-1}} \left(\frac{p_t(j)}{P_t}\right)^{\frac{1}{\omega-1}} \frac{Y_t}{\left(F_t(j)^{\tau+\frac{\rho-1}{\rho}}\right)^{\frac{\rho}{\rho-1}}}, \tag{3.9}$$

where $p_t(j,i)$ is the price of $m_t(j,i)$, and the sectoral price index is equal to

$$p_t(j) = \frac{1}{F_t(j)^{\tau+\frac{\rho-1}{\rho}}} \left[\sum_{i=1}^{F_t(j)} p_t(j,i)^{\frac{\rho}{\rho-1}}\right]^{\frac{\rho-1}{\rho}}. \tag{3.10}$$

Finally, firms sell the final good to households and the government in a competitive fashion.

3.2.2 Intermediate goods production

Each intermediate good, $m_t(j,i)$, is produced using capital and labor purchased in competitive markets, according to the following constant-returns-to-scale production technology:

$$m_t(j,i) = k_{t-1}(j,i)^{\mu} \left[\left(n_t^b(j,i)\right)^{\alpha} \left(n_t^l(j,i)\right)^{1-\alpha}\right]^{1-\mu} - \varphi, \qquad \alpha, \mu \in]0;1[, \tag{3.11}$$

where $\varphi > 0$ is a fixed cost of production, $k_{t-1}(j,i)$ denotes the firm-level capital input, while $n_t^b(j,i)$ and $n_t^l(j,i)$ denote the firm-level labor inputs supplied by impatient and patient households, respectively.

Firms sell their intermediate goods to the final good firms in a monopolistically competitive market within each sector. In doing so, they account for the effect they exert on the sectoral price index, $p_t(j)$, but not on the final good price, P_t, as in Jaimovich (2007). Thus, the elasticity of demand for the intermediate goods firm, according to the demand curve (3.9) and the price index (3.10), is

$$\varepsilon_{m_t(j,i)} = \frac{1}{\rho-1} + \left(\frac{1}{\omega-1} - \frac{1}{\rho-1}\right) \left(\frac{p_t(j,i)}{p_t(j)F_t(j)^{\tau}}\right)^{\frac{\rho}{\rho-1}} \frac{1}{F_t(j)}. \tag{3.12}$$

We assume that the elasticity of substitution within sectors is higher than the elasticity of substitution across sectors, $\frac{1}{1-\omega} < \frac{1}{1-\rho}$.[19] This implies that if an individual firm increases its price,

[19]This assumption is consistent with the evidence by Broda and Weinstein (2006), who show that as product categories are disaggregated, varieties become increasingly substitutable.

$p_t(j,i)$, relative to the sectoral price index adjusted for the variety effect, $p_t(j)F_t(j)^\tau$, the elasticity of demand increases, since the demand for the aggregate sectoral good falls through the firm's effect on the sectoral price index.

The elasticity of demand in equation (3.12) results in firms setting prices at the following markup over the marginal cost:

$$x_t(j,i) = \frac{\varepsilon_{m_t(j,i)}}{1+\varepsilon_{m_t(j,i)}}, \tag{3.13}$$

which is a decreasing function of the number of firms. This highlights the competition effect associated with endogenous entry. Note that the markup converges to the standard constant markup $\frac{1}{\rho}$ as $F_t(j) \to \infty$, and to $\frac{1}{\omega}$ as $F_t(j) \to 1$. Hence, the markup is bounded between $\frac{1}{\rho}$ and $\frac{1}{\omega}$.

Firms' cost minimization results in the following cost function:

$$C_t(j,i) = A\left(r_t^k\right)^\mu \left(w_t^b\right)^{\alpha(1-\mu)} \left(w_t^l\right)^{(1-\alpha)(1-\mu)} \left(m_t(j,i)+\varphi\right), \tag{3.14}$$

where $A \equiv \frac{1}{(1-\mu)^{1-\mu}(1-\alpha)^{(1-\alpha)(1-\mu)}\mu^\mu\alpha^{\alpha(1-\mu)}}$.

We assume that firms can enter and exit sectors freely. They do so until profits are driven to zero, which results in the following free entry condition:

$$\frac{p_t(j,i)}{P_t} m_t(j,i) = C_t(j,i). \tag{3.15}$$

Finally, combining the free entry condition with the cost function in equation (3.14) and the pricing schedule $\frac{p_t(j,i)}{P_t} = x_t(j,i) \cdot \frac{\partial C_t(j,i)}{\partial m_t(j,i)}$ pins down each firm's production as a function of fixed costs and the markup:

$$m_t(j,i)\left(x_t(j,i)-1\right) = \varphi. \tag{3.16}$$

3.2.3 Symmetric firm equilibrium

Intermediate firms face identical technology, entry costs and demand curves for their goods. Thus, we focus on a symmetric equilibrium in which the number of firms is equalized across sectors, all firms set identical prices and produce the same quantity of output using the same amount of production factors. Formally, $\forall (j,i) \in [0;1]\times[1,F_t(j)] : F_t(j) = F_t$, $p_t(j,i) = p_t$, $x_t(j,i) = x_t$, $m_t(j,i) = m_t$, $k_{t-1}(j,i) = k_{t-1}$, $n_t^b(j,i) = n_t^b$, $n_t^l(j,i) = n_t^l$. In addition, market clearing in the capital and labor markets implies that $k_{t-1} = \frac{K_{t-1}}{F_t}$, $n_t^b = \frac{N_t^b}{F_t}$, and $n_t^l = \frac{N_t^l}{F_t}$.

Combining the aggregate price index with the sectoral price index, allows us to express the price of an intermediate good relative to that of the final good, $\frac{p_t}{P_t}$, as a function of the number of firms:

$$\frac{p_t}{P_t} = F_t^\tau. \tag{3.17}$$

Moreover, setting $m_t(j,i) = m_t$ in (3.7) results in

$$Y_t = F_t^{1+\tau} m_t. \tag{3.18}$$

Equations (3.17) and (3.18) yield two key insights about the *variety effect* and its interplay with the *competition effect*. First, p_t/P_t increases in the number of firms, for $\tau > 0$. Increasing the taste for variety lowers the marginal cost of the final good firm, thereby lowering the price of the final good relative to that of the intermediate goods. Second, a larger number of intermediate firms increases final output more than one-for-one, for given intermediate firm-level production. Thus, there are increasing returns to the number of firms, while the production technology at the intermediate-firm level features constant returns to scale.

Analogous considerations about the role of the two effects at play in the model can be made with respect to their impact on TFP. To this end, combining identical price setting with (3.13) returns the markup as a decreasing function of the number of firms:

$$x_t = \frac{(1-\omega)\,F_t - (\rho - \omega)}{\rho\,(1-\omega)\,F_t - (\rho-\omega)}. \tag{3.19}$$

Thus, we can use (3.11), (3.16), and (3.18), together with market clearing in the factor market, to write aggregate output as

$$Y_t = TFP_t K_{t-1}^{\mu} \left[\left(N_t^b\right)^{\alpha} \left(N_t^l\right)^{1-\alpha} \right]^{1-\mu}, \tag{3.20}$$

where $TFP_t \equiv F_t^{\tau}/x_t$ implies that the entry and exit of firms result in endogenous procyclical TFP variations through the competition and the variety effects. As emphasized by Jaimovich and Floetotto (2008), the competition effect stimulates TFP through the impact that changes in the number of firms exert on the markup. To see this, consider an increase in the number of firms fostered by a fiscal expansion, which lowers the markup through more intense competition. In turn, this induces firms to increase production to cover their fixed costs, thus driving TFP up. Furthermore, TFP is affected by the variety effect, as long as $\tau > 0$: A higher number of firms has a direct expansionary impact on TFP, as it raises aggregate output for given primary production factors. A higher τ amplifies this channel. Section 4 will discuss how these effects formally combine to produce a conditional increase in house prices.

3.3 Fiscal policy

Government spending follows an autoregressive process:

$$G_t = (1-\gamma_g)\bar{G} + \gamma_g G_{t-1} + \epsilon_{g,t}, \qquad \epsilon_{g,t} \sim N\left(0, \sigma_g^2\right). \tag{3.21}$$

Each type of household is assumed to pay a fixed share of the lump-sum tax revenue, τ_t^{TOT}, corresponding to their labor income share:

$$\tau_t^b = \alpha \tau_t^{TOT}, \tag{3.22}$$

$$\tau_t^l = (1-\alpha)\,\tau_t^{TOT}. \tag{3.23}$$

The government is allowed to run a non-balanced budget and finance a part of its spending by issuing debt. The government budget constraint is given by:

$$R_{t-1}B^g_{t-1} + G_t = \tau_t^{TOT} + B^g_t. \tag{3.24}$$

Following Leeper et al. (2017), we assume that the tax level adjusts to deviations of the debt-to-GDP ratio from steady state with inertia:

$$\tau_t^{TOT} = \left(\tau_{t-1}^{TOT}\right)^{\rho_\tau} \left(\frac{B^g_{t-1}}{Y_{t-1}}\right)^{(1-\rho_\tau)\gamma_\tau}, \tag{3.25}$$

where $\rho_\tau \in [0; 1[$ measures the degree of inertia in the tax level, and $\gamma_\tau > 0$ is the responsiveness of the tax level to past deviations in the debt level.

3.4 Market clearing

The market clearing conditions are:

$$Y_t = C_t + G_t + I_t, \tag{3.26}$$

$$C_t = C_t^b + C_t^l, \tag{3.27}$$

$$H = H_t^l + H_t^b, \tag{3.28}$$

where H is a fixed stock of housing in the economy.

Lastly, the mortgage market clears when patient-household lending equals impatient-household borrowing:

$$B_t^l = B_t^b. \tag{3.29}$$

4 Inspecting the key mechanisms

In standard models with no endogenous firm entry and no taste for variety, an expansionary shock to fiscal spending produces an increase in labor supply that leads to a drop in the real wage and a simultaneous fall in consumption; the usual crowding-out effect of fiscal spending induced by an increase in the present value of lump-sum taxes (see Baxter and King, 1993). Along with implying counterfactual conditional movements in nondurable consumption and the real wage, this also represents a problem for the conditional response of the price of housing, in virtually any model where Ricardian agents benefit from housing services. To see why, consider patient households' Euler equation for housing (i.e., equation B.6 in Appendix B), which may be solved forward to yield an expression for their shadow value of housing:

$$q_t \lambda_t^l = \Upsilon^l \mathrm{E}_t \left\{ \sum_{t=i}^{\infty} \left(\beta^l\right)^i \left(H_{t+i}^l\right)^{-\sigma_h} \right\} \equiv \Lambda_t. \tag{4.1}$$

Since housing does not depreciate, H_t^l is effectively an "idealized durable" according to Barsky et al. (2007): This means that the intertemporal elasticity of substitution in housing demand is close to infinite. As a result, short-term movements in H_t^l—as those generated by a temporary shock to fiscal spending—will affect the right-hand side of (4.1) relatively little, given that β^l is close to one. So, it is possible to approximate

$$q_t \lambda_t^l = \Lambda_t \approx \Lambda. \tag{4.2}$$

According to this, movements in the price of housing are forced to mirror movements in patient households' shadow value of income, as our quantitative analysis in Section 6 will confirm. In light of this, any model where a Ricardian household participates in the housing market—even as the only type of household in the economy—may be able to generate a conditional expansion in house prices only to the extent that it is capable of generating a decline in λ_t^l.[20] In the absence of channels that break the approximate constancy of the shadow value of housing, overcoming the negative wealth effect of an increase in public spending—by inducing a positive response of nondurable consumption and a concurrent drop in λ_t^l—represents the only viable option.[21]

Our solution consists of combining two mechanisms that generate a sizeable increase in the real wage, following a positive shock to fiscal spending, so that patient households' nondurable consumption and, therefore, the price of housing, both increase. First, we allow for endogenous entry in the (monopolistically competitive) intermediate goods market, which implies aggregate TFP to be endogenously determined by the markup, x_t: What is crucial, in this sense, is that expansionary shocks to public spending stimulate the entry of new producers due to enhanced profit opportunities, thus exerting a positive impulse on TFP. While increasing returns via free entry have already been introduced in a model with monopolistic competition by Devereux et al. (1996), they find that generating a crowding-in of private consumption is only possible through counterfactually high levels of the markup. Thus, to enhance the capacity of the competition effect to overcome the negative wealth effect of an expansion in fiscal spending, we allow for a certain degree of taste for variety, in the vein of Benassy (1996). This feature induces the aggregate production technology to exhibit increasing returns to scale in the primary factors of production, so that we may complement the competition effect, while allowing for plausible values of the markup. Both channels induce labor demand to shift outward in the face of the fiscal stimulus, overcoming the expansion in labor supply, and ultimately increasing the real wage.

To illustrate how these mechanisms formally combine, we solve analytically a simplified ver-

[20]In other words, in any model in which a Ricardian household participates in the housing market, this agent effectively determines how the house price moves.

[21]In this respect, the alternative frameworks considered by Khan and Reza (2017) are not able to reproduce a conditional drop in the shadow value of income, even when attaining a crowding-in of patient households' consumption by appealing. For instance, this is the case when imposing complementarity between private and public spending, as in Bouakez and Rebei (2007).

sion of the model in Section 3. We assume the economy to be solely populated by financially unconstrained households that exhibit logarithmic nondurable consumption utility, and intermediate goods firms featuring a production technology that is linear in labor, the only production input. The resulting list of equilibrium conditions is reported in Appendix C. These can be combined to show that the response of log-linear aggregate production is such that[22]

$$\hat{y}_t = \frac{\theta x (1+\tau)\left[(\rho x - 1) - (\rho - 1)(1+\tau)\right]}{x\left[(\tau - \psi)(1-\theta) - (1+\tau)\right]\left[(\rho - 1)(1+\tau) - (\rho x - 1)\right] - (\rho x - 1)(1-\theta)(1+\psi)(x - 1 - \tau)}\hat{g}_t, \tag{4.3}$$

where θ is the steady-state share of fiscal spending-to-GDP, so that $\hat{c}_t = \frac{1}{1-\theta}\hat{y}_t - \frac{\theta}{1-\theta}\hat{g}_t$.

The first step consists of evaluating the role of endogenous entry in isolation, thus setting $\tau = 0$, so that (4.3) reduces to:

$$\hat{y}_t = \frac{\theta \rho x}{x\rho\left[1 + \psi(1-\theta)\right] - (\rho x - 1)(1-\theta)(1+\psi)}\hat{g}_t. \tag{4.4}$$

In light of this, we can show that a necessary condition to observe a crowding-in of nondurable consumption—i.e., $\hat{y}_t > \theta \hat{g}_t$ which, given the assumption of log-utility, is sufficient to obtain a positive response of the price of housing—is

$$\rho x > 1 + \psi. \tag{4.5}$$

As x is bounded below by $\frac{1}{\rho}$, so that $\rho x > 1$, the condition is satisfied—conditional on conventional values of ρ and x—only in the presence of a relatively elastic labor supply (recall that ψ is the inverse of the Frisch elasticity). This is because, under a relatively low ψ, households are more prone to substitute out of leisure and into consumption in response to the increase in TFP induced by entry in the intermediate goods market. With this in mind, it is easy to see how (4.5) embodies the problems encountered in the existing literature when trying to generate consumption crowding-in through endogenous firm entry: Conditional on a realistic value of ρ, the condition can only be satisfied for unconventionally high values of the markup, x, consistent with the numerical results of Devereux et al. (1996); or for values of the Frisch elasticity, $\frac{1}{\psi}$, that are inconsistent with microeconometric studies, as discussed by Bilbiie (2011). A similar point was made by Lewis and Winkler (2017).

In the general case with taste for variety, instead, it is possible to show that the following condition suffices to ensure a positive response of consumption and the house price:

$$\tau > \frac{(\rho x - 1 - \psi)(1 - x)}{x(1-\rho)}. \tag{4.6}$$

Notice how this condition embeds (4.5): As long as this is satisfied, (4.6) always holds, as the overall expression on its right side is negative. Should this not be the case, crowding-in of nondurable consumption would still be attainable through a taste for variety that is large enough.

[22] Log-linear variables are hatted.

This is typically the case for realistically calibrated values of the elasticity of labor supply and the markup, as we will see in the next section.

5 Estimation and calibration

We split the parameters of the model in Section 3 into two groups. The first group of parameters is calibrated, while the second group is estimated via impulse-response matching.

