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

Rasmus Bisgaard Larsen† Søren Hove Ravn‡ Emiliano Santoro§

September 2021

Abstract

We report new regional evidence indicating that U.S. house prices increase persistently in the face of positive shocks to fiscal spending. In sharp contrast with this fact, though, house prices fall in conventional dynamic general equilibrium models where Ricardian households benefit from the flow of housing services. The inconsistency rests on the negative wealth effect exerted by the concurrent increase in the present-value tax burden, which increases the marginal utility of consumption, even in the presence of mechanisms that produce consumption crowding-in. Due to the quasi-constant shadow value of housing, this property inevitably depresses house prices. To address this problem, we devise a model with two layers of production: A final-good, fully competitive 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 aggregate production. This helps overcoming the negative wealth effect, thus flipping the response of Ricardian households' marginal utility of consumption and, thus, that of house prices. We match the impulse responses from the model to those obtained from a Bayesian Vector Autoregression framework featuring house prices and other macroeconomic variables, and show that the model can account for the aggregate effects of federal fiscal spending.

*The authors wish to thank Søren Leth-Petersen, Vivien Lewis, Luigi Paciello, Ivan Petrella, and Celine Poilly for their comments. The usual disclaimer applies.

†Danmarks Nationalbank. *E-mail*: rbl@nationalbanken.dk.

‡Department of Economics, University of Copenhagen. *E-mail*: Soren.Hove.Ravn@econ.ku.dk.

§Department of Economics, University of Copenhagen. *E-mail*: Emiliano.Santoro@econ.ku.dk.

1 Introduction

The Great Recession has shed light on the key role that housing plays 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, have 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 academia as well as in policy circles. Yet, surprisingly few studies have investigated the direct link between changes in government spending and house prices, empirically as well as theoretically. Upon documenting that house prices display a positive response to an increase in government spending, the main focus of this paper is to produce a theoretical model capable of framing the transmission of shocks to fiscal spending on house prices.

Using contract data from the U.S. Department of Defense (DoD), we report that a positive change in federal government spending in a given city expands the price of housing relative to other cities. In support of our subsequent modeling strategy we also show, through the use of U.S. Census data on County Business Patterns, that this fact is accompanied by a marked increase in the number of establishments in the area being perturbated, and more so once we account for the sectoral bias in DoD spending. These findings add to existing structural Vector Autoregression (VAR) studies showing how an expansion in fiscal spending produces a rise in house prices, along with various other macroeconomomic aggregates, such as output, consumption, and net business formation (e.g., Khan and Reza, 2017; Lewis and Winkler, 2017; Auerbach et al., 2019).

In sharp contrast with these findings, house prices are found to fall in a large variety of dynamic stochastic general equilibrium (DSGE) models. In fact, it is possible to show that 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, will feature this type of property. 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, from the perspective of a Ricardian household, housing features an approximately constant shadow value. Two key elements lead to 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—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

[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).

value of housing remains approximately constant, the relative price of housing must track the behavior of Ricardian households' nondurable consumption. As discussed by 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 final-good, fully competitive 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 at the aggregate level. This 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 the labor market equilibrium. An expansion in government spending typically leads to an increase in labor supply, at given factor prices. In a standard economy with no entry, holding the number of intermediate goods producers fixed would consequently lead to a fall in the real wage, thus exacerbating the fall in the present value of disposable income.

With free entry, instead, enhanced profit opportunities determine an increase in the number of intermediate producers. This leads to a decline in the markup of intermediate goods prices over their marginal cost; the so-called *competition effect* (see, e.g., Lewis and Winkler, 2017). Markup countercyclicality entails an outward shift in the demand curves for goods and labor, thus stimulating the response of consumption and the real wage (see, e.g., Hall et al., 2009). The increase in the number of firms also raises total factor productivity (TFP). Thus, 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. We complement this channel with taste for variety à la Benassy (1996), which implies that, as the number of intermediate goods producers within a sector increases, the aggregate sectoral good expands for a given input of intermediate goods; the *variety effect*. As a result of this combination, following a fiscal stimulus our model entails a robust increase in TFP, which drives further up the marginal product of labor and, thus, the wage rate. 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.[2]

[2]The empirical plausibility of our proposed mechanism is supported by Epstein et al. (2020), 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.