5.1 Calibration

The vector $\omega_1 = \{\alpha, \beta^b, \beta^l, \delta, \theta, \mu, m, \Xi, \omega, \rho\}$ contains the parameters that we choose to calibrate. We set the income share of borrowers' labor to $\alpha = 0.21$, in line with the estimate of Iacoviello and Neri (2010). The discount factors of borrowers and lenders are set to $\beta^b = 0.97$ and $\beta^l = 0.99$, respectively, as in Jensen et al. (2018). The depreciation rate of capital is set at $\delta = 0.025$, while the income share of capital is set to $\mu = 0.25$. These values imply ratios of investment to output and of capital to output of 0.18 and 1.8, respectively, both of which are broadly in line with the corresponding average values for the U.S. economy. We set the loan-to-value ratio m to 0.85, as in Iacoviello and Neri (2010). The share of government spending to output, denoted by θ, is set to 0.24, while the ratio of public debt to output, denoted by Ξ, is set to 0.7. Both of these numbers are closely in line with the average values for the U.S. over the past decades. We then turn to the parameters governing the elasticity of substitution within and across sectors, ρ and ω. We set $\rho = 0.9$ and $\omega = 0.75$, in order to obtain elasticities of substitution of 10 (within sectors) and 4 (across sectors), respectively. The latter value is closely in line with Bilbiie et al. (2019), who use an elasticity of substitution of 3.8, while the former is chosen to reflect that varieties are increasingly substitutable as product categories are disaggregated, as found by Broda and Weinstein (2006), who estimate elasticities of substitution ranging from 1.2 to 17. Note that the values of ρ and ω encompass the common practice in the New Keynesian literature of setting the elasticity of substitution in one-sector models to 6; see, e.g., Rotemberg and Woodford (1992). We collect the calibrated parameters in Panel A of Table 1.[23]

5.2 Estimation strategy

The remaining parameters are estimated by impulse-response matching, as in Christiano et al. (2005) and Iacoviello (2005), among others. This is done by matching the model-implied impulse responses to a government spending shock to the empirical responses from our BVAR in Section

[23]We also need to set values for the parameters measuring the (dis)utility weights of labor and housing. We set Υ to ensure a ratio of housing wealth to output of 1.45 at the annual frequency, as in Jensen et al. (2018). The weight on labor disutility only affects the scale of the economy, and is simply set to 1.

2.1. We collect in $\omega_2 = \{\gamma, h^b, h^l, \sigma_c, \sigma_h, \tau, \phi, \psi, x, \rho_\tau, \gamma_\tau, \gamma_G, \sigma_g\}$ the parameters to be estimated. Let $\Gamma(\omega_2)$ denote the model-implied impulse responses, which are functions of the parameters, while $\widehat{\Gamma}$ denotes the corresponding empirical estimates from our BVAR model. We obtain the vector of parameter estimates $\widehat{\omega}_2$ by solving:

$$\widehat{\omega}_2 = \arg\min_{\omega_2} \left(\Gamma(\omega_2) - \widehat{\Gamma}\right)' W \left(\Gamma(\omega_2) - \widehat{\Gamma}\right). \tag{5.1}$$

The weighting matrix W is diagonal, with the inverse of the sample variances of the BVAR-based impulse responses as entries. Effectively, this means that we are attaching higher weights to those impulse responses that are estimated most precisely. We match the impulse responses of all variables in our BVAR (except the forecast error series, FE_t) for the first 25 quarters after the shock.[24] To obtain credible sets for the parameter estimates, we employ the entire distribution of impulse responses used to construct the credible sets in Figure 1; that is, we run 1000 estimations of the DSGE model.

5.2.1 Estimation results

Panel B of Table 1 reports the median parameter estimates, as well as the associated credible sets based on the 16th and 84th percentiles of the distribution of the estimates.[25] We first note that most parameters take on values that are generally in line with the existing literature. The degree of inertia in mortgage debt is close to the estimate of Guerrieri and Iacoviello (2017) of 0.7. The estimate of ψ implies a Frisch elasticity somewhat above 3, which is not uncommon in business cycle models with flexible prices.[26]

A distinctive trait of our estimates is that the data seem to emphasize the role of the variety effect more than that of the competition effect. In fact, the steady-state markup, x, is estimated relatively close to the lower bound given by $\frac{1}{\rho} = 1.11$. Under these circumstances—given the estimate of ψ—the condition to obtain consumption crowding-in in the stylized model of Section 4 calls for a sufficiently high value of τ. In fact, under a low steady-state markup, fixed costs are relatively small, and there are relatively many firms with little market power within each sector. As a result, the markup is relatively insensitive to fiscal shocks. Thus, to produce sizable upward changes in TFP—which are key to bringing about an equilibrium increase in patient households' nondurable consumption and, thus, house prices—a relatively high τ is necessary, so as to amplify

[24] We implement a penalty function to drive the procedure away from areas of the parameter space for which the model has no unique and determinate solution.

[25] The credible regions are not symmetric, in part because some of the parameters are bounded above and/or below.

[26] In the estimation, we impose an upper bound of 4 on the Frisch elasticity. While this value is well above microeconometric estimates, it allows traditional RBC models to match business-cycle data (see the discussion by Chetty et al., 2011).

Table 1: Parameter values

Panel A: Calibrated parameters

Parameter	Description	Value
β^l	Discount factor, lenders	0.99
β^b	Discount factor, borrowers	0.97
μ	Capital share of production	0.25
δ	Capital depreciation rate	0.025
α	Income share of impatient households	0.21
θ	Ratio of government spending to output	0.24
Ξ	Ratio of government debt to output	0.7
m	Loan-to-value ratio of borrowers	0.85
ρ	Substitution parameter within sectors	0.9
ω	Substitution parameter across sectors	0.75

Panel B: Estimated parameters

Parameter	Description	Value
σ_c	Curvature in utility of consumption	1.210 [0.711−2.741]
σ_h	Curvature in utility of housing	0.293 [0.105−1.202]
h^l	Habit formation, lenders	0.380 [0.156−0.536]
h^b	Habit formation, borrowers	0.614 [0.405−0.775]
ψ	Inverse Frisch elasticity	0.309 [0.258−0.534]
ϕ	Capital adjustment cost parameter	9.826 [3.994−10.187]
γ	Inertia of mortgage debt	0.742 [0.356−0.871]
τ	Love for variety parameter	4.196 [3.494−4.557]
x	Steady-state value of markup	1.139 [1.127−1.156]
γ_τ	Tax response to government debt	0.529 [0.045−0.778]
ρ_τ	Inertia of tax level	0.485 [0.383−0.804]
γ_G	Persistence of government spending shock	0.942 [0.922−0.952]
σ_g	Std. dev. of government spending shock	0.097 [0.083−0.118]

Notes: 68 percent credible sets for the estimated parameters are reported in brackets.

the effect of F_t on TFP.[27] These arguments explain why we obtain an estimate of $\tau = 4.196$, which is somewhat higher than what most of the literature has typically contemplated; in fact, Bilbiie et al. (2019) consider values between 0 and 1, while Corsetti et al. (2007) consider a value of 2. However, very little empirical evidence exists about this parameter (Chugh and Ghironi, 2011; Bilbiie et al., 2019). For this reason, in the next subsection we analyze the sensitivity of our findings to different degrees of love of variety, and explain its interplay with the competition effect.

6 Model dynamics

In Figure 3, we report (dashed blue lines) the estimated impulse response functions from the model, computed as the pointwise median of the distribution of impulse responses we obtain. We also report the empirical counterparts from the BVAR model. As the figure shows, we are able to match the sign and shape of the responses of all variables, including the hump-shaped dynamics exhibited by most variables. In terms of magnitudes, the model-implied responses are almost always inside the credible regions from the BVAR model, although some variables tend to respond too strongly, and others too little. Of particular interest is the increase in the house price, which is found to be somewhat smaller in the model than in the data.

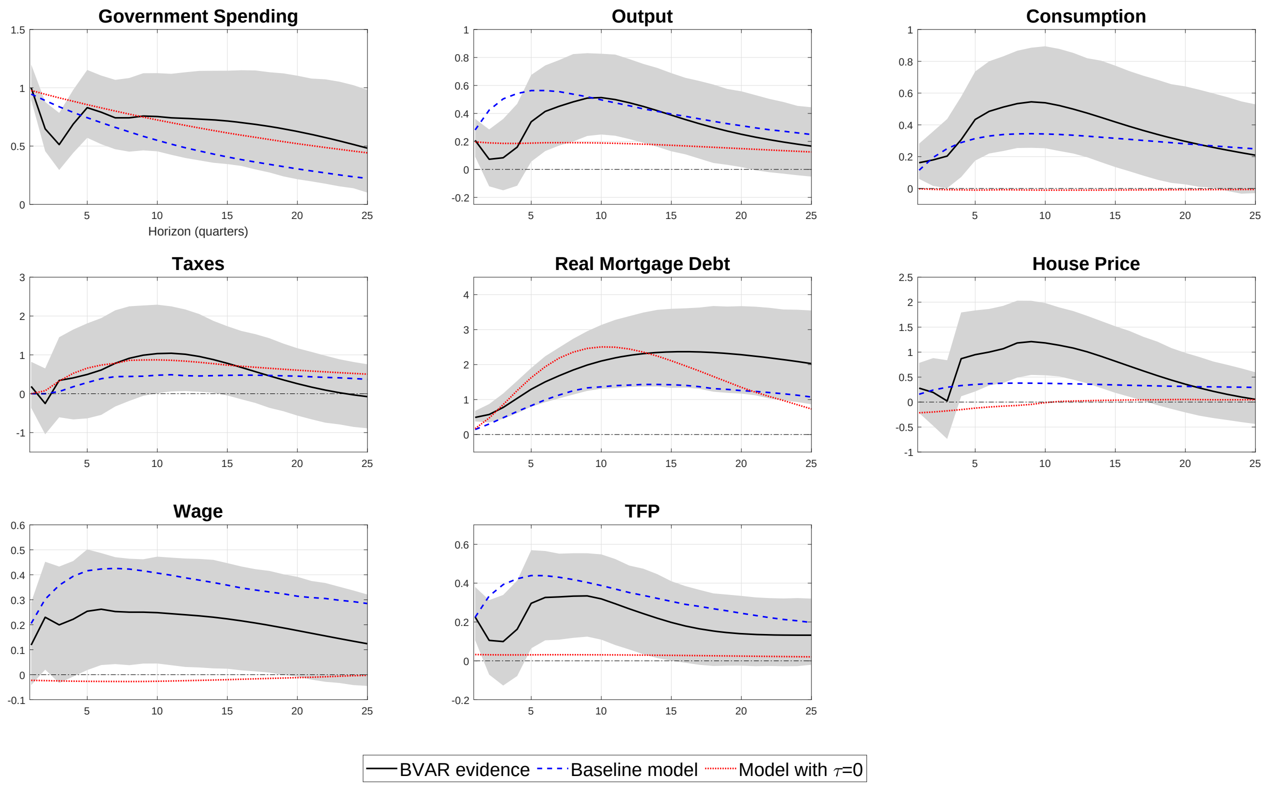
We also estimate a version of the model in which we switch off the taste-for-variety channel, by imposing $\tau = 0$. The estimated impulse responses from this model also appear in Figure 3 (dotted red lines). As the figure makes clear, this model version fails to generate the increase in TFP and the wage rate that are required to obtain a positive response of consumption and the house price. Instead, the responses of these four variables are virtually flat, with the house price even displaying a negative reaction for the first two-three years after the shock. This also implies a counterfactual negative comovement between house prices and mortgage debt over the same time window. Overall, the results from this model echo those of Devereux et al. (1996) and Lewis and Winkler (2017), who show that, for realistic parameter values, firm entry *per se* is not sufficient to crowd-in aggregate consumption.[28]

To gauge the quantitative performance of the model in producing a conditional increase in house prices, Table 2 reports the present-value cumulative fiscal multipliers for the house price that we obtain from the BVAR, the baseline model, and the alternative model (with $\tau = 0$), at

[27]Figure D.5 in Appendix D sheds additional light on the interplay between the taste for variety and the degree of market power by reporting the combinations of these two parameters for which an increase in the house price is obtained. The figure confirms that a lower steady-state markup reduces the minimum value of τ required to obtain an increase in the house price.

[28]In fact, as seen from Table D.1 in Appendix D, the estimate of the steady-state markup, x, is driven to its upper bound of $\frac{1}{\omega} = 1.33$, while the estimated inverse Frisch elasticity, ψ, reaches the lower bound of 0.25 that we impose. This reflects the insights obtained from condition (4.5) in Section 4.

Figure 3: Estimated effects of a government spending shock



Notes: The figure shows the effects of a shock to government spending. Solid black line: BVAR model. Grey areas: 68 percent credible sets from BVAR model. Dashed blue line: Estimated baseline DSGE model. Dotted red line: Estimated DSGE model without taste for variety ($\tau = 0$).

three different horizons (on impact, after two years, and at the end of the 25-quarter horizon considered in Figure 3). As seen from the table, the baseline model accounts for between half and two thirds of the increase in house prices observed in the data, with multipliers well within the credible sets from the BVAR. In contrast, the multipliers of the alternative model without love of variety always remain negative and outside (or on the boundary of) the estimated credible sets.

To inspect the mechanism behind the increase in house prices, Figure 4 reports the response of some selected variables in both the baseline model economy and some alternative economies featuring lower or no taste for variety (keeping all other coefficients at the calibrated/estimated values reported in Table 1).[29] As discussed in Section 3.2.3, the positive response of TFP is magnified by a positive degree of taste for variety, which amplifies the effect on TFP of the increase in the number of firms, as compared with what happens under $\tau = 0$ (despite the

[29]Note that the impulse responses in Figures 4, 5, and D.6 do not correspond exactly to those in Figure 3. The reason is that the former display impulse responses based on the median parameter estimates reported in Table 1, while the latter displays the median impulse responses. As is well known, these two objects do not necessarily coincide. This, however, does not affect any of our qualitative conclusions.

Table 2: Present-value house price fiscal multipliers			
	Impact	8 quarters	25 quarters
Data (BVAR)	0.28 [−0.24;0.65]	0.87 [−0.06;1.40]	0.99 [0.16;1.35]
Baseline model	0.17	0.39	0.55
Alternative model ($\tau = 0$)	−0.22	−0.16	−0.07

Notes: The table reports the present-value cumulative fiscal multipliers for the house price obtained from the BVAR (including 68 percent credible sets in square brackets), along with those from the baseline DSGE model and the alternative model ($\tau = 0$). The multiplier at horizon j is computed as $M_j = \frac{\sum_{i=1}^{j}(1+r)^{-j}\hat{q}_i}{\sum_{i=1}^{j}(1+r)^{-j}\hat{g}_i}$, where $\hat{q}_i$ and $\hat{g}_i$ denote the responses of government spending and the house price at i, respectively. We use the sample average of the 3-month Treasury Bill rate over the period 1966:Q4-2019:Q4, for which the BVAR is estimated. This yields $r = 4.69$.

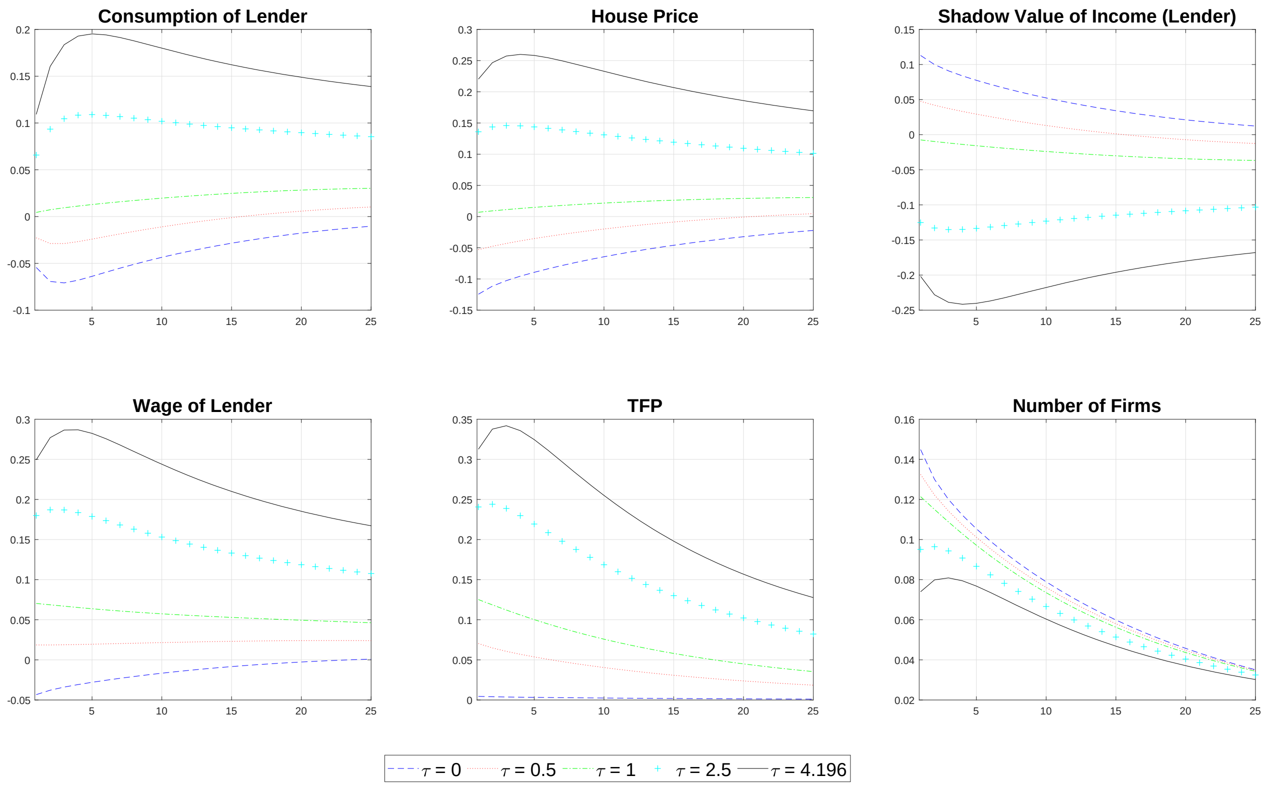
response of the number of firms itself being more modest when τ is high). For a sufficiently high τ, this reflects into an outward shift in the demand for labor that counteracts the drop in labor supply, ultimately leading to a rise in the real wage.[30] Otherwise, under $\tau = 0$ the contraction in labor supply dominates and the real wage drops. The top row of the figure thus confirms the message from the stylized model in Section 4 summarized in condition (4.6): When the taste for variety is sufficiently strong, the model produces a positive response of Ricardian agents' consumption, a decline in their shadow value of income, and thus an increase in the house price.