We first consider a simplified version of the model which may be solved analytically. Within this context, we show that the competition effect in isolation may be sufficient to generate a positive response of private consumption and the house price to a government spending shock, but only if the steady-state markup or the Frisch elasticity of labor supply is very high. Once we introduce the variety effect, a joint increase in consumption and the house price obtains—for realistic values of the markup and the Frisch elasticity—when the taste for variety is sufficiently high.

We then proceed to analyze the full model from a quantitative viewpoint, matching its impulse responses to the empirical ones from a Bayesian Vector Autoregression (BVAR) featuring house prices, output, consumption, TFP, mortgage debt, a measure of the real wage, along with federal government spending. In line with our model, all these variables increase following a positive shock to fiscal spending. In fact, matching the real wage response helps us obtain a large increase in TFP, which is key to generate a crowding-in effect on private consumption and, thus, an increase in house prices. Moreover, 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. 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.

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 these 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.[3] The specification in Lewis and Winkler (2017), instead, embeds the competition effect as in our setup, but with 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), who estimate a structural VAR model for the US, reporting a positive effect. 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

[3] 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).

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 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.[4] Second, the specification used in Lewis and Winkler (2017) assumes that the flow of government spending is productivity-enhancing (or, equivalently, that the public capital stock depreciates entirely each period). If one assumes 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 have found 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.[5]

We choose instead to draw on an emerging literature combining endogenous firm entry with love 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 for variety introduced by Benassy (1996). This function disentangles market power from love for 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 love for variety (Chugh and Ghironi, 2011; Bilbiie et al., 2019). Lewis and Poilly (2012) estimate a DSGE model featuring a CES function with generalized love of variety using impulse-response matching to monetary policy shocks, and report that the love of variety parameter is poorly identified. Similar to their findings, our matching exercise returns a rather imprecise estimate, though our model produces a positive response of the house price to a government spending shock for a wide range of values of the parameter. Nonetheless, our findings echo the calls of

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

[5] These results are available upon request.

Bilbiie et al. (2012) and Bilbiie (forthcoming) 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.[6] 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 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. In Section 5, we describe the calibration and estimation of the model, and then discuss its qualitative and quantitative implications. Section 6 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. Furthermore, as our proposed explanation builds on the entry and exit of firms, we also examine the response of the number of establishments to fiscal spending shocks. More specifically, we study how a change in federal government spending in a given city relative to others affects relative house price movements and net firm entry. 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.

2.1 Data and methodology

Our analysis 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

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

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.1.1. To measure local house prices, we use the Freddie Mac House Price Index, while we normalize DoD spending by local activity using GDP from the Bureau of Economic Analysis (BEA). Net firm entry is measured as the growth in the number of establishments within the MSA. Establishment data are taken from the County Business Patterns from the U.S. Census, which contains information on the stock of establishments at the county level. We aggregate these data to get MSA-level series on establishment counts. The final panel data set covers 380 MSAs from 2001 through 2019, at the annual frequency.

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

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

where $Z_{i,t}$ is either the house price index or the number of establishments, $G_{i,t}$ is DoD spending, $X_{i,t}$ is a vector of controls, and $Y_{i,t}$ is GDP. The MSA fixed effect, $\alpha_{i,h}$, controls for MSA-specific trends in house prices and firm entry, while the time fixed effect, η_{t+h}, controls for common, national variation in house prices and firm entry.[7] All variables are measured in nominal terms, though we obtain similar results when using the MSA-level GDP deflator.[8]

The coefficient of interest is β_h, which measures the growth in house prices or establishments from t to $t+h$ relative to other MSAs, as a result of an increase in DoD spending by 1 % of initial GDP from period t to $t+1$.[9] 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).

[7] The MSA-level normalized change in DoD spending is winsorized at the 1 % 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.1.3.

[8] 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 instead differences in industry composition. In a recent paper, Hazell et al. (2020) circumvent this imputation issue by constructing regional price indices using BLS micro-data. However, these indices are only constructed for 34 states.

[9] 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 structural VAR that will serve as a basis for the calibration of the model.

We deal with the potential bias by instrumenting the change in local DoD spending with a Bartik (1991) instrument: The change in national DoD spending interacted with the MSA's average share of national DoD spending over the sample period. This instrument identifies the effect of spending on house prices and firm entry 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.1.3 plots the period-by-period first-stage Kleibergen-Papp F-statistics from the regression with house price growth as the dependent variable.[10] The F-statistics are in the range 45-134, which is well above the cluster-robust threshold for weak instruments of 23.1 provided by Montiel Olea and Pflueger (2013), so the instrument is quite strong.