To dig deeper into the endogenous drivers of the model, we find it useful to consider Figure 5, which reports the impulse responses for different values of the steady-state markup, x, in a setting where the variety effect is shut off by imposing $\tau = 0$. Recall that, for the model to match the data, a large increase in TFP is required—both because TFP itself is found to rise in the BVAR, and because this is crucial in overturning the negative wealth effect, thus producing a positive response of the house price. In the absence of a variety effect, the model solely relies on the competition effect to generate an increase in TFP. To this end, a large drop in the markup is required, as implied by $TFP_t = 1/x_t$. This can be achieved not only through a large increase in the number of firms, but also—*ceteris paribus*—through a high value of the steady-state markup x, which measures the strength of the competition effect, and has a key impact on the relationship between the number of firms and the markup.[31] This can be explained as follows: When the steady-state markup is relatively high, the economy is characterized by poor competition and few firms—each with substantial market power—while fixed costs are high. Consequently, a marginal entrant has a rather large effect on the degree of competition, and thus on the response of the markup to a fiscal shock. By contrast, as the steady-state markup is lowered, the economy approaches perfect

[30] The TFP amplification also reflects into a marked increase in the rental rate of capital, which adds to the upward movement in the real wage, ultimately increasing patient households' income.

[31] As seen from Figure 5, when $\tau = 0$ is imposed in the estimation, the estimated value of $x = 1.33$ implies a relatively weak response of the number of firms.

Figure 4: Effects of a government spending shock for different values of τ

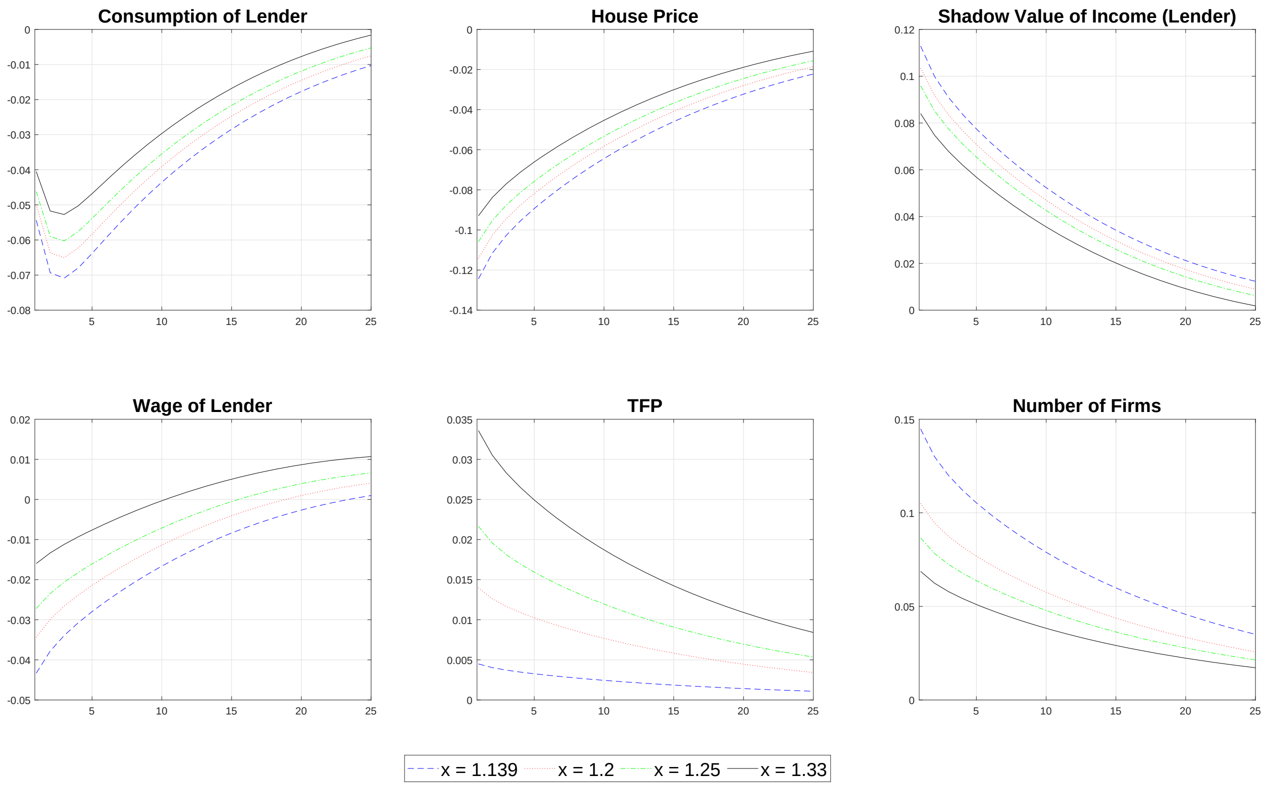


Notes: The figure shows the effects of a shock to government spending for various values of the love-of-variety parameter τ. Dashed blue line: $\tau = 0$. Dotted red line: $\tau = 0.5$. Dashed-dotted green line: $\tau = 1$. Crossed cyan line: $\tau = 2.5$. Solid black line: $\tau = 4.196$ (estimated value). All other parameters are kept at their baseline values.

competition, and the marginal effect of an additional entrant is heavily reduced. To obtain a large response of the markup, and thus a large increase in TFP as seen in the data, the estimation therefore prefers an economy characterized by poor competition, thus returning a high value of x. Even so, as the figure confirms, this is not sufficient to generate an increase in the house price for the range of realistic values of x considered here.

These arguments are turned around once we account for the variety effect. An increase in the number of firms now has a direct positive impact on the TFP response, as discussed in Section 3.2.3, alongside the indirect effect through the markup discussed above. The variety effect relies on a large increase in the number of firms to produce the maximal impact on TFP. This explains why the estimation of our baseline model returns a low value of the steady-state markup: The estimation procedure prefers an environment with strong competition and low entry costs, so that a fiscal shock leads to a large increase in the number of operating firms. This is true despite the fact that a low steady-state markup entails a rather weak competition effect, implying that TFP is only affected by a small decline in the markup. These arguments are confirmed by Figure D.6 in Appendix D, which shows that TFP, the number of firms, and the

Figure 5: Government spending shock for different values of x without variety effect ($\tau = 0$)



Notes: The figure shows the effects of a shock to government spending for various values of the steady-state markup x. Dashed blue line: $x = 1.139$ (estimated value in baseline model). Dotted red line: $x = 1.2$. Dashed-dotted green line: $x = 1.25$. Solid black line: $x = 1.33$ (estimated value in model with $\tau = 0$). The love-of-variety parameter τ has been set to zero. All other parameters are kept at their baseline values.

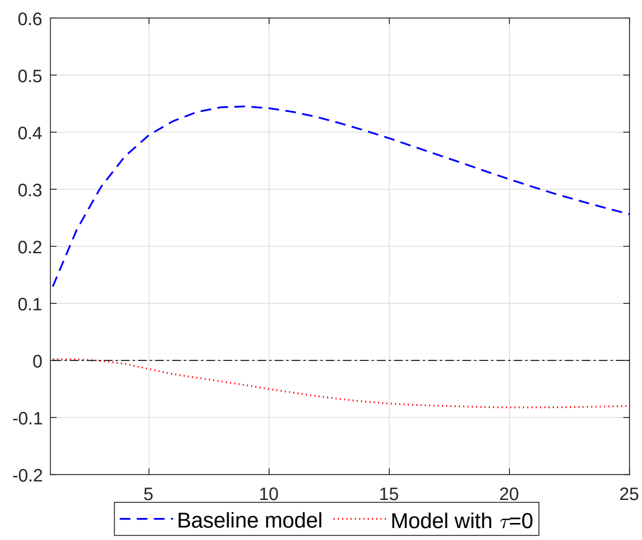
house price increase in tandem when the variety effect operates, and more so the lower x is.

6.1 The consumption response of the borrowers

Our model-based analysis has closely examined the consumption response of patient households, so as to diagnose the reaction of house prices to a government spending shock. We now turn our attention to the consumption response of impatient households. While this response is of limited importance for movements in the house price in the present environment, it contains important insights regarding the transmission of fiscal policy onto aggregate demand. Moreover, it offers a way to validate our model.

The existing empirical literature has produced widespread direct and indirect evidence that credit-constrained households raise their consumption in response to a government spending shock. Alpanda et al. (2021) use data from the U.S. Consumer Expenditure Survey to construct disaggregated time-series data for the consumption of mortgage borrowers, outright homeowners, and renters. In response to an increase in government spending, they find that mortgage borrowers—which are proxied by impatient households in our model—display a significant increase in consumption. Klein (2017) and Bernardini and Peersman (2018) both report that private consumption responds more strongly to fiscal policy during times of high private debt, suggesting that the potential for fiscal policy to crowd in private spending relies to a large extent on the consumption of credit-constrained households. Brinca et al. (2016) report cross-country evidence that fiscal multipliers are positively correlated with wealth inequality, and suggest that the share of consumers facing binding credit constraints may explain this finding.

Figure 6: Response of Impatient Households' Consumption



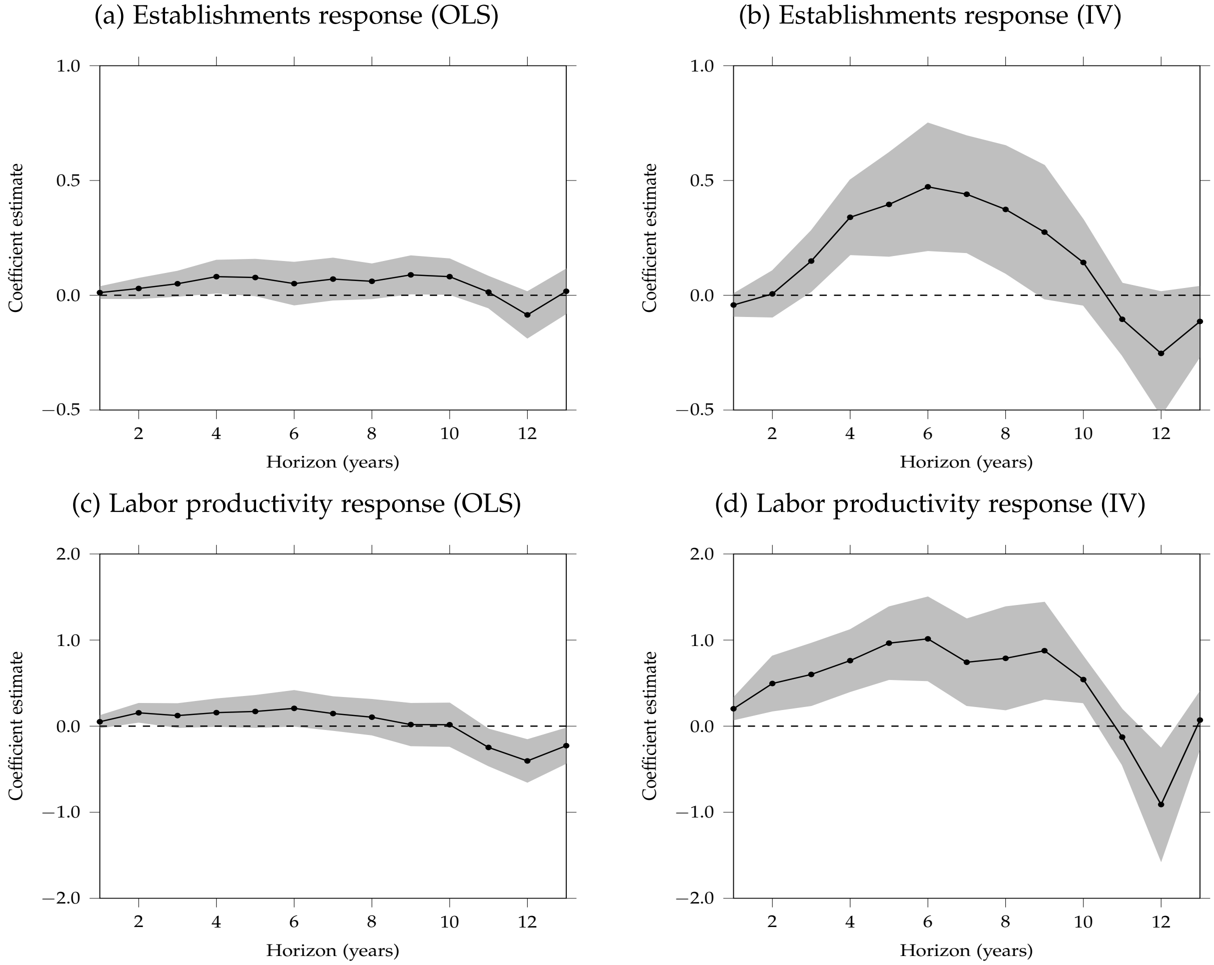
Notes: The figure shows the response of impatient households' consumption to a shock to government spending. Dashed blue line: Estimated baseline DSGE model. Dotted red line: Estimated DSGE model without taste for variety ($\tau = 0$).

We now proceed to evaluate our model's ability to speak to these findings. In Figure 6, we report the consumption response of impatient households to a government spending shock. The message is quite clear: In line with the empirical evidence, impatient households raise their consumption in the baseline model. In contrast, the alternative model without love of variety fails to replicate the positive response of credit-constrained households' consumption, which instead displays a decline. Thus, while the response of impatient households' consumption is not directly targeted in the estimation of the model, our baseline model appears to outperform the alternative model also along this dimension.

7 Validating the mechanism

The procyclical entry of firms, in conjunction with the increase in total and factor-specific productivities, represents the essence of the transmission we have relied on to frame the expansionary effect of fiscal spending on house prices. The regional data employed in Section 2.2 provide us

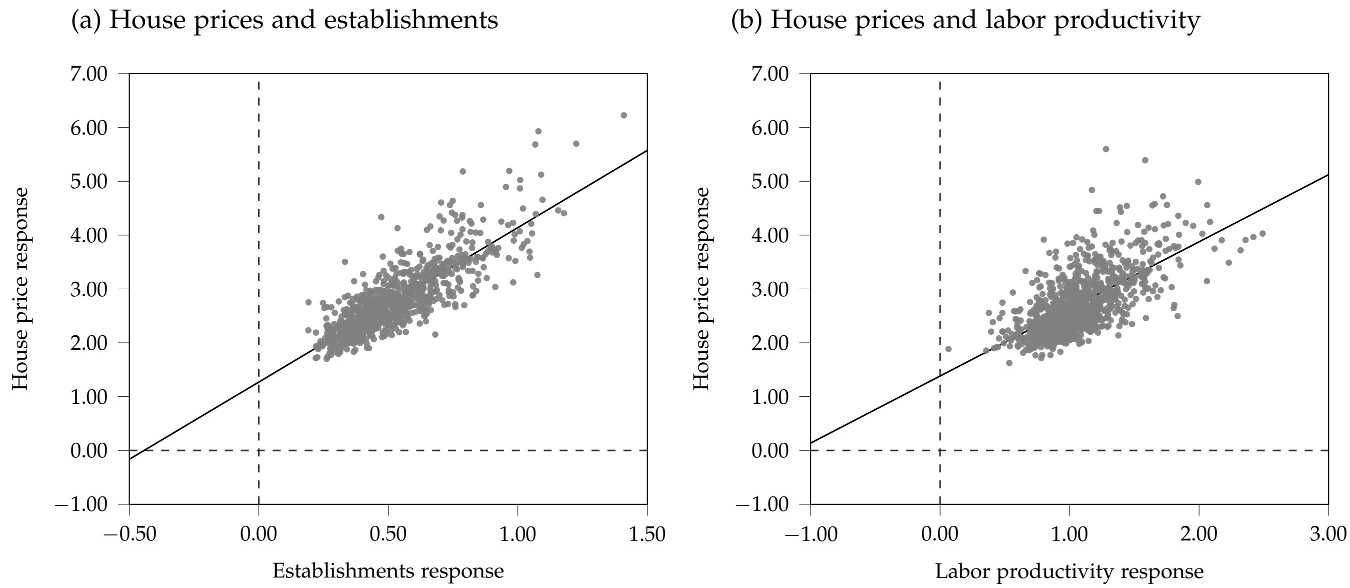
Figure 7: The regional response of establishments and labor productivity to military spending



Notes: The figure shows the estimates of β_h from regression (2.1) based on an annual panel of 380 CBSAs covering the period 2001-2019. The OLS and IV estimates for the establishments response are plotted in panels (a) and (b), while the OLS and IV estimates for the response of labor productivity are shown in panels (c) and (d). As controls, the regression includes two lags of the one-year growth in establishments in panels (a) and (b) and labor productivity in panels (c) and (d), two lags of the instrument, and two lags of the one-year change in local spending normalized by GDP. Grey areas indicate 95 percent confidence bands constructed using heteroskedasticity-robust standard errors clustered by CBSA.

with a unique opportunity to validate this mechanism. Thus, we return to regression (2.1) used to assess the responsiveness of house prices to a change in local DoD spending; this time using (the growth rates of) either local net firm entry or local labor productivity as the dependent variable. Both of these variables are available at the MSA level and for the entire sample we consider. Net firm entry is measured as the growth in the number of establishments within the MSA. Data on establishments are from the County Business Patterns from the U.S. Census Bureau, which contains information on the stock of establishments at the county level. We aggregate these data to get an MSA-level series on establishment counts. Labor productivity is measured as local GDP divided by local employment using data from the BEA.