The identifying assumption behind this approach is that, conditional on controls, there are no confounding factors—not only contemporaneously, but also at leads and lags—affecting local house price and firm entry growth 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:[11]

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

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

The average DoD spending share in a 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 or establishments is codetermined with the local share of national DoD spending, equation (2.2) will 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 and firm entry 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 $\frac{G^{nat}_{t+1} - G^{nat}_t}{Y_{i,t}}$

[10] The F-statistics from the regression of establishment growth are almost identical.

[11] 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).

in the time-series dimension. We address some of these concerns in Appendix A.1.3, 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 if the instrument itself is serially correlated. 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 include 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, $\frac{G_{i,t+1} - G_{i,t}}{Y_{i,t}}$. We also add two lags of the one-year growth in the dependent variable, $\frac{Z_{i,t+1} - Z_t}{Z_{i,t}}$.

2.2 Results

We estimate h separate (2.1) regressions, for $h = 1, 2, \ldots, 12$ horizons, and present the estimates in Figure 1. The OLS and IV estimates for the house price response are shown in panels (a) and (b), respectively, while the corresponding estimates for the number of establishments are shown in panels (c) and (d). Standard errors are heteroskedasticity-robust and clustered at the MSA level, so as to account for within-MSA correlation of the error term. 95 percent confidence bands based on the point estimate standard errors are indicated by the grey areas.

According to the IV estimates, the response of house prices to an expansion in government spending follows a hump-shaped pattern, peaking at the sixth year, thus reverting back to the trend. In terms of magnitude, we appreciate a relative increase in house prices of 1.6 % over six years, as a result of an increase in spending of 1 % of GDP in the first year. A similar hump-shaped response is observed for the number of establishments, albeit the estimates are smaller in magnitude and not as statistically significant, with an increase of 0.2 % after six years, prior to the reversal. For both the house price and establishment responses, the OLS estimates follow a hump-shaped pattern, but are biased toward zero.

These estimates are robust to a number of alternative specifications of (2.1), as reported in Appendix A.1.3. 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 prices and establishments movements associated with potential confounding factors, such as local industry composition and exposure to regional business cycles. In addition, we examine the robustness of our results to outliers and the set of controls included in the baseline regression.

2.3 Establishment responses at the sectoral level

The response of the number of establishments in Figure 1 is positive, yet barely statistically significant. We now show that this is because firms enter a given sector to a greater extent, as compared with other sectors, when the local government increases its purchases of goods and services from that specific sector. Such a differential response cannot be estimated by regression (2.1), since this is designed to capture the average effect on the number of establishments in all sectors, when government spending changes irrespective of which sector spending is aimed at.

In order to understand if the response of establishments in a given sector is larger when changes in government spending are targeted to that sector, we disaggregate our contract data. We split these by sectors, as defined by two-digit NAICS codes, and aggregate the contracts such that we measure total annual military spending in sector c located in MSA i, $G_{i,c,t}$. For each sector c, we also construct a variable for sectoral military spending, $G'_{i,c,t}$, as the sum of contracts going to all other sectors than c, $G'_{i,c,t} = \sum_{k \neq c} G_{i,k,t}$. By regressing the growth in the number of sector-c establishments on these two government spending variables normalized by local aggregate GDP, we can distinguish the effect of own-sector government spending from that of government spending outside that sector:

$$\frac{Z_{i,c,t+h} - Z_{i,c,t}}{Z_{i,c,t}} = \alpha_{i,c,h} + \eta_{c,t+h} + \beta_{same,h} \frac{G_{i,c,t+1} - G_{i,c,t}}{Y_{i,t}} + \beta_{other,h} \frac{G'_{i,c,t+1} - G'_{i,c,t}}{Y_{i,t}} + \gamma_{c,h} X_{i,c,t} + \varepsilon_{i,c,t+h}. \tag{2.3}$$