Figure 8: The joint distributions of house prices, establishments, and labor productivity responses



Notes: This figure shows the estimated joint distribution of the six-year horizon responses of house prices and establishments (left figure), and house prices and labor productivity (right figure). The joint distribution is obtained from bootstrap estimation of regression (2.1) using 1,000 iterations from a cluster bootstrap drawing sets of MSAs with replacement. Each point represents 0.1 percent probability.

Figure 7 presents the estimates for the response of establishments and labor productivity to a change in DoD spending. The OLS and IV estimates for the establishments response are shown in panels (a) and (b), respectively, while the corresponding estimates for the response of labor productivity are shown in panels (c) and (d). Finally, 95 percent confidence bands based on standard errors clustered by MSA are indicated by the grey areas. The IV estimates for the responses of labor productivity and establishments to military spending are in line with our model's predictions. Both increase following a change in government spending and eventually revert back to their local trends, thus validating the central feature of our model. The positive responses of labor productivity and business formation to an increase in local DoD spending are in line with the evidence reported by Auerbach et al. (2019).

Figure 8 conveys further insights into the empirical connection between business formation, labor productivity, and house prices. We report the scatter plots of the impulse-response estimates (at a six-year horizon), based on 1,000 iterations from a cluster bootstrap drawing sets of MSAs from our sample. Panel (a) plots the joint distribution of the responses of house prices and establishments, while the joint distribution of the responses of house prices and labor productivity is shown in panel (b). There is a clear positive relationship across MSAs between the magnitudes of the responses of house prices, establishments, and labor productivity to a change in DoD spending. That is, for the sets of MSAs where the estimated house price response is larger, so are the responses of labor productivity and establishments. This link between response estimates is relatively tight, with pairwise correlation coefficients ranging between 0.7 and 0.8.

8 Concluding remarks

We report new evidence for the U.S. economy indicating that house prices increase following an unanticipated expansion in fiscal spending. We add to the existing time-series evidence pointing to this fact, showing that also the number of establishments and labor productivity increase. This is central to explaining the impact of fiscal spending on house prices, according to a dynamic general equilibrium economy where we combine endogenous entry with a certain degree of taste for variety.

We overcome a longstanding limitation of dynamic frameworks featuring Ricardian households that participate in the housing market. In these economies, fiscal expansions are ultimately responsible for a drop in Ricardian households' nondurable consumption, whose movements are tightly connected to those in house prices, as it is generally the case for any type of shock that does not exert a direct impact on the shadow value of housing (see Barsky et al., 2007). By generating a crowding-in effect on Ricardian households' nondurable consumption and a concurrent drop in their shadow value of income, we are able to induce an increase in house prices following a fiscal expansion.

While our estimated model resolves the house price puzzle emerging in the face of shocks to fiscal spending, the house price response is somewhat weaker than what is observed in the data. In future work, we aim at improving the quantitative account of house price dynamics, also in response to other types of business-cycle perturbations. To do so, we plan to combine our approach—which rests on the role of Ricardian households' consumption choices in pricing housing—with a more recent practice that admits credit-constrained households to exert a non-negligible influence on house prices in equilibrium, in line with the approach of Greenwald and Guren (2021). We leave this challenge for future research.

References

Alpanda, S., Song, H., and Zubairy, S. (2021). Household debt and the effects of fiscal policy. Technical report.

Auerbach, A. J. and Gorodnichenko, Y. (2012). Measuring the Output Responses to Fiscal Policy. *American Economic Journal: Economic Policy*, 4(2):1–27.

Auerbach, A. J., Gorodnichenko, Y., and Murphy, D. (2019). Macroeconomic frameworks. *NBER Working Papers*, 26365.

Auerbach, A. J., Gorodnichenko, Y., and Murphy, D. (2020a). Effects of fiscal policy on credit markets. *AEA Papers and Proceedings*, 110:119–24.

Auerbach, A. J., Gorodnichenko, Y., and Murphy, D. (2020b). Local fiscal multipliers and fiscal spillovers in the USA. *IMF Economic Review*, 68(1):195–229.

Barsky, R. B., House, C. L., and Kimball, M. S. (2007). Sticky-price models and durable goods. *American Economic Review*, 97(3):984–998.

Bartik, T. J. (1991). *Who Benefits from State and Local Economic Development Policies?* W.E. Upjohn Institute.

Baxter, M. and King, R. G. (1993). Fiscal policy in general equilibrium. *The American Economic Review*, 83(3):315–334.

Bayer, C., Born, B., and Luetticke, R. (2020). The liquidity channel of fiscal policy. CEPR Discussion Paper No. DP14883.

Benassy, J.-P. (1996). Taste for variety and optimum production patterns in monopolistic competition. *Economics Letters*, 52(1):41–47.

Bergin, P. R. and Corsetti, G. (2008). The extensive margin and monetary policy. *Journal of Monetary Economics*, 55(7):1222–1237.

Bernardini, M. and Peersman, G. (2018). Private debt overhang and the government spending multiplier: Evidence for the united states. *Journal of Applied Econometrics*, 33(4):485–508.

Bilbiie, F. O. (2011). Nonseparable Preferences, Frisch Labor Supply, and the Consumption Multiplier of Government Spending: One Solution to a Fiscal Policy Puzzle. *Journal of Money, Credit and Banking*, 43(1):221–251.

Bilbiie, F. O. (2021). Monetary Neutrality with Sticky Prices and Free Entry. *The Review of Economics and Statistics*, 103(3):492–504.

Bilbiie, F. O., Fujiwara, I., and Ghironi, F. (2014). Optimal monetary policy with endogenous entry and product variety. *Journal of Monetary Economics*, 64:1–20.

Bilbiie, F. O., Ghironi, F., and Melitz, M. J. (2012). Endogenous entry, product variety, and business cycles. *Journal of Political Economy*, 120(2):304–345.

Bilbiie, F. O., Ghironi, F., and Melitz, M. J. (2019). Monopoly Power and Endogenous Product Variety: Distortions and Remedies. *American Economic Journal: Macroeconomics*, 11(4):140–174.

Bilbiie, F. O., Meier, A., and Müller, G. J. (2008). What accounts for the changes in us fiscal policy transmission? *Journal of Money, Credit and Banking*, 40(7):1439–1470.

Bouakez, H. and Rebei, N. (2007). Why does private consumption rise after a government spending shock? *Canadian Journal of Economics*, 40(3):954–979.

Brinca, P., Holter, H. A., Krusell, P., and Malafry, L. (2016). Fiscal multipliers in the 21st century. *Journal of Monetary Economics*, 77:53–69.

Broda, C. and Weinstein, D. E. (2006). Globalization and the Gains From Variety. *Quarterly Journal of Economics*, 121(2):541–585.

Bureau of Economic Analysis (2015). GDP by Metropolitan Area Methodology.

Chetty, R., Guren, A., Manoli, D., and Weber, A. (2011). Are micro and macro labor supply elasticities consistent? a review of evidence on the intensive and extensive margins. *American Economic Review*, 101(3):471–75.

Christiano, L. J., Eichenbaum, M., and Evans, C. L. (2005). Nominal rigidities and the dynamic effects of a shock to monetary policy. *Journal of Political Economy*, 113(1):1–45.

Chugh, S. and Ghironi, F. (2011). Optimal Fiscal Policy with Endogenous Product Variety. *NBER Working Papers*, 17319.

Corsetti, G., Martin, P., and Pesenti, P. (2007). Productivity, terms of trade and the 'home market effect'. *Journal of International Economics*, 73(1):99–127.

d'Alessandro, A., Fella, G., and Melosi, L. (2019). Fiscal stimulus with learning-by-doing. *International Economic Review*, 60(3):1413–1432.

Devereux, M. B., Head, A. C., and Lapham, B. J. (1996). Monopolistic competition, increasing returns, and the effects of government spending. *Journal of Money, Credit and Banking*, 28(2):233.

Dixit, A. K. and Stiglitz, J. E. (1975). Monopolistic Competition and Optimum Product Diversity. *University of Warwick, Economic Research Papers, No. 64*.

Dixit, A. K. and Stiglitz, J. E. (1977). Monopolistic Competition and Optimum Product Diversity. *American Economic Review*, 67(3):297–308.

Epstein, B., Finkelstein Shapiro, A., and González Gómez, A. (2022). Firm creation, entry costs, and house-price volatility. *Scandinavian Journal of Economics*, forthcoming.

Fernald, J. (2014). A quarterly, utilization-adjusted series on total factor productivity. Working Paper, Federal Reserve Bank of San Francisco.

Ferrero, A. (2015). House Price Booms, Current Account Deficits, and Low Interest Rates. *Journal of Money, Credit and Banking*, 47(S1):261–293.

Galí, J., López-Salido, J. D., and Vallés, J. (2007). Understanding the effects of government spending on consumption. *Journal of the European Economic Association*, 5(1):227–270.

Garriga, C., Kydland, F. E., and Šustek, R. (2021). Monk: Mortgages in a new-keynesian model. *Journal of Economic Dynamics and Control*, 123:104059.

Garriga, C., Manuelli, R., and Peralta-Alva, A. (2019). A macroeconomic model of price swings in the housing market. *American Economic Review*, 109(6):2036–72.

Giannone, D., Lenza, M., and Primiceri, G. E. (2015). Prior selection for vector autoregressions. *Review of Economics and Statistics*, 97(2):436–451.

Greenwald, D. L. and Guren, A. (2021). Do Credit Conditions Move House Prices? *NBER Working Papers*, 29391.

Guerrieri, L. and Iacoviello, M. (2017). Collateral constraints and macroeconomic asymmetries. *Journal of Monetary Economics*, 90:28–49.

Guren, A. M., McKay, A., Nakamura, E., and Steinsson, J. (2021). Housing wealth effects: The long view. *The Review of Economic Studies*, 88(2):669–707.

Hall, R. E. (2009). By how much does gdp rise if the government buys more output? *Brookings Papers on Economic Activity*, 40(2):183–249.

Hazell, J., Herreno, J., Nakamura, E., and Steinsson, J. (2022). The slope of the phillips curve: Evidence from us states. *The Quarterly Journal of Economics*, 137(3):1299–1344.

Iacoviello, M. (2005). House Prices, Borrowing Constraints, and Monetary Policy in the Business Cycle. *American Economic Review*, 95(3):739–764.

Iacoviello, M. and Neri, S. (2010). Housing Market Spillovers: Evidence from an Estimated DSGE Model. *American Economic Journal: Macroeconomics*, 2(2):125–164.

Jaimovich, N. (2007). Firm dynamics and markup variations: Implications for sunspot equilibria and endogenous economic fluctuations. *Journal of Economic Theory*, 137(1):300–325.

Jaimovich, N. and Floetotto, M. (2008). Firm dynamics, markup variations, and the business cycle. *Journal of Monetary Economics*, 55(7):1238–1252.

Jensen, H., Ravn, S. H., and Santoro, E. (2018). Changing credit limits, changing business cycles. *European Economic Review*, 102:211–239.

Jordà, O., Schularick, M., and Taylor, A. M. (2015). Betting the house. *Journal of International Economics*, 96:S2–S18.

Jørgensen, P. L. and Ravn, S. H. (2022). The inflation response to government spending shocks: A fiscal price puzzle? *European Economic Review*, 141:103982.

Justiniano, A., Primiceri, G. E., and Tambalotti, A. (2019). Credit supply and the housing boom. *Journal of Political Economy*, 127(3):1317–1350.

Khan, H. and Reza, A. (2017). House prices and government spending shocks. *Journal of Money, Credit and Banking*, 49(6):1247–1271.

Kiyotaki, N. and Moore, J. (1997). Credit cycles. *Journal of Political Economy*, 105(2):211–248.

Klein, M. (2017). Austerity and private debt. *Journal of Money, Credit and Banking*, 49(7):1555–1585.

Leeper, E. M., Traum, N., and Walker, T. B. (2017). Clearing Up the Fiscal Multiplier Morass. *American Economic Review*, 107(8):2409–2454.

Leeper, E. M., Walker, T. B., and Yang, S.-C. S. (2010). Government investment and fiscal stimulus. *Journal of Monetary Economics*, 57(8):1000–1012.

Lewis, V. and Poilly, C. (2012). Firm entry, markups and the monetary transmission mechanism. *Journal of Monetary Economics*, 59(7):670–685.

Lewis, V. and Winkler, R. (2017). Government spending, entry, and the consumption crowding-in puzzle. *International Economic Review*, 58(3):943–972.

Liu, Z., Wang, P., and Zha, T. (2013). Land-price dynamics and macroeconomic fluctuations. *Econometrica*, 81(3):1147–1184.

Montiel Olea, J. L. and Pflueger, C. (2013). A Robust Test for Weak Instruments. *Journal of Business & Economic Statistics*, 31(3):358–369.

Nakamura, E. and Steinsson, J. (2014). Fiscal Stimulus in a Monetary Union: Evidence from US Regions. *American Economic Review*, 104(3):753–792.

Ramey, V. A. (2016). Macroeconomic shocks and their propagation. *Handbook of Macroeconomics*, 2:71–162.

Ramey, V. A. and Zubairy, S. (2018). Government Spending Multipliers in Good Times and in Bad: Evidence from US Historical Data. *Journal of Political Economy*, 126(2):850–901.

Rotemberg, J. J. and Woodford, M. (1992). Oligopolistic pricing and the effects of aggregate demand on economic activity. *Journal of Political Economy*, 100(6):1153–1207.

Saiz, A. (2010). The Geographic Determinants of Housing Supply. *Quarterly Journal of Economics*, 125(3):1253–1296.

Stock, J. H. and Watson, M. W. (2018). Identification and Estimation of Dynamic Causal Effects in Macroeconomics Using External Instruments. *The Economic Journal*, 128(610):917–948.

A Appendix to the empirical analysis

This appendix contains additional details on the data used in the VAR model and the cross-MSA analyses, as well as some robustness checks.

A.1 Appendix to the BVAR analysis

We estimate a BVAR model with four lags and a constant. The model can be stated as:

$$\mathbf{X}_t = \Theta + B^{-1}A(L)\mathbf{X}_{t-1} + B^{-1}e_t, \tag{A.1}$$

where $\mathbf{X}_t$ is the vector of endogenous variables, e_t is a vector of i.i.d. structural shocks with unit variance, $A(L)$ comprises the coefficients on the lagged endogenous variables, L is the lag operator, and B denotes the coefficients on the contemporaneous endogenous variables. The list of variables and their ordering is the following:

$$\mathbf{X}_t = \begin{bmatrix} FE_t & G_t & Y_t & C_t & T_t & B_t & Q_t & W_t & TFP_t \end{bmatrix}'.$$

This ordering reflects our identification strategy: The forecast errors are ordered first in the system, as these are assumed to be orthogonal to the economy in the sense that they do not respond to any of the other variables within-quarter. This allows us to recover a truly unexpected shock to government spending. We follow Auerbach and Gorodnichenko (2012) and order government spending immediately after FE_t, while the ordering of the remaining variables is not of importance for the results.

Most of the data used in the baseline specification of our BVAR model are taken from the Federal Reserve Economic Data (FRED) database. The series are described in detail below, with series names in FRED indicated in brackets. The only exceptions are the forecast errors of Auerbach and Gorodnichenko (2012) and the TFP series of Fernald (2014).

G_t: Real Government Consumption Expenditures and Gross Investment (GCEC1, seasonally adjusted, Chained 2009 $).

Y_t: Real Gross Domestic Product (GDPC1, seasonally adjusted, Chained 2009 $).

C_t: Real Personal Consumption Expenditures (PCECC96, seasonally adjusted, Chained 2009 $).

T_t: Government current tax receipts (W054RC1Q027SBEA) + Government income receipts on assets (W058RC1Q027SBEA) + Government current transfer receipts (W060RC1Q027SBEA) - Government current transfer payments (A084RC1Q027SBEA) - Government interest payments (A180RC1Q027SBEA) - Government subsidies (GDISUBS).[32] All series are seasonally adjusted. We convert from nominal to real terms using the GDP deflator (GDPDEF).

[32]Since the series turns negative at some points in time, we add a constant to it before taking logs.

B_t: Home mortgages (liabilities) of households and nonprofit organizations from the Flow of Funds (HMLBSHNO). We convert the series to real terms using the GDP deflator.

Q_t: Median Sales Price of Houses Sold for the United States (MSPUS). We convert the series to real terms using the GDP deflator.

W_t: Real Compensation Per Hour in the Nonfarm Business Sector (COMPRNFB, Seasonally Adjusted, 2012=100).

TFP_t: Raw (non-utilization-adjusted) Total Factor Productivity series of Fernald (2014). The data can be collected from https://www.frbsf.org/economic-research/indicators-data/total-factor-productivity-tfp/

The first five series are converted to per capita terms using the Census Bureau Civilian Population (All Ages) estimates, which we also collect from the FRED database (POP). We then take logs of all variables, and detrend them using a linear and a quadratic trend.

Finally, we use the following series of "narrative" shocks to government spending:

FE_t: Forecast error of government spending, computed as the difference between forecasts (obtained from the Greenbook data of the Federal Reserve Board combined with the Survey of Professional Forecasters) and the actual, first-release data for the growth rate of government spending. We use the updated series covering the entire sample 1966:Q4-2019:Q4, which we obtain from Jørgensen and Ravn (2022).

A.2 Appendix to the regional analysis

A.2.1 Data used in the regional analysis

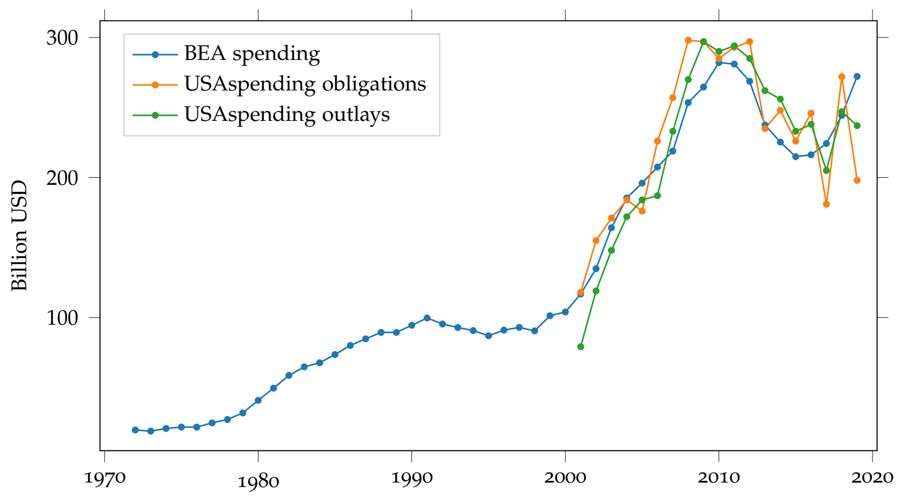
We collect data on contracts signed by firms with the Department of Defense from USAspending.gov to construct the data used in Section 2. The data cover all DoD prime contracts signed from 2001 through 2019, including terminated contracts. The dataset does not contain information on the timing of actual outlays to contractors but it does contain information on the duration and total dollar amount obligated per contract. Additionally, the dataset contains the name of the contractor and the primary place of work performance at the ZIP code level.

The raw data are cleaned using the same approach as Auerbach et al. (2020b). First, we match a terminated contract with its original contract if a de-obligated dollar amount falls within 0.5 percent of dollars obligated in another contract, and both contracts have the same contractor ID and ZIP code. These matched obligations and de-obligations are removed from the dataset. Second, we remove long-term contracts that terminate after our sample period by removing all contracts that terminate after 2023.

Our baseline estimates use variation in obligations rather than actual outlays. This assigns the entire obligated amount to the first year of the contract. As a robustness check, we construct a proxy for outlays per contract by dividing the dollars obligated in each contract evenly among the months of the contract's duration. We then sum these amounts annually by MSA to get a proxy for total annual outlays to the MSAs.

Our data track official data on national military spending from the BEA well in terms of both magnitude and movements. This is seen in Figure A.1, which plots national obligations and our proxy for outlays according to the data from USAspending.gov, together with intermediate goods and services purchased for national defense from the BEA's NIPA tables.

Figure A.1: Military spending according to USAspending.gov and BEA data



Notes: The blue line is "Intermediate goods and services purchased" in the BEA's NIPA Table 3.11.5, "National Defense Consumption Expenditures and Gross Investment by Type". Orange and green lines are annual obligations and outlays constructed using USAspending.gov data.

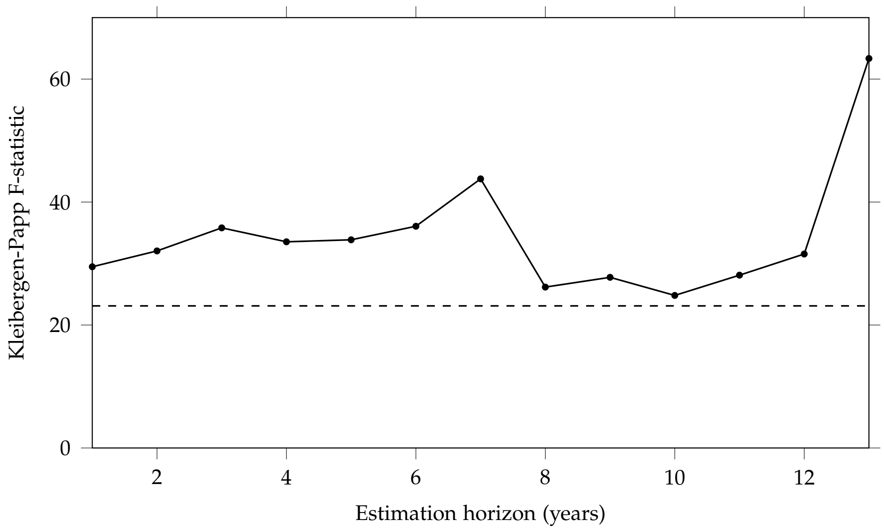
A.2.2 First-stage estimates

Figure A.2 shows the Kleibergen-Papp F-statistics over different estimation horizons from the first-stage regression

$$\frac{G_{i,t+1} - G_{i,t}}{Y_{i,t}} = \tilde{\alpha}_{i,h} + \tilde{\eta}_{t+h} + \tilde{\beta}_h \bar{G}_i \times \frac{G^{nat}_{t+1} - G^{nat}_{t}}{Y_{i,t}} + \tilde{\gamma}_h X_{i,t} + \epsilon_{i,t+1}. \tag{A.2}$$

We only show the F-statistics from the first stage to the regression with house price growth as dependent variable. This regression only differs from the first stage to the regressions with establishments or labor productivity growth in that it has two lags of house price growth as controls, rather than two lags of establishments or labor productivity growth. Although not shown, the F-statistics from these three sets of first-stage regressions are almost identical.

Figure A.2: F-statistics from first-stage regression



Notes: The figure shows the Kleibergen-Papp F-statistics from the first-stage regression (A.2) over different estimation horizons. Heteroskedasticity-robust standard errors are clustered by MSA. The dashed line indicates the Montiel Olea and Pflueger (2013) critical value for the F-statistic under a null hypothesis of the IV bias exceeding 10% of the OLS bias at the 5% significance level.

A.2.3 Pre-trend analysis of the house price estimates

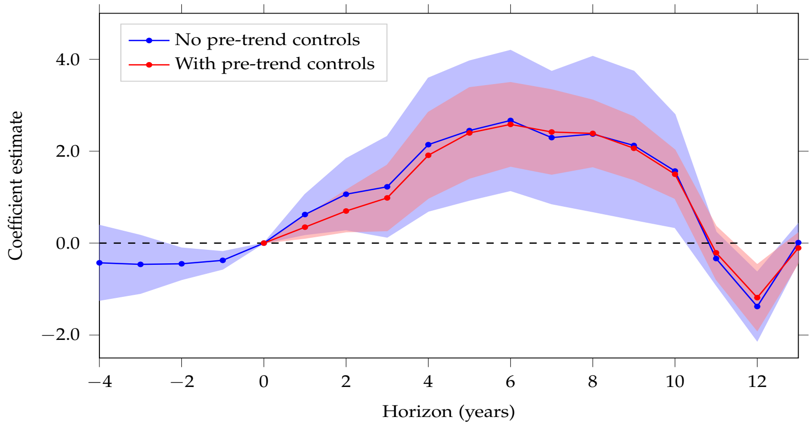
This section presents a pre-trend analysis of our regional estimates of the house price response in Section 2. The blue line in Figure A.3 shows the estimates of the house price response from a modification of our baseline regression in which we do not control for lags of house price growth and extend the estimation horizon to four years prior to the spending change (that is, for horizons $h = -4, -3, \ldots, 13$). The red line plots the estimates from our baseline regression.

The blue line indicates that house prices are slightly, albeit significantly, lower in the two years before a change in military spending. Reassuringly, we see that the estimates of the house price response after a spending change are similar for both regressions. Hence, whether or not we control for the relatively small pre-trend in house prices matters little for our results.

A.2.4 Endogeneity analysis of the estimates

This section presents an endogeneity analysis of our regional estimates of the house price response in Section 2. We test the null hypothesis of the endogenous regressor in the IV regression being exogenous using a Hausman test (i.e. testing that the IV and OLS estimates are identical). The resulting p-values are very close to zero up until year 11 as shown in Figure A.4, supporting the need for using an IV approach.

Figure A.3: Pre-trend analysis of house price responses to military spending



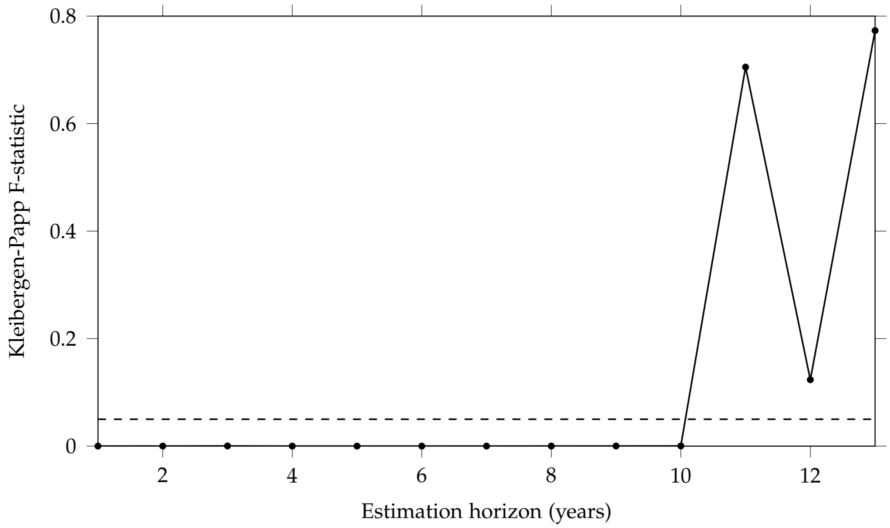
Notes: The figure shows the IV estimates of β_h from regression (2.1) based on an annual panel of 380 MSAs covering the period 2001-2019. The red line shows the baseline specification, which include as controls two lags of the one-year growth in house prices, two lags of the instrument, and two lags of the one-year change in local spending normalized by GDP. The blue line plots the estimates from the same regression except that the two lags of the one-year growth in house prices are not included in the set of controls. Red and blue areas indicate the 95 percent confidence bands constructed using heteroskedasticity-robust standard errors clustered by MSA.

A.2.5 Robustness of the regional house price estimates

This section analyzes the robustness of our regional estimates of the house price response in Section 2 to alternative specifications and potential outliers. We also show the sensitivity of the estimates to the inclusion of control variables.

Table A.1 shows the IV estimates from alternative specifications of regression (2.1). Estimates from the baseline specification also shown in Figure 2 are presented in column (1). Column (2) shows the estimates from a regression in which house prices, DoD spending, and GDP have been deflated by the MSA-level GDP deflator. Column (3) reports estimates from a regresion in which we use the proxy for outlays described in Appendix A.2.1 to measure DoD spending. Columns (4) and (5) present estimates with alternative normalizations of DoD spending (by personal income and population in thousand persons, respectively). Column (6) controls for house price movements associated with industry composition, by adding to the regression 2-digit industry employment shares multiplied by year dummies. Column (7) controls for differential exposure to aggregate house price movements by adding to the regression three time-invariant controls multiplied by year dummies: the Wharton Regulation Index, the Saiz (2010) instrument, and the Bartik-like instrument for sensitivity to regional house price movements by Guren et al. (2021). Column (8) adds state × year fixed effects to control for state-specific house price growth fluctuations. Finally, column (9) reports estimates from a version of the regression in which the instrument is constructed using the average DoD spending share over 2001 and 2002. By using beginning-of-sample spending shares, we avoid endogeneity issues that could potentially stem

Figure A.4: p values from endogeneity test



Notes: The figure shows robust p-values from Hausman tests of endogeneity over different estimation horizons. The null hypothesis is that the endogeneous regressor can be treated as exogenous. Dashed line indicates 5 % confidence level.

from the within-sample DoD spending shares being correlated with economic fluctuations.

Controlling for local industry composition or differential housing exposure tends to reduce the size of the house price response. Assuringly, the estimates are still statistically significant. We also want to highlight the estimates from the specification using the proxy for outlays and the specification controlling for state $\times$ year fixed effects. The latter only uses within-state variation and reduces estimates but they are still significant and display a hump-shaped pattern. When using the proxy for outlays instead of obligations, the estimates become substantially larger.

Next, we show in Table A.2 the sensitivity of our estimates to the outliers. The baseline estimates are shown in column (1). In column (2) we remove all MSAs in the bottom and top 5th percentiles of the distribution of DoD spending shares used to construct the instrument. Column (3) reports estimates from a regression in which we use the non-winsorized change in local spending. Finally, column (4) shows the estimates when we remove all winsorized observations. Using non-winsorized changes in local spending has little impact on the estimates, while removing MSAs in the top and bottom DoD spending share distribution or winsorized observations amplifies the house price response.

Finally, we show the sensitivity of the estimates to the inclusion of control variables in Table A.3. Column (1) shows estimates from regressions without any controls. Adding lagged spending and instruments lowers the estimates across the entire horizon, as seen in column (2). This suggests that the lag exogeneity condition is not met unless we condition on past values of both spending growth and the instrument itself. When lags of the dependent variable are added as controls, this has limited impact on the estimates, as shown in column (3). Adding lags of the

Table A.1: Robustness of regional house price estimates (alternative specifications)

	(1) Baseline	(2) Real variables	(3) Outlays	(4) Normalize by income	(5) Normalize by population	(6) Control for industry comp.	(7) Control for housing exposure	(8) Control for state	(9) Pre-sample shares
	Dependent variable: House price growth								
1-year	0.35***	0.35**	0.61***	0.29***	0.0064**	0.34**	0.18***	0.23***	0.53**
	(0.12)	(0.14)	(0.16)	(0.10)	(0.00)	(0.16)	(0.05)	(0.07)	(0.21)
2-year	0.70***	0.68***	1.23***	0.59***	0.013***	0.64**	0.41***	0.51***	0.99***
	(0.23)	(0.25)	(0.32)	(0.20)	(0.00)	(0.30)	(0.11)	(0.15)	(0.35)
4-year	1.91***	1.89***	3.29***	1.52***	0.039***	1.60***	1.17***	1.42***	2.64***
	(0.48)	(0.50)	(0.63)	(0.39)	(0.01)	(0.52)	(0.32)	(0.36)	(0.86)
6-year	2.58***	2.44***	4.20***	2.04***	0.055***	2.28***	1.40***	1.79***	3.24***
	(0.47)	(0.48)	(0.63)	(0.38)	(0.01)	(0.56)	(0.52)	(0.39)	(0.92)
10-year	1.50***	1.18***	3.34***	1.11***	0.027***	1.44***	0.78**	1.02***	1.74***
	(0.27)	(0.29)	(0.57)	(0.21)	(0.01)	(0.33)	(0.31)	(0.23)	(0.46)
MSAs	380	380	380	380	380	380	255	373	380

Notes: The table presents the IV estimates from alternative specifications of regression (2.1). Column (1) presents the baseline estimates. Column (2) shows the estimates when house prices, DoD spending, and GDP are deflated by the MSA-level GDP deflator. Column (3) uses DoD spending measured by the outlay proxy described in Appendix A.2.1. Column (4) normalizes DoD spending by the BEA's measure of personal income. Column (5) normalizes DoD spending by the BEA's measure of population (in thousand persons). Column (6) adds year dummies multiplied the average two-digit industry employments shares over the sample period. The employment shares are calculated using data from the Census Bureau's County Business Patterns. Column (7) adds year dummies interacted with three time-invariant measures of exposure to aggregate house price flucuations (the Wharton Regulation Index, the Saiz (2010) instrument and the Guren et al. (2021) instrument). This reduces the sample size since the Wharton Regulation Index and the Saiz (2010) instrument are not available for all MSAs. Column (8) adds state $\times$ year fixed effects. Column (9) uses pre-sample (2001-2002) DoD spending shares to construct the instrument. Heteroskedasticity-robust standard errors clustered by MSA are shown in parentheses. ***, **, and * denote significance at the 0.01, 0.05 and 0.1 level, respectively.

dependent variables to the regression with lagged spending and instruments—thereby retrieving our baseline results—reduces the standard errors (see column (4)), thus improving the efficiency of the estimators.