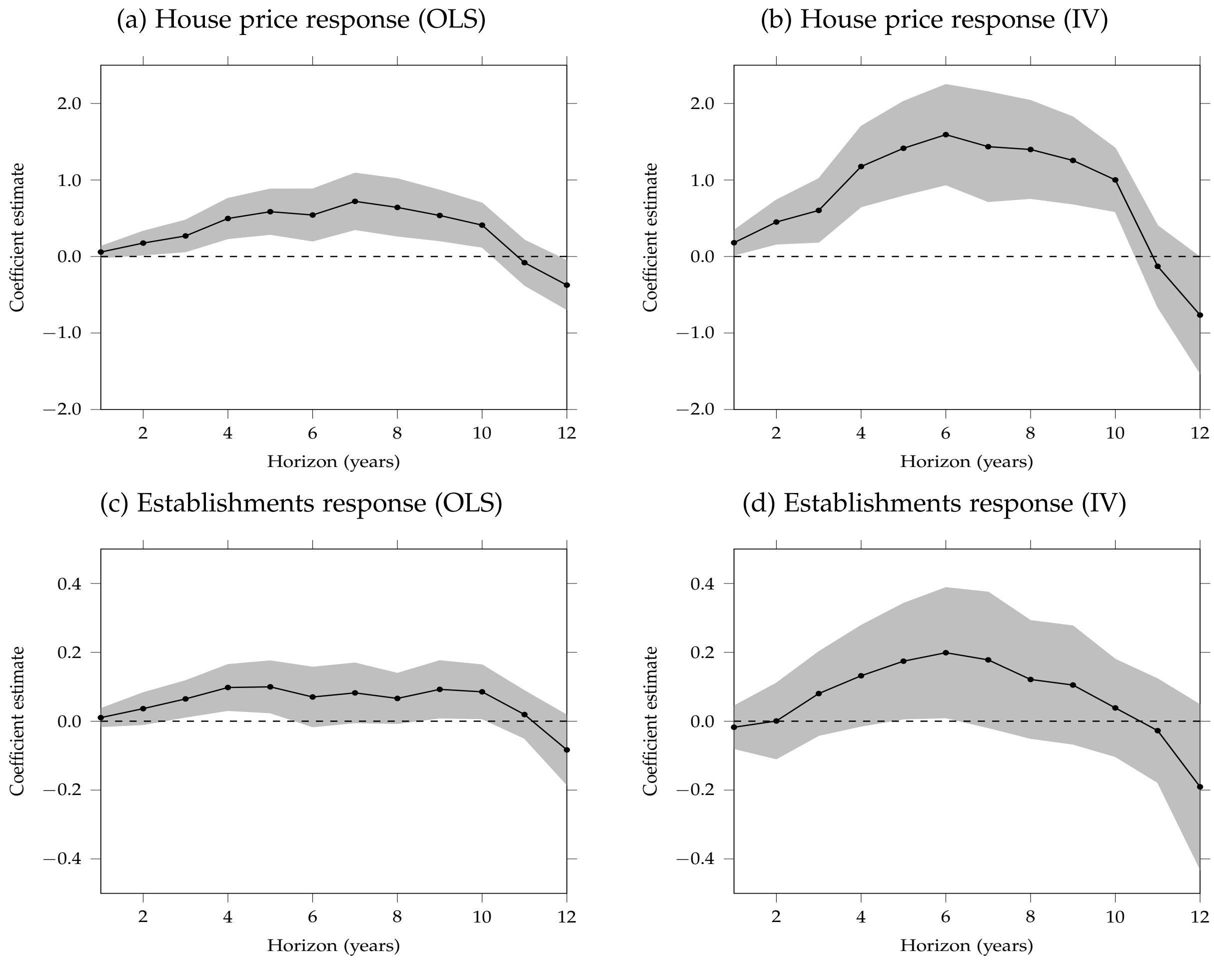
Two instruments are needed in order to identify $\beta_{same,h}$ and $\beta_{other,h}$, so we extend our existing IV strategy. First, we instrument for the change in own-sector government spending, $\frac{G_{i,c,t+1} - G_{i,c,t}}{Y_{i,t}}$, with the change in national own-sector spending interacted with MSA i's average share of national own-sector spending, $\bar{G}_i^{nat} \cdot \frac{G^{nat}_{i,c,t+1} - G^{nat}_{i,c,t}}{Y_{i,t}}$. Second, the change in spending outside sector c, $\frac{G'_{i,c,t+1} - G'_{i,c,t}}{Y_{i,t}}$, is instrumented with the change in national spending outside sector c interacted with MSA i's average share of national spending outside that sector, $\bar{G}_i'^{nat} \cdot \frac{G'^{nat}_{i,c,t+1} - G'^{nat}_{i,c,t}}{Y_{i,t}}$.

The set of controls included in regression (2.3) mirrors that included in regression (2.1). We include MSA $\times$ sector fixed effects, $\alpha_{i,c,h}$, as well time $\times$ sector fixed effects, $\eta_{c,t+h}$, in order to control for sector-specific MSA and time trends. In addition, we control for two lags of the one-year growth in the dependent variable, $\frac{Z_{i,c,t+1} - Z_{i,c,t}}{Z_{i,c,t}}$, as well as two lags of the two endogenous variables and their instruments.

The IV estimates from regression (2.3) are presented in Figure 2, along with their 95 percent confidence bands based on standard errors clustered by MSAs. The estimates for the number of sector-level establishments to a change in own-sector government spending, $\beta_{same,h}$, are plotted in the left panel, and show that the number of establishments in a sector increases significantly when government spending aimed at that sector increases. The peak response of 1.3 % is about six times larger than that of aggregate establishments to an aggregate government spending change.

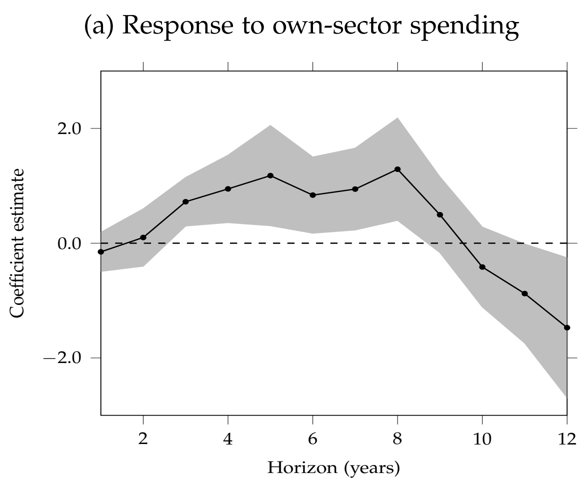
Conversely, the estimates for other-sectors government spending, $\beta_{other,h}$, show that the number of sector-level establishments barely responds when government spending changes outside the sector (right panel). This muted response is estimated relatively precisely, while resulting statistically insignificant, except for the first year after the change in government spending.

Figure 1: Regional 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), while the OLS and IV estimates for the response of establishments are shown in panels (c) and (d). The regressions include as controls two lags of the one-year growth in house prices and establishments, 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.

Figure 2: Regional responses of establishments to own-sector and other-sectors military spending



Notes: The figure shows the IV estimates of $\beta_{same,h}$ (left panel) and $\beta_{other,h}$ (right panel) from regression (2.3), based on an annual panel of 380 MSAs covering the period 2001-2019. The set of controls includes two lags of the one-year growth in establishments, two lags of the two instruments, two lags of the one-year change in local own-sector spending normalized by GDP, and two lags of the one-year change in local other-sectors spending normalized by GDP. Grey areas indicate 95 percent confidence bands constructed using heteroskedasticity-robust standard errors clustered by MSA.

3 The model

We now turn to developing a structural model that can account for the results in the previous section. 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 non-durables 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 non-durables 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 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 the 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. Appendix B reports the first-order conditions for both the borrower and other agents in the model economy.

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). As in Guerrieri and Iacoviello (2017), we allow for inertia in the dynamics of mortgage debt, as measured by $\gamma \in [0;1[$. 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 $\phi > 0$.

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 goods production

The final good, Y_t, is produced by a representative firm using a CES production function that aggregates a continuum of measure one 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, $1/(1-\rho)$.[12] 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 eliminated.[13]

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

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

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

[13] 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.

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 the intermediate good 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, following Jaimovich (2007). Thus, the elasticity of demand for the intermediate 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, $\frac{1}{1-\omega} < \frac{1}{1-\rho}$.[14] This implies that if an individual firm increases its price, $p_t(j,i)$, 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.[15]

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 $\frac{1}{\omega}$.

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

[15] Notice from (3.10) that the sectoral price index is not equal to an average of the individual firms' prices due to the variety effect.

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 into

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

Equations (3.17) and (3.18) yield two insights about the *variety effect*, as well as about its interplay with the *competition effect*. First, p_t/P_t increases in the number of firms. This results from the increased variety lowering marginal costs for 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 goods output more than one-for-one, for given 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 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 results in endogenous procyclical TFP variations through the competition effect and the variety effect. 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 cost, 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 combine to produce a conditional increase in houses 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:

$$R_{t-1} B_{t-1}^g + G_t = \tau_t^{TOT} + B_t^g. \tag{3.24}$$

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:

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

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 (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, any 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 5.3 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.[16] 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.[17]

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: An expansionary shock to public spending stimulates the entry of a new group of producers due to enhanced profit opportunities, thus increasing aggregate productivity. Second, we allow for a certain degree of taste for variety, as captured by τ, which implies the aggregate production technology to exhibit increasing returns to scale in the primary factors of production. While increasing returns to scale via free entry have already been introduced in a model with monopolistic competition by Devereux et al. (1996), they also notice how generating a crowding-in of private consumption is only possible through counterfactually high levels of the markup. Introducing taste for variety à la Benassy (1996) complements this channel, while allowing for plausible values of the markup. In turn, both channels induce labor demand to shift outward, thus allowing to overcome the expansion in labor supply, ultimately increasing the real wage.