A.2.6 Robustness of the regional establishments and labor productivity estimates

This section reports some robustness exercises regarding the responses of labor productivity and the number of establishments. Table A.4 reports the IV estimates from the same alternative specifications as those reported in Table A.1. The estimates that differ from the baseline are those from the specification controlling for housing exposure, which are smaller than the baseline and in many cases insignificant; see column (7). For the labor productivity response, this seems to be driven by the model being estimated on the subsample of MSAs for which we have data

Table A.2: Robustness of regional house price estimates (outliers)

	(1) Baseline	(2) Remove extreme DoD shares	(3) Non-winsorized	(4) Remove winsorized
	Dependent variable: House price growth			
1-year	0.35^{***}	0.86^{***}	0.38^{***}	0.92^{***}
	(0.12)	(0.23)	(0.09)	(0.18)
2-year	0.70^{***}	1.61^{***}	0.72^{***}	1.51^{***}
	(0.23)	(0.41)	(0.17)	(0.30)
4-year	1.91^{***}	4.93^{***}	1.91^{***}	3.90^{***}
	(0.48)	(1.01)	(0.40)	(0.63)
6-year	2.58^{***}	6.22^{***}	2.42^{***}	4.59^{***}
	(0.47)	(1.35)	(0.44)	(0.75)
10-year	1.50^{***}	3.50^{***}	1.49^{***}	2.35^{***}
	(0.27)	(0.82)	(0.45)	(0.44)
MSAs	380	342	380	379

Notes: The table presents the IV estimates from regression (2.1). Column (1) presents the baseline estimates. Column (2) shows the estimates when removing MSAs in the bottom and top 5th percentiles of the distribution of average DoD spending shares used to construct the instrument. Column (3) presents estimates when the cumulative change in DoD spending is not winsorized. Column (4) removes all winsorized observations. Heteroskedasticity-robust standard errors clustered by MSA are shown in parentheses. ***, **, and * denote significance at the 0.01, 0.05 and 0.1 level, respectively.

Table A.3: Robustness of regional house price estimates (controls)

	(1)	(2)	(3)	(4)
	Dependent variable: House price growth			
1-year	0.49^{***}	0.62^{***}	0.30^{***}	0.35^{***}
	(0.14)	(0.22)	(0.11)	(0.12)
2-year	1.15^{***}	1.06^{***}	0.79^{***}	0.70^{***}
	(0.26)	(0.40)	(0.18)	(0.23)
4-year	3.56^{***}	2.14^{***}	3.09^{***}	1.91^{***}
	(0.62)	(0.74)	(0.51)	(0.48)
6-year	4.77^{***}	2.67^{***}	4.33^{***}	2.58^{***}
	(0.89)	(0.78)	(0.71)	(0.47)
10-year	4.05^{***}	1.57^{**}	3.36^{***}	1.50^{***}
	(0.94)	(0.63)	(0.54)	(0.27)
Control for lagged spending and instruments	No	Yes	No	Yes
Control for lagged dependent variable	No	No	Yes	Yes

Notes: The table presents the IV estimates from regression (2.1). Column (1) shows results from regressions without any controls. Column (2) adds two lags of the change in spending and the instrument. Column (3) adds two lags of the one-period growth in house prices. Column (4) adds both sets of controls. Heteroskedasticity-robust standard errors clustered by MSA are shown in parentheses. ***, **, and * denote significance at the 0.01, 0.05 and 0.1 level, respectively.

on exposure to aggregate house price fluctuations. If we estimate the model on this subsample but do not control for differential exposure, the results are broadly similar to those in column

(7). For the establishments response, the results reflect a combination of the reduced number of

Table A.4: Robustness of regional establishments and labor productivity estimates (alternative specifications)

	(1) Baseline	(2) Real variables	(3) Outlays	(4) Normalize by income	(5) Normalize by population	(6) Control for industry comp.	(7) Control for housing exposure	(8) Control for state	(9) Pre-sample shares
	Dependent variable: Establishment growth								
1-year	-0.04*	-0.04*	-0.078	-0.031	-0.001	-0.08*	-0.055	-0.041	-0.09
	(0.03)	(0.03)	(0.06)	(0.02)	(0.00)	(0.04)	(0.04)	(0.03)	(0.08)
2-year	0.006	0.011	0.069	0.011	0.0002	-0.023	-0.024	0.0006	-0.02
	(0.05)	(0.05)	(0.12)	(0.04)	(0.00)	(0.08)	(0.06)	(0.04)	(0.08)
4-year	0.34***	0.36***	0.61***	0.28***	0.0069***	0.34***	0.095	0.34***	0.37***
	(0.08)	(0.09)	(0.14)	(0.07)	(0.00)	(0.11)	(0.09)	(0.08)	(0.12)
6-year	0.47***	0.50***	0.69***	0.37***	0.0091***	0.46***	0.11	0.42***	0.46**
	(0.14)	(0.15)	(0.17)	(0.11)	(0.00)	(0.15)	(0.15)	(0.12)	(0.18)
10-year	0.14	0.14	0.54***	0.100	0.0015	0.17*	0.093	0.12*	0.23
	(0.10)	(0.10)	(0.14)	(0.07)	(0.00)	(0.10)	(0.07)	(0.06)	(0.16)
	Dependent variable: Labor productivity growth								
1-year	0.20***	0.19**	0.37***	0.16***	0.0055***	0.26***	0.17**	0.042	0.08
	(0.07)	(0.07)	(0.13)	(0.05)	(0.00)	(0.09)	(0.08)	(0.05)	(0.10)
2-year	0.50***	0.42***	0.68***	0.38***	0.011***	0.39**	0.46***	0.25***	0.50***
	(0.16)	(0.15)	(0.23)	(0.13)	(0.00)	(0.19)	(0.15)	(0.08)	(0.18)
4-year	0.76***	0.62***	1.12***	0.60***	0.017***	0.56**	0.45*	0.31**	0.72***
	(0.18)	(0.17)	(0.22)	(0.15)	(0.01)	(0.25)	(0.26)	(0.15)	(0.27)
6-year	1.01***	0.91***	1.47***	0.79***	0.021***	0.88***	0.38	0.37*	0.83**
	(0.25)	(0.20)	(0.28)	(0.21)	(0.01)	(0.31)	(0.30)	(0.21)	(0.35)
10-year	0.54***	0.41***	1.53***	0.36***	0.011***	0.39	0.37*	0.29**	0.69***
	(0.14)	(0.14)	(0.44)	(0.11)	(0.00)	(0.25)	(0.22)	(0.14)	(0.23)
MSAs	380	380	380	380	380	380	255	373	380

Notes: The table presents the IV estimates from alternative specifications of regression (2.1). Column (1) presents the baseline estimates. Column (2) shows the estimates when house prices, DoD spending, and GDP are deflated by the MSA-level GDP deflator. Column (3) uses DoD spending measured by the outlay proxy described in Appendix A.2.1. Column (4) normalizes DoD spending by the BEA's measure of personal income. Column (5) normalizes DoD spending by the BEA's measure of population (in thousand persons). Column (6) adds year dummies multiplied the average two-digit industry employment shares over the sample period. The employment shares are calculated using data from the Census Bureau's County Business Patterns. Column (7) adds year dummies interacted with three time-invariant measures of exposure to aggregate house price flucuations (the Wharton Regulation Index, the Saiz (2010) instrument and the Guren et al. (2021) instrument). This reduces the sample size since the Wharton Regulation Index and the Saiz (2010) instrument are not available for all MSAs. Column (8) adds state × year fixed effects. Column (9) uses pre-sample (2001-2002) DoD spending shares to construct the instrument. Heteroskedasticity-robust standard errors clustered by MSA are shown in parentheses. ***, **, and * denote significance at the 0.01, 0.05 and 0.1 level, respectively.

MSAs in the regression and the fact that, unconditionally, local establishment growth is positively correlated with local house price growth (see also Epstein et al., 2022), which in turn is closely related to the measures of house price exposure that we control for. In contrast, there are no statistically significant pairwise correlations between the house price exposure measures and our instrument, which rules out potential endogeneity concerns stemming from differential exposure to aggregate house price fluctuations.

Turning to the sensitivity to outliers, Table A.5 shows the sensitivity of the labor productivity and establishments response estimates to outliers using the same checks as those in Table A.2. The results are similar to those from the house price response estimates.

Table A.5: Robustness of regional establishments and labor productivity estimates (outliers)

	(1) Baseline	(2) Remove extreme DoD shares	(3) Non-winsorized	(4) Remove winsorized
	Dependent variable: Establishment growth			
1-year	-0.042*	0.014	-0.040	-0.078
	(0.03)	(0.08)	(0.03)	(0.07)
2-year	0.0060	0.15	0.0095	0.065
	(0.05)	(0.11)	(0.05)	(0.09)
4-year	0.34***	0.73***	0.33***	0.48**
	(0.08)	(0.26)	(0.08)	(0.20)
6-year	0.47***	1.23***	0.41***	0.80***
	(0.14)	(0.27)	(0.13)	(0.23)
10-year	0.14	0.55***	0.17*	0.36**
	(0.10)	(0.18)	(0.09)	(0.16)
	Dependent variable: Labor productivity growth			
1-year	0.20***	0.45**	0.22***	0.25*
	(0.07)	(0.20)	(0.06)	(0.13)
2-year	0.50***	1.13***	0.48***	0.95***
	(0.16)	(0.29)	(0.14)	(0.22)
4-year	0.76***	1.94***	0.72***	1.28***
	(0.18)	(0.41)	(0.18)	(0.24)
6-year	1.01***	2.96***	0.89***	1.75***
	(0.25)	(0.81)	(0.24)	(0.44)
10-year	0.54***	0.99**	0.38**	1.00***
	(0.14)	(0.50)	(0.17)	(0.33)
MSAs	380	342	380	379

Notes: The table presents the IV estimates from regression (2.1). Column (1) presents the baseline estimates. Column (2) shows the estimates when removing MSAs in the bottom and top 5th percentiles of the distribution of average DoD spending shares used to construct the instrument. Column (3) presents estimates when the cumulative change in DoD spending is not winsorized. Column (4) removes all winsorized observations. Heteroskedasticity-robust standard errors clustered by MSA are shown in parentheses. ***, **, and * denote significance at the 0.01, 0.05 and 0.1 level, respectively.

Table A.6 reports the sensitivity of our baseline estimates to the control variables. The main take-aways from this exercise are in line with those from Table A.3. While controlling for lags of spending growth and of the instrument generally lowers the response estimates, adding lags of the dependent variables tends to reduce the standard errors.

Table A.6: Robustness of regional establishments and labor productivity estimates (controls)

	(1)	(2)	(3)	(4)
	Dependent variable: Establishment growth			
1-year	0.018	-0.020	0.0086	-0.042*
	(0.02)	(0.02)	(0.02)	(0.03)
2-year	0.079*	0.029	0.070	0.0060
	(0.05)	(0.05)	(0.04)	(0.05)
4-year	0.53***	0.34***	0.52***	0.34***
	(0.15)	(0.08)	(0.15)	(0.08)
6-year	0.65***	0.43***	0.65***	0.47***
	(0.20)	(0.14)	(0.21)	(0.14)
10-year	0.46**	0.13	0.44**	0.14
	(0.19)	(0.11)	(0.18)	(0.10)
	Dependent variable: Labor productivity growth			
1-year	0.13**	0.20***	0.13**	0.20***
	(0.06)	(0.06)	(0.06)	(0.07)
2-year	0.45***	0.49***	0.44***	0.50***
	(0.15)	(0.16)	(0.15)	(0.16)
4-year	1.04***	0.74***	1.04***	0.76***
	(0.29)	(0.17)	(0.27)	(0.18)
6-year	1.46***	0.97***	1.24***	1.01***
	(0.44)	(0.24)	(0.34)	(0.25)
10-year	1.62***	0.66***	0.48***	0.54***
	(0.40)	(0.17)	(0.15)	(0.14)
Control for lagged spending and instruments	No	Yes	No	Yes
Control for lagged dependent variable	No	No	Yes	Yes

Notes: The table presents the IV estimates from regression (2.1). Column (1) shows results from regressions without any controls. Column (2) adds two lags of the change in spending and the instrument. Column (3) adds two lags of the one-period growth in establishments/labor productivity. Column (4) adds both sets of controls. Heteroskedasticity-robust standard errors clustered by MSA are shown in parentheses. ***, **, and * denote significance at the 0.01, 0.05 and 0.1 level, respectively.

B Model appendix

We now turn to presenting the additional details of our general equilibrium model.

B.1 Households' first-order conditions

Impatient households' behavior is described by the following first-order conditions for consumption, housing, labor, and debt, respectively:

$$\lambda_t^b = \left(C_t^b - h^b C_{t-1}^b\right)^{-\sigma_c} - \beta^b h E_t \left\{ \left(C_{t+1}^b - h^b C_t^b\right)^{-\sigma_c} \right\}, \tag{B.1}$$

$$q_t \lambda_t^b = \Upsilon^b \left(H_t^b\right)^{-\sigma_h} + \beta^b E_t \left\{ \lambda_{t+1}^b q_{t+1} \right\} + E_t \left\{ \mu_t^b m(1-\gamma) \frac{q_{t+1}}{R_t} \right\}, \tag{B.2}$$

$$w_t^b \lambda_t^b = \psi \left(N_t^b\right)^{\psi}, \tag{B.3}$$

$$\lambda_t^b + \beta^b \gamma E_t \left\{ \mu_{t+1}^b \right\} = \mu_t^b + \beta^b E_t \left\{ \lambda_{t+1}^b R_t \right\}, \tag{B.4}$$

where λ_t^b and μ_t^b are the multipliers on the budget and borrowing constraints, respectively.

Patient households' first-order conditions with respect to C_t^l, H_t^l, N_t^l, B_t, K_t and I_t are

$$\lambda_t^l = \left(C_t^l - h^l C_{t-1}^l\right)^{-\sigma_c} - \beta^l h^l E_t \left\{ \left(C_{t+1}^l - h^l C_t^l\right)^{-\sigma_c} \right\}, \tag{B.5}$$

$$q_t \lambda_t^l = \Upsilon^l \left(H_t^l\right)^{-\sigma_h} + \beta^l E_t \left\{ \lambda_{t+1}^l q_{t+1} \right\}, \tag{B.6}$$

$$w_t^l \lambda_t^l = \psi \left(N_t^l\right)^{\psi}, \tag{B.7}$$

$$\lambda_t^l = E_t \left\{ \lambda_{t+1}^l \beta^l R_t \right\}, \tag{B.8}$$

$$q_t^k = \beta^l E_t \left\{ \frac{\lambda_{t+1}^l}{\lambda_t^l} \left[r_{t+1}^k + q_{t+1}^k \left((1-\delta) - \phi \left(\frac{I_{t+1}}{K_t} - \delta \right) \left(\frac{1}{2} \left(\frac{I_{t+1}}{K_t} - \delta \right) - \frac{I_{t+1}}{K_t} \right) \right) \right] \right\}, \tag{B.9}$$

$$q_t^k = \left[1 - \phi \left(\frac{I_t}{K_{t-1}} - \delta \right) \right]^{-1}, \tag{B.10}$$

where λ_t^l is the multiplier on the budget constraint and q_t^k is the relative price of capital in terms of consumption.

B.2 Final good firms

The representative final good firm maximizes profits:

$$P_t Y_t - \int_0^1 Q_t(j) p_t(j) dj, \tag{B.11}$$

subject to the production technologies

$$Y_t = \left[\int_0^1 Q_t(j)^{\omega} dj \right]^{\frac{1}{\omega}}, \tag{B.12}$$

$$Q_t(j) = F_t(j)^{\tau + \frac{\rho-1}{\rho}} \left[\sum_{i=1}^{F_t(j)} m_t(j,i)^{\rho} \right]^{\frac{1}{\rho}}. \tag{B.13}$$

The problem is solved in two steps. First, the input of aggregate sectoral goods is found by solving

$$\min_{\{Q_t(j)\}_{j=0}^{1}} \int_0^1 Q_t(j)p_t(j)dj \qquad \text{subject to } Y_t = \left[\int_0^1 Q_t(j)^{\omega}dj\right]^{\frac{1}{\omega}}. \tag{B.14}$$

This leads to the standard demand function and price index:

$$Q_t(j) = \left(\frac{p_t(j)}{P_t}\right)^{\frac{1}{\omega-1}} Y_t, \tag{B.15}$$

$$P_t = \left[\int_0^1 p_t(j)^{\frac{\omega}{\omega-1}}dj\right]^{\frac{\omega-1}{\omega}}. \tag{B.16}$$

Second, the firm decides the mix of inputs within each sector by solving the following:

$$\min_{\{m_t(j,i)\}_{i=1}^{F_t(j)}} \sum_{i=1}^{F_t(j)} p_t(j,i)m_t(j,i) \qquad \text{s.t. } Q_t(j) = F_t(j)^{\tau+\frac{\rho-1}{\rho}} \left[\sum_{i=1}^{F_t(j)} m_t(j,i)^{\rho}\right]^{\frac{1}{\rho}}, \tag{B.17}$$

which has the first-order condition

$$p_t(j,i) - p_t(j)\frac{1}{\rho}F_t(j)^{\tau+\frac{\rho-1}{\rho}} \left[\sum_{i=1}^{F_t(j)} m_t(j,i)^{\rho}\right]^{\frac{1}{\rho}-1} \rho m_t(j,i)^{\rho-1} = 0. \tag{B.18}$$

Rewriting the first-order condition and inserting the expression for $Q_t(j)$ results in the following demand function:

$$m_t(j,i) = \left(\frac{p_t(j,i)}{p_t(j)}\right)^{\frac{1}{\rho-1}} \frac{Q_t(j)}{\left(F_t(j)^{\tau+\frac{\rho-1}{\rho}}\right)^{\frac{\rho}{\rho-1}}} = \left(\frac{p_t(j,i)}{p_t(j)}\right)^{\frac{1}{\rho-1}} \left(\frac{p_t(j)}{P_t}\right)^{\frac{1}{\omega-1}} \frac{Y_t}{\left(F_t(j)^{\tau+\frac{\rho-1}{\rho}}\right)^{\frac{\rho}{\rho-1}}}. \tag{B.19}$$

Lastly, we derive the consumption-based price index for sector j by inserting the demand function into the cost function $Q_t(j)p_t(j) = \sum_{i=1}^{F_t(j)} p_t(j,i)m_t(j,i)$:

$$p_t(j) = \frac{1}{F_t(j)^{\tau+\frac{\rho-1}{\rho}}} \left[\sum_{i=1}^{F_t(j)} p_t(j,i)^{\frac{\rho}{\rho-1}}\right]^{\frac{\rho-1}{\rho}}. \tag{B.20}$$

B.3 Intermediate goods firms

The intermediate goods firm i in sector j maximizes real profits:

$$\frac{p_t(j,i)}{P_t}m_t(j,i) - w_t^l n_t^l(j,i) - w_t^b n_t^b(j,i) - r_t^k k_{t-1}(j,i), \tag{B.21}$$

subject to the production function, the demand for its good, and the sectoral price index:

$$m_t(j,i) = k_{t-1}(j,i)^{\mu}\left[n_t^b(j,i)^{\alpha} n_t^l(j,i)^{1-\alpha}\right]^{1-\mu} - \varphi, \tag{B.22}$$

$$m_t(j,i) = \left(\frac{p_t(j,i)}{p_t(j)}\right)^{\frac{1}{\rho-1}} \left(\frac{p_t(j)}{P_t}\right)^{\frac{1}{\omega-1}} \frac{Y_t}{\left(F_t(j)^{\tau+\frac{\rho-1}{\rho}}\right)^{\frac{\rho}{\rho-1}}}, \tag{B.23}$$

$$p_t(j) = \frac{1}{F_t(j)^{\tau+\frac{\rho-1}{\rho}}}\left[\sum_{i=1}^{F_t(j)} p_t(j,i)^{\frac{\rho}{\rho-1}}\right]^{\frac{\rho-1}{\rho}}. \tag{B.24}$$

The first order conditions with respect to $k_{t-1}(j,i)$, $n_t^b(j,i)$ and $n_t^l(j,i)$ are

$$r_t^k = \mu\frac{p_t(j,i)}{P_t}\frac{k_{t-1}(j,i)^{\mu}\left[n_t^b(j,i)^{\alpha} n_t^l(j,i)^{1-\alpha}\right]^{1-\mu}}{x_t(j,i)k_{t-1}(j,i)}, \tag{B.25}$$

$$w_t^b = (1-\mu)\alpha\frac{p_t(j,i)}{P_t}\frac{k_{t-1}(j,i)^{\mu}\left[n_t^b(j,i)^{\alpha} n_t^l(j,i)^{1-\alpha}\right]^{1-\mu}}{x_t(j,i)n_t^b(j,i)}, \tag{B.26}$$

$$w_t^l = (1-\mu)(1-\alpha)\frac{p_t(j,i)}{P_t}\frac{k_{t-1}(j,i)^{\mu}\left[n_t^b(j,i)^{\alpha} n_t^l(j,i)^{1-\alpha}\right]^{1-\mu}}{x_t(j,i)n_t^l(j,i)}. \tag{B.27}$$

The elasticity of demand according to the demand curve and the sectoral price index is given by

$$\varepsilon_{m_t(j,i)} = \left(\frac{m_t(j,i)}{p_t(j,i)}\frac{1}{\rho-1} + \left(\frac{1}{\omega-1} - \frac{1}{\rho-1}\right)\frac{m_t(j,i)}{p_t(j)}\frac{\rho-1}{\rho}\frac{p_t(j)}{\sum_{i=1}^{F_t} p_t(j,i)^{\frac{\rho}{\rho-1}}}\frac{\rho}{\rho-1}p_t(j,i)^{\frac{\rho}{\rho-1}-1}\right)\frac{p_t(j,i)}{m_t(j,i)}. \tag{B.28}$$

Reducing this and substituting out $\sum_{i=1}^{F_t} p_t(j,i)^{\frac{\rho}{\rho-1}}$ results in the following expression:

$$\varepsilon_{m_t(j,i)} = \frac{1}{\rho-1} + \left(\frac{1}{\omega-1} - \frac{1}{\rho-1}\right)\left(\frac{p_t(j,i)}{p_t(j)F_t(j)^{\tau}}\right)^{\frac{\rho}{\rho-1}}\frac{1}{F_t}. \tag{B.29}$$

Since the firm sells the good in a monopolistic competitive market, it will set its price at a markup over marginal costs. The markup follows from inserting the elasticity into the standard markup rule:

$$x_t(j,i) = \frac{1}{1+\frac{1}{\varepsilon_{m_t(j,i)}}} = \frac{\varepsilon_{m_t(j,i)}}{1+\varepsilon_{m_t(j,i)}}. \tag{B.30}$$

B.4 Steady state

We now describe the non-stochastic steady state of the economy. In the remainder, variables without time subscripts denote steady-state values.

We derive the interest rate and the capital rental rate in steady state from (B.8), (B.9), and (B.10):

$$R = \frac{1}{\beta^l}, \tag{B.31}$$

$$r^k = \frac{1}{\beta^l} - (1-\delta). \tag{B.32}$$

The capital-to-output ratio is derived as

$$\frac{K}{Y} = \frac{\mu}{\frac{1}{\beta^l} - (1-\delta)},$$

while steady-state government spending as a share of output is determined by the parameter θ:

$$\frac{G}{Y} = \theta.$$

Combining the two ratios above with the capital accumulation schedule (3.5) and the aggregate resource constraint (3.26) gives us the consumption-to-output ratio:

$$\frac{C}{Y} = 1 - \bar{G} - \frac{\delta\mu}{\frac{1}{\beta^l} - (1-\delta)}. \tag{B.33}$$

Next, we derive the income shares by using (3.20) and firms' cost-minimization conditions:

$$\frac{r^k K}{Y} = \mu, \tag{B.34}$$

$$\frac{w^b N^b}{Y} = (1-\mu)\alpha, \tag{B.35}$$

$$\frac{w^l N^l}{Y} = (1-\mu)(1-\alpha). \tag{B.36}$$

The steady-state markup follows directly from (3.19):

$$x = \frac{(1-\omega)F - (\rho-\omega)}{\rho(1-\omega)F - (\rho-\omega)}.$$

The steady-state version of the government budget constraint (3.24) implies:

$$\frac{\tau^{TOT}}{Y} = \left(\frac{1}{\beta^l} - 1\right)\frac{B^g}{Y} + \frac{G}{Y}, \tag{B.37}$$

which determines the tax level, since both $\frac{G}{Y} = \theta$ and $\frac{B^g}{Y} = \Xi$ are exogenously determined.

Lastly, we compute the consumption and housing shares of the two households. The housing demand equation (B.2) and the Euler equation (B.4) in combination with (B.31) are given by

$$q\lambda^b = \Upsilon^b \left(H^b\right)^{-\sigma_h} + \beta^b\lambda^b q + \mu^b m\beta^l q(1-\gamma), \tag{B.38}$$

$$\mu^b = \lambda^b \frac{1 - \frac{\beta^b}{\beta^l}}{1 - \beta^b\gamma}. \tag{B.39}$$

Substituting the latter equation into the former yields

$$Y^b \left(H^b\right)^{-\sigma_h} = q\lambda^b \left[1-\beta^b - \frac{\beta^l-\beta^b}{1-\beta^b\gamma} m(1-\gamma)\right]. \tag{B.40}$$

The housing demand equation for the patient households is given by

$$Y^l \left(H^l\right)^{-\sigma_h} = q\lambda^l \left(1-\beta^l\right). \tag{B.41}$$

Dividing (B.41) by (B.40) and inserting the steady state expressions for the budget constraint multipliers—which follow directly from (B.1) and (B.5)—together with the consumption and housing market clearing conditions (3.27) and (3.28) gives us an expression for the housing and consumption shares of the impatient households:

$$\begin{aligned}
\frac{Y^l \left(H^l\right)^{-\sigma_h}}{Y^b \left(H^b\right)^{-\sigma_h}} &= \frac{q\lambda^l\left(1-\beta^l\right)}{q\lambda^b\left[1-\beta^b-\frac{\beta^l-\beta^b}{1-\beta^b\gamma}m(1-\gamma)\right]} \\
\left(\frac{H}{H^b}-1\right)^{-\sigma_h} &= \frac{Y^b}{Y^l}\frac{1-\beta^l}{1-\beta^b-\frac{\beta^l-\beta^b}{1-\beta^b\gamma}m(1-\gamma)}\frac{\lambda^l}{\lambda^b} \\
\left(\frac{H}{H^b}-1\right)^{-\sigma_h} &= \frac{Y^b}{Y^l}\frac{1-\beta^l}{1-\beta^b-\frac{\beta^l-\beta^b}{1-\beta^b\gamma}m(1-\gamma)}\frac{\left(1-h^l\beta^l\right)\left(\left(1-h^l\right)\left(C-C^b\right)\right)^{-\sigma_c}}{\left(1-h^b\beta^b\right)\left(\left(1-h^b\right)C^b\right)^{-\sigma_c}} \\
\left(\frac{H}{H^b}-1\right)^{-\sigma_h} &= \frac{Y^b}{Y^l}\frac{1-\beta^l}{1-\beta^b-\frac{\beta^l-\beta^b}{1-\beta^b\gamma}m(1-\gamma)}\frac{1-\beta^l h^l}{1-\beta^b h^b}\left(\frac{1-h^l}{1-h^b}\right)^{-\sigma_c}\left(\frac{C}{C^b}-1\right)^{-\sigma_c}.
\end{aligned} \tag{B.42}$$

Similarly, we derive an additional expression for the housing and consumption shares of the impatient households by inserting the borrowing constraint (3.3) into their budget constraint (3.2), and using the interest rate (B.31), the labor income share (B.35), and the lump-sum tax payment (3.22):

$$\frac{C^b}{C} = \frac{Y}{C}\left[\left(\beta^l-1\right)m\frac{qH}{Y}\frac{H^b}{H} + \alpha\left(1-\mu-\frac{\tau^{TOT}}{Y}\right)\right]. \tag{B.43}$$

The housing wealth-to-output ratio, $\frac{qH}{Y}$, is calibrated, while the consumption share, $\frac{C}{Y}$, follows from (B.33), so (B.42) and (B.43) are solved numerically for $\frac{H^b}{H}$ and $\frac{C^b}{C}$. The steady-state budget constraint of the patient households has not been used in the derivation of the steady state but will hold by Walras' law.

B.5 Log-linearized model

The model is log-linearized around the non-stochastic steady state. For any generic variable X_t, we let $\hat{X}_t = \ln X_t - \ln X$ denote its log-deviation from steady state. We replace B_t^l and B_t^b by B_t throughout the following.

B.5.1 Optimality conditions of the impatient households

Log-linearization of (B.1), (B.3), and (3.3) gives us the following:

$$\hat{\lambda}_t^b = -\frac{\sigma_c^b}{(1-\beta^b h^b)(1-h^b)}\left(\hat{C}_t^b - h^b \hat{C}_{t-1}^b - \beta^l h^b E_t\left\{\hat{C}_{t+1}^b - h^b \hat{C}_t^b\right\}\right), \tag{B.44}$$

$$\hat{w}_t^b + \lambda_t^b = \psi \hat{N}_t^b, \tag{B.45}$$

$$\widehat{B}_t = \gamma \widehat{B}_{t-1} + (1-\gamma)\left(\mathrm{E}_t \widehat{q}_{t+1} + \widehat{H}_t^b - \widehat{R}_t\right). \tag{B.46}$$

Log-linearization of (B.2) results in

$$q\lambda^b\left(\hat{q}_t + \hat{\lambda}_t^b\right) = -\sigma_h Y^b \left(H^b\right)^{-\sigma_h} \hat{H}_t^b + \beta^b \lambda^b q E_t\left\{\hat{\lambda}_{t+1}^b + \hat{q}_{t+1}\right\} + \mu^b m (1-\gamma)\frac{q}{R} E_t\left\{\hat{\mu}_t^b + \hat{q}_{t+1} - \hat{R}_t\right\},$$

which is rewritten using (B.31), (B.40), and (B.39):

$$\begin{aligned} q\lambda^b\left(\hat{q}_t + \hat{\lambda}_t^b\right) &= -\sigma_h q \lambda^b\left[1-\beta^b - \frac{\beta^l-\beta^b}{1-\beta^b\gamma} m(1-\gamma)\right]\hat{H}_t^b + \beta^b\lambda^b q E_t\left\{\hat{\lambda}_{t+1}^b + \hat{q}_{t+1}\right\} \\ &\quad + \lambda^b \frac{1-\frac{\beta^b}{\beta^l}}{1-\beta^b\gamma} m(1-\gamma)\frac{q}{R}E_t\left\{\hat{\mu}_t^b + \hat{q}_{t+1} - \hat{R}_t\right\}, \end{aligned}$$

$$\begin{aligned} \hat{q}_t + \hat{\lambda}_t^b &= -\sigma_h\left[1-\beta^b - \frac{\beta^l-\beta^b}{1-\beta^b\gamma} m(1-\gamma)\right]\hat{H}_t^b + \beta^b E_t\left\{\hat{\lambda}_{t+1}^b + \hat{q}_{t+1}\right\} \\ &\quad + \frac{\beta^l-\beta^b}{1-\beta^b\gamma} m(1-\gamma) E_t\left\{\hat{\mu}_t^b + \hat{q}_{t+1} - \hat{R}_t\right\}. \end{aligned} \tag{B.47}$$

In addition, (B.4) becomes

$$\lambda^b \widehat{\lambda}_t^b + \beta^b \gamma \mu^b E_t\left\{\widehat{\mu}_{t+1}^b\right\} = \mu^b \widehat{\mu}_t^b + \beta^b \lambda^b R E_t\left\{\widehat{\lambda}_{t+1}^b + \widehat{R}_t\right\}.$$

Rewriting this using (B.39) results in

$$\widehat{\lambda}_t^b + \beta^b\gamma \frac{1-\frac{\beta^b}{\beta^l}}{1-\beta^b\gamma} E_t\left\{\widehat{\mu}_{t+1}^b\right\} = \frac{1-\frac{\beta^b}{\beta^l}}{1-\beta^b\gamma}\widehat{\mu}_t^b + \frac{\beta^b}{\beta^l}E_t\left\{\widehat{\lambda}_{t+1}^b + \widehat{R}_t\right\}. \tag{B.48}$$

The log-linearized budget constraint becomes

$$\begin{aligned} &\frac{C^b}{C}\frac{C}{Y}\hat{C}_t^b + \frac{qH}{Y}\frac{H^b}{H}\left(\hat{H}_t^b - \hat{H}_{t-1}^b\right) + m\frac{qH}{Y}\frac{H^b}{H}\left(\hat{R}_{t-1} + \hat{B}_{t-1}\right) = \\ &(1-\mu)\alpha\left(\hat{w}_t^b + \hat{N}_t^b\right) + m\frac{qH}{Y}\frac{H^b}{H}\beta^l \hat{B}_t - \alpha\frac{\tau^{TOT}}{Y}\widehat{\tau}_t^{TOT}. \end{aligned} \tag{B.49}$$

B.5.2 Optimality conditions of the patient households

Log-linearization of (B.5), (B.7), (B.8), and (3.5) results in

$$\hat{\lambda}_t^l = -\frac{\sigma_c^l}{(1-\beta^l h^l)(1-h^l)} \left(\hat{C}_t^l - h^l \hat{C}_{t-1}^l - \beta^l h^l E_t \left\{ \hat{C}_{t+1}^l - h^l \hat{C}_t^l \right\} \right), \tag{B.50}$$

$$\hat{\lambda}_t^l = E_t \left\{ \hat{\lambda}_{t+1}^l \right\} + \hat{R}_t, \tag{B.51}$$

$$\hat{w}_t^l + \hat{\lambda}_t^l = \psi \hat{N}_t^l, \tag{B.52}$$

$$\hat{K}_t = (1-\delta)\hat{K}_{t-1} + \delta \hat{I}_t. \tag{B.53}$$

Similar to the derivation for the impatient households, log-linearization of (B.6) in combination with (B.41) becomes

$$\hat{q}_t + \hat{\lambda}_t^l = -\sigma_h \left(1-\beta^l\right) \hat{H}_t^l + \beta^l E_t \left\{ \hat{q}_{t+1} + \hat{\lambda}_{t+1}^l \right\}. \tag{B.54}$$

The log-linearized first-order conditions for capital and investment, (B.9) and (B.10), are

$$\hat{q}_t^k = E_t \left\{ \hat{\lambda}_{t+1}^l \right\} - \hat{\lambda}_t^l + \beta^l r^k \hat{r}_{t+1}^k + \beta^l (1-\delta) E_t \left\{ q_{t+1}^k \right\} + \beta^l \delta^2 \phi E_t \left\{ \hat{I}_{t+1} - \hat{K}_t \right\},$$

$$\hat{q}_t^k = \phi \delta \left(\hat{I}_t - \hat{K}_{t-1} \right).$$

Combining these two equations to eliminate $\hat{q}_t^k$ and inserting (B.53) results in

$$\phi \left(\hat{K}_t - \hat{K}_{t-1} \right) + \hat{\lambda}_t^l = E_t \left\{ \hat{\lambda}_{t+1}^l + \beta^l r^k \hat{r}_{t+1}^k + \beta^l \phi \left(\hat{K}_{t+1} - \hat{K}_t \right) \right\}. \tag{B.55}$$