To illustrate how these mechanisms formally combine, we solve analytically a simplified version 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[18]

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

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

[16] 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.

[17] 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).

[18] Log-linear variables are hatted.

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

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

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 in this simple model is sufficient to obtain a positive response of the price of housing, is that

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

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:

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

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. 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 $\omega_1 = \left\{\alpha, \beta^b, \beta^l, \delta, \theta, \mu, m, \Xi, \omega, \rho\right\}$ 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 US 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-sector) 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 number 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.[19]

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 a Bayesian Vector Autoregression (BVAR) model estimated for the U.S. economy. To account for potential anticipation effects in the BVAR model, we include the survey-based forecast errors of the growth rate of government spending (denoted by FE_t) computed by Auerbach and Gorodnichenko (2012) to identify truly unexpected government spending shocks. The BVAR model 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 (D_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.[20] The data sample is 1966:Q4-2010:Q3, as dictated by the availability of FE_t from Auerbach

[19] 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.

[20] 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 prioritize the availability of a long data sample.

and Gorodnichenko (2012). We use a standard Minnesota prior and follow the implementation of Giannone et al. (2015). Additional details regarding the data are provided in Appendix A.

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 a diagonal matrix with the inverse of the sample variances of the BVAR-based impulse responses along the diagonal. Effectively, this means that we are attaching higher weights to those impulse responses that are estimated most precisely. We match impulse responses for all variables except net tax revenues, using the responses during the first 25 quarters after the shock.[21]

5.2.1 Estimation results

Panel B of Table 1 reports the estimated parameter values, as well as the associated confidence bands, which are obtained using bootstrap methods based on 1000 replications of our BVAR model.[22] 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 substantially lower than the estimate of Guerrieri and Iacoviello (2017) of 0.7. The estimate of ψ implies a Frisch elasticity of around 3.75, which is not uncommon in business cycle models with flexible prices.[23]

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 very closely 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,

[21] 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.

[22] We obtain 68 percent confidence bands by reporting the 16th and 84th percentiles of the distribution of parameter estimates. The bands are not symmetric, in part because some of the parameter bounds we have imposed in the estimation procedure are reached. This is the case for σ_h, which is bounded below at 0.1; h^b, which is bounded below at 0; ψ, which is bounded below at 0.25; γ, which is bounded above at 0.95; τ, which is bounded above at 5; x, which is bounded below at 1.12; and ρ_τ, which is bounded above at 0.9. Note also that some parameter estimates fall outside their confidence bands. This is possible because the estimates implied by the median impulse response are not necessarily similar to the median of the parameter estimates implied by the distribution of impulse responses.

[23] Note that the upper bound of 4 which we impose on the Frisch elasticity in the estimation is well above microeconometric estimates, but allows traditional RBC models to match business-cycle data (see the discussion by Chetty et al., 2011).

the markup is relatively insensitive to fiscal shocks. Thus, to produce sizable upward changes in TFP—which are key to bring about an equilibrium increase in patient households' nondurable consumption and, thus, house prices—a relatively high τ is necessary, so as to amplify the effect of F_t on TFP.[24] In light of these arguments, we obtain an estimate of $\tau = 4.925$, 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). Moreover, τ is rather imprecisely estimated, as also found by Lewis and Poilly (2012). For these reasons, 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.

[24]Figure D.3 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 obtains. The figure confirms that a lower steady-state markup reduces the minimum value of τ required to obtain an increase in the house price.

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	1.527 [0.704−2.686]
σ_h	Curvature in utility of housing	0.107 [0.100−1.945]
h^l	Habit formation, lenders	0.634 [0.352−0.777]
h^b	Habit formation, borrowers	0.037 [0.000−0.851]
ψ	Inverse Frisch elasticity	0.264 [0.250−9.001]
ϕ	Capital adjustment cost parameter	4.798 [5.383−24.986]
γ	Inertia of mortgage debt	0.366 [0.434−0.950]
τ	Love for variety parameter	4.925 [1.870−5.000]
x	Steady-state value of markup	1.123 [1.120−1.184]
γ_τ	Tax response to government debt	0.238 [0.250−0.823]
ρ_τ	Inertia of tax level	0.377 [0.142−0.900]
γ_G	Persistence of government spending shock	0.971 [0.940−0.989]
σ_g	Std. dev. of government spending shock	0.090 [0.087−0.117]

Note: 68 percent confidence bands for the estimated parameters are reported in brackets.