Lastly, the budget constraint (3.4) becomes

$$\begin{aligned} &\frac{C^l}{C}\frac{C}{Y}\hat{C}_t^l + \frac{qH}{Y}\frac{H^l}{H}\left(\hat{H}_t^l - \hat{H}_{t-1}^l\right) + \frac{\delta K}{Y}\hat{I}_t + m\frac{qH}{Y}\frac{H^b}{H}\beta^l \hat{B}_t + \Xi \hat{B}_t^g = \\ &(1-\mu)(1-\alpha)\left(\hat{w}_t^l + \hat{N}_t^l\right) + m\frac{qH}{Y}\frac{H^b}{H}\left(\hat{R}_{t-1} + \hat{B}_{t-1}\right) + \mu\left(\hat{r}_t^k + \hat{K}_{t-1}\right) \\ &+ \frac{1}{\beta^l}\Xi\left(\hat{R}_{t-1} + \hat{B}_{t-1}^g\right) - (1-\alpha)\frac{\tau_t^{TOT}}{Y}\hat{\tau}_t^{TOT}. \end{aligned} \tag{B.56}$$

B.5.3 Symmetric firm equilibrium conditions

The log-linearized factor prices read as

$$\hat{r}_t^k = (1+\tau)\left(\left(\left(\mu - \frac{1}{1+\tau}\right)\hat{K}_{t-1} + (1-\mu)\left(\alpha \hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t, \tag{B.57}$$

$$\hat{w}_t^b = (1+\tau)\left(\left(\mu \hat{K}_{t-1} + (1-\mu)\left(\left(\alpha - \frac{1}{(1+\tau)(1-\mu)}\right)\hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t, \tag{B.58}$$

$$\hat{w}_t^l = (1+\tau)\left(\left(\mu \hat{K}_{t-1} + (1-\mu)\left(\alpha \hat{N}_t^b + \left(1-\alpha - \frac{1}{(1+\tau)(1-\mu)}\right)\hat{N}_t^l\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t. \tag{B.59}$$

while log-linearization of (3.20) results in

$$\hat{Y}_t = (1+\tau)\left(\left(\mu\hat{K}_{t-1} + (1-\mu)\left(\alpha\hat{N}^b_t + (1-\alpha)\hat{N}^l_t\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t. \tag{B.60}$$

We can combine the production function (3.11) with (3.16) to obtain:

$$F_t = \frac{x_t - 1}{x_t\varphi}K^{\mu}_{t-1}\left[\left(N^b_t\right)^{\alpha}\left(N^l_t\right)^{1-\alpha}\right]^{1-\mu}.$$

Combining this with (3.20), we obtain the number of firms as a function of output and the markup:

$$F_t = \left(\frac{x_t - 1}{\varphi}\right)^{\frac{1}{1+\tau}} Y_t^{\frac{1}{1+\tau}},$$

which is log-linearized as

$$\hat{F}_t = \frac{1}{1+\tau}\left(\hat{Y}_t + \frac{x}{x-1}\hat{x}_t\right). \tag{B.61}$$

We rewrite the markup (3.19) as

$$F_t\left(\rho x_t - 1\right)\left(1-\omega\right) = \left(x_t - 1\right)\left(\rho - \omega\right),$$

which is log-linearized as

$$F\left(\rho x - 1\right)\left(1-\omega\right)\hat{F}_t + \rho x F\left(1-\omega\right)\hat{x}_t = x\left(\rho-\omega\right)\hat{x}_t.$$

Inserting $F = \frac{(x-1)(\rho-\omega)}{(\rho x-1)(1-\omega)}$ into the equation above and rearranging yields

$$\hat{F}_t = \frac{x}{x-1}\frac{\rho-1}{\rho x-1}\hat{x}_t. \tag{B.62}$$

The log-linearized expression for TFP is

$$T\hat{F}P_t = \tau\hat{F}_t - \hat{x}_t. \tag{B.63}$$

B.5.4 Fiscal policy, market clearing conditions, and shock processes

The log-linearized version of the government's budget constraint (3.24) is

$$\frac{1}{\beta^l}\Xi Y\left(\hat{R}_{t-1} + \hat{B}^g_{t-1}\right) + \theta\hat{G}_t = \frac{\tau^{TOT}}{Y}\hat{\tau}^{TOT}_t + \Xi\hat{B}^g_t, \tag{B.64}$$

while the adjustment rule for the tax level is given by

$$\hat{\tau}^{TOT}_t = \rho_\tau\hat{\tau}^{TOT}_{t-1} + \left(1-\rho_\tau\right)\gamma_\tau\left(\hat{B}^g_{t-1} - \hat{Y}_{t-1}\right). \tag{B.65}$$

The market clearing conditions (3.26), (3.27), and (3.28) become

$$\hat{Y}_t = \frac{C}{Y}\hat{C}_t + \theta\hat{G}_t + \frac{I}{Y}\hat{I}_t,$$

$$\hat{C}_t = \frac{C^b}{C}\hat{C}^b_t + \frac{C^l}{C}\hat{C}^l_t, \tag{B.66}$$

$$0 = \frac{H^b}{H}\hat{H}^b_t + \frac{H^l}{H}\hat{H}^l_t. \tag{B.67}$$

The good market clearing condition is not included in the Dynare code since it is redundant by Walras' law.

The log-linearized shock process for government spending (3.21) is

$$\hat{G}_t = \gamma_g \hat{G}_t + \epsilon_{g,t}. \tag{B.68}$$

The 25 equations (B.44) to (B.68) define the log-linearized model for the 25 endogeneous variables $\hat{\lambda}^b_t$, $\hat{\lambda}^l_t$, $\hat{C}^b_t$, $\hat{C}^l_t$, $\hat{w}^b_t$, $\hat{w}^l_t$, $\hat{N}^b_t$, $\hat{N}^l_t$, $\hat{H}^b_t$, $\hat{H}^l_t$, $\hat{B}_t$, $\hat{q}_t$, $\hat{R}_t$, $\hat{\mu}^b_t$, $\hat{G}_t$, $\hat{\tau}^{TOT}_t$, $\hat{B}^g_t$, $\hat{K}_t$, $\hat{I}_t$, $\hat{r}^k_t$, $\hat{x}_t$, $\hat{F}_t$, $\hat{Y}_t$, $\hat{C}_t$, and $T\hat{F}P_t$.

C Stylized model

We assume the economy to be solely populated by financially unconstrained households that exhibit logarithmic nondurable consumption utility, and intermediate goods firms featuring a production technology that is linear in labor, the only production input. Under these assumptions, we can retrieve the following set of log-linearized equations, corresponding to (B.66), (B.52), (B.59), (B.60), (B.61), and (B.62), respectively:

$$\hat{y}_t = (1-\theta)\hat{c}_t + \theta\hat{g}_t, \tag{C.1}$$

$$\hat{w}_t = \psi\hat{n}_t + \hat{c}_t, \tag{C.2}$$

$$\hat{w}_t = \tau\hat{n}_t - \frac{x-(1+\tau)}{x-1}\hat{x}_t, \tag{C.3}$$

$$\hat{y}_t = (1+\tau)\hat{n}_t - \frac{x-(1+\tau)}{x-1}\hat{x}_t, \tag{C.4}$$

$$\hat{F}_t = \frac{1}{1+\tau}\left(\hat{y}_t + \frac{x}{x-1}\hat{x}_t\right), \tag{C.5}$$

$$\hat{F}_t = \frac{x}{x-1}\frac{\rho-1}{\rho x-1}\hat{x}_t. \tag{C.6}$$

First, combine (C.1) and (C.2) to eliminate $\hat{c}_t$. Thus, plug (C.3) in the resulting equation and rearrange to obtain

$$\frac{(\tau-\psi)(1-\gamma)-(1+\tau)}{(1+\tau)(1-\gamma)}\hat{y}_t + \frac{[\tau-\psi-(1+\tau)][x-(1+\tau)]}{(1+\tau)(x-1)}\hat{x}_t = -\frac{\gamma}{1-\gamma}\hat{g}_t. \tag{C.7}$$

Separately, combine (C.5) and (C.6) to eliminate $\hat{F}_t$ and obtain

$$\hat{x}_t = \frac{(x-1)(\rho x - 1)}{x\left[(\rho-1)(1+\tau) - (\rho x - 1)\right]}\hat{y}_t. \tag{C.8}$$

Finally, combine (C.7) and (C.8) to obtain the response of $\hat{y}_t$ to $\hat{g}_t$:

$$\hat{y}_t = \frac{\gamma x (1+\tau)\left[(\rho x-1) - (\rho-1)(1+\tau)\right]}{x\left[(\tau-\psi)(1-\gamma) - (1+\tau)\right]\left[(\rho-1)(1+\tau) - (\rho x-1)\right] + (\rho x-1)(1-\gamma)\left[\tau-\psi-(1+\tau)\right]\left[x-(1+\tau)\right]}\hat{g}_t.$$

Combining this solution with (C.1), we are able to infer when $\hat{g}_t$ implies movements in $\hat{c}_t$ (and, thus, in house prices, given that $\hat{c}_t \approx \hat{q}_t$) of the same sign, which is the case whenever $\frac{\hat{y}_t}{\hat{g}_t} > \gamma$.

D Additional numerical results

This appendix contains additional results based on the quantitative model presented in Section 3. Table D.1 presents the estimated parameter values from the alternative model version without taste for variety, i.e. where we impose $\tau = 0$ from the outset.

Table D.1: Estimated parameters in the model with $\tau = 0$

Parameter	Description	Value
σ_c	Curvature in utility of consumption	4.997 [4.776−5.000]
σ_h	Curvature in utility of housing	0.926 [0.314−1.638]
h^l	Habit formation, lenders	0.808 [0.532−0.900]
h^b	Habit formation, borrowers	0.768 [0.306−0.900]
ψ	Inverse Frisch elasticity	0.250 [0.250−0.267]
ϕ	Capital adjustment cost parameter	24.583 [10.256−24.996]
γ	Inertia of mortgage debt	0.950 [0.933−0.950]
x	Steady-state value of markup	1.330 [1.322−1.330]
γ_τ	Tax response to government debt	0.895 [0.407−0.900]
ρ_τ	Inertia of tax level	0.294 [0.101−0.833]
γ_G	Persistence of government spending shock	0.967 [0.924−0.980]
σ_g	Std. dev. of government spending shock	0.099 [0.082−0.113]

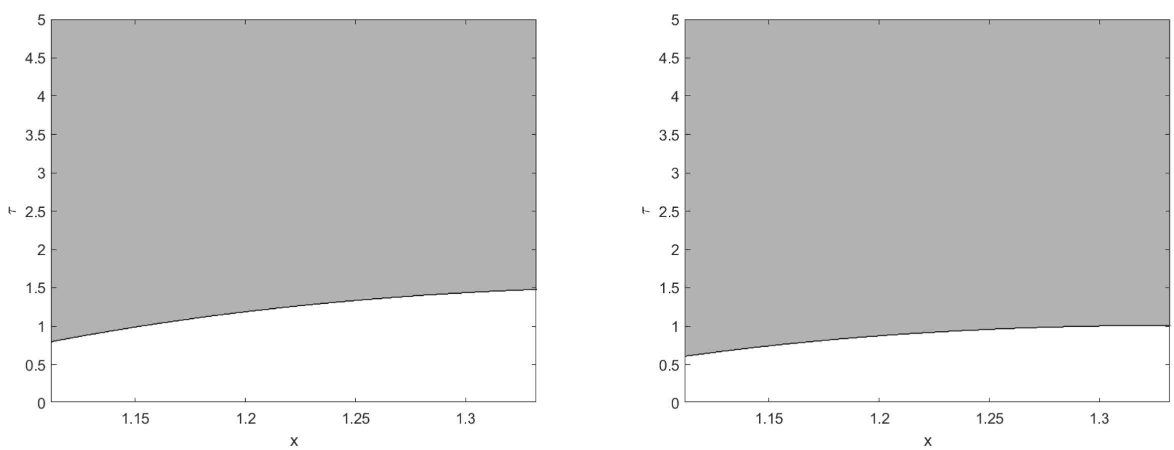
Note: 68 percent credible sets for the estimated parameters are reported in brackets.

As the table indicates, a number of parameters (or their credible sets) reach the bounds we have imposed in the estimation. This is the case for σ_c, which is bounded above at 5; h^l and h^b, which are bounded above at 0.9; ψ, which is bounded below at 0.25; ϕ, which is bounded above at 25; γ, which is bounded above at 0.95; x, which is bounded above at $\frac{1}{\omega} = 1.33$; γ_τ, which is bounded above at 0.9; and ρ_τ, which is bounded below at 0.1. These bounds are imposed in order to avoid areas of the parameter space where the model was found to be unstable, and/or to prevent parameter estimates that are considered unrealistic by the existing literature. The same set of bounds were also imposed in the estimation of the baseline model, but were found to be less relevant.

We then return to the baseline model characterized by the parameter estimates reported in Section 5.2.1. We rely on Figure D.5 to shed additional light on the interplay between the taste for variety and the steady-state markup in the model. In the left panel, we report the combinations of the taste for variety parameter, τ, and the steady-state markup, x, for which the model generates an increase in the house price on impact, holding all other parameters fixed at the values reported in Table 1. In the right panel, we focus on the cumulative response of the house price.

Figure D.5: House price response for different parameter combinations

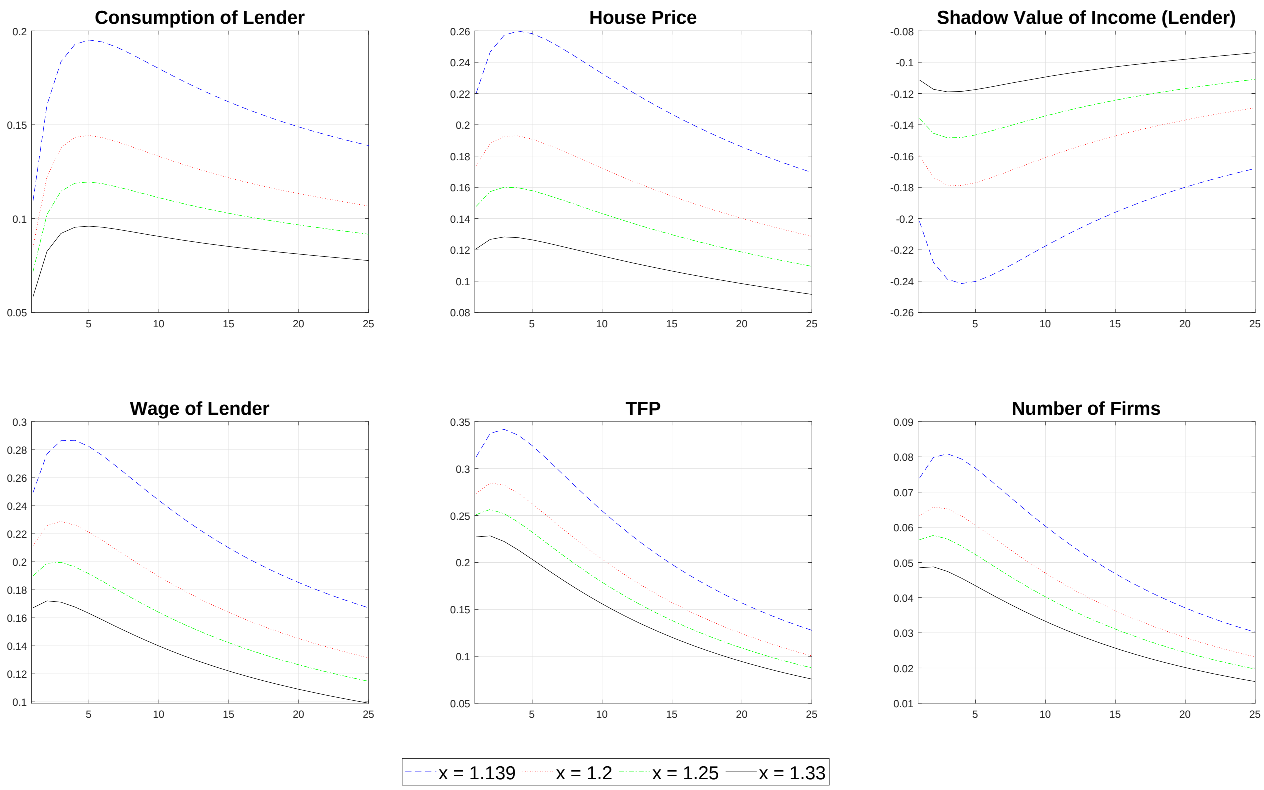
(a) Impact response (b) Cumulative response



Notes: The figure shows the model outcomes for different combinations of the parameter values of τ and x. The grey (white) area indicates parameter combinations where the model produces an increase (a decline) in the house price in response to a government spending shock. The left panel considers the impact response of the house price, while the right panel considers the cumulated response over 25 periods.

We finally report the effects of an increase in government spending for various values of the steady-state markup x, holding all other parameters fixed at their baseline values.

Figure D.6: Effects of a government spending shock for different values of x



Notes: The figure shows the effects of a shock to government spending for various values of the steady-state markup x. Dashed blue line: $x = 1.139$ (estimated value). Dotted red line: $x = 1.2$. Dashed-dotted green line: $x = 1.25$. Solid black line: $x = 1.33$. All other parameters are kept at their baseline values.