5.3 Model dynamics

We report the estimated impulse response functions from the model (dashed-blue lines) in Figure 3, alongside their empirical counterparts from the BVAR model. We are able to match the sign and shape of the responses of all variables, with the model-implied responses mostly remaining within the confidence bands from the BVAR model. The increase in the house price in the model is somewhat smaller than in the BVAR—whose magnitude is in the ballpark of the regional estimates obtained via OLS and IV, as shown in Figure 2—while the reverse is true for the responses of output, TFP and, to a smaller extent, the real wage. The positive response of consumption and the wage rate implied by the model are in line with most of the existing literature, e.g. Galí et al. (2007), while the increase in TFP is in line with recent empirical studies by d'Alessandro et al. (2019) and Jørgensen and Ravn (2018), among others. The increase in mortgage debt, which closely matches that in the data, is consistent with 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.

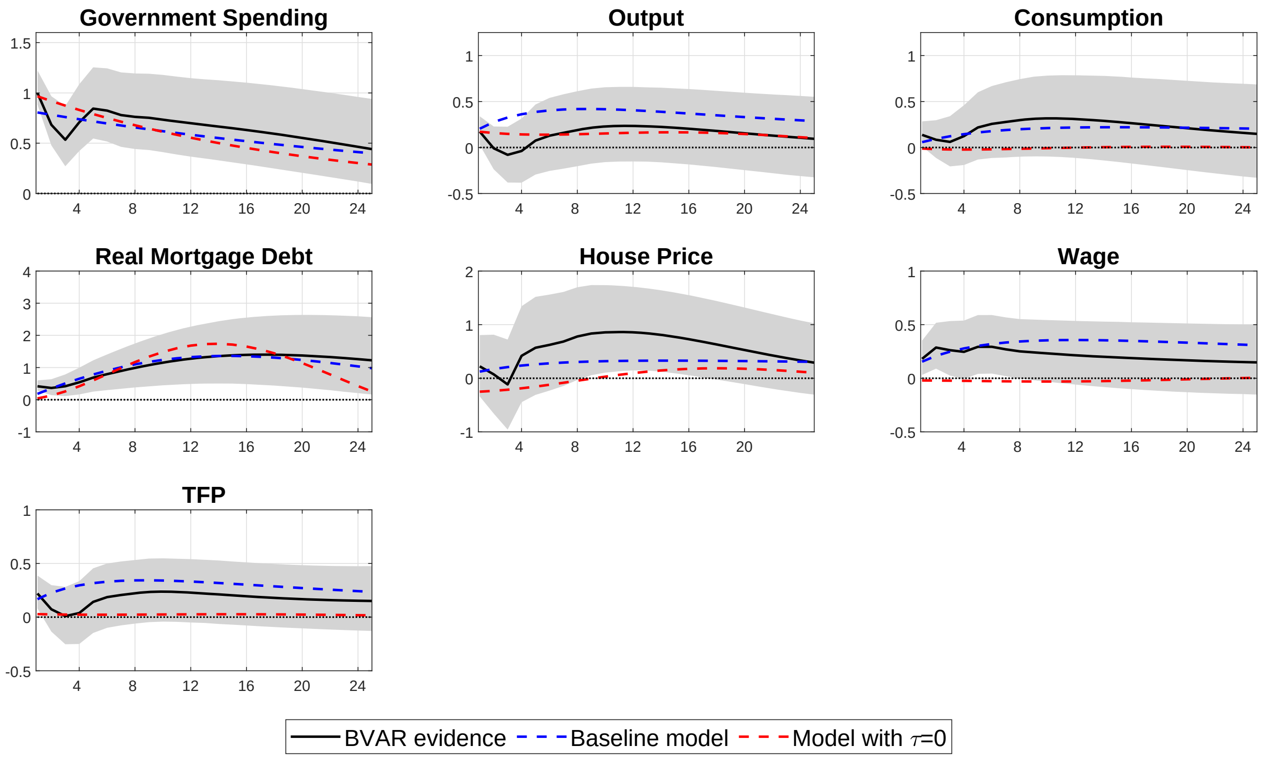
We also estimate a version of the model in which we switch off the taste-for-variety channel, by imposing $\tau = 0$. The estimated impulse responses from this model also appear in Figure 3 (dashed-red lines). As the figure makes it clear, this model version fails to generate the increase in TFP and 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. These results echo those of Devereux et al. (1996) and Lewis and Winkler (2017), who show that, for realistic parameter values, firm entry is *per se* not sufficient to crowd-in aggregate consumption.[25]

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). 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 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.[26] Otherwise, under $\tau = 0$ the contraction in labor supply dominates and the real wage drops. The top row thus confirms the message from the stylized model in Section 4 summarized in condition (4.6): When the taste for variety is sufficiently

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

[26] In fact, 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.

Figure 3: Estimated effects of a government spending shock



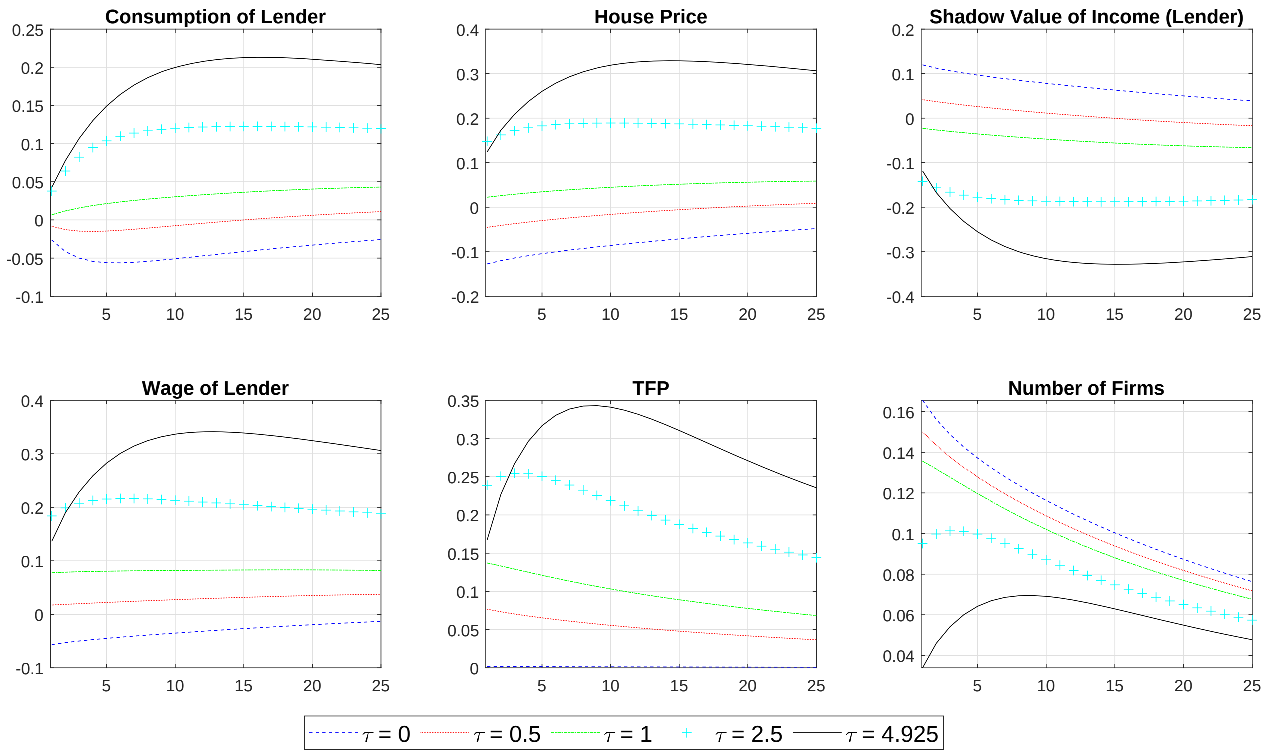
Notes: The figure shows the effects of a shock to government spending. Black line: BVAR model. Grey areas: 68 percent credible sets from BVAR model. Dashed blue line: Estimated DSGE model. Dashed red line: Estimated DSGE model without taste for variety ($\tau = 0$).

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 dynamics 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 strong impact on the relationship between the number of firms and the markup.[27] 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

[27] In fact, as seen from Figure 5, the estimation prefers a value of x for which the response of the number of firms is relatively weak.

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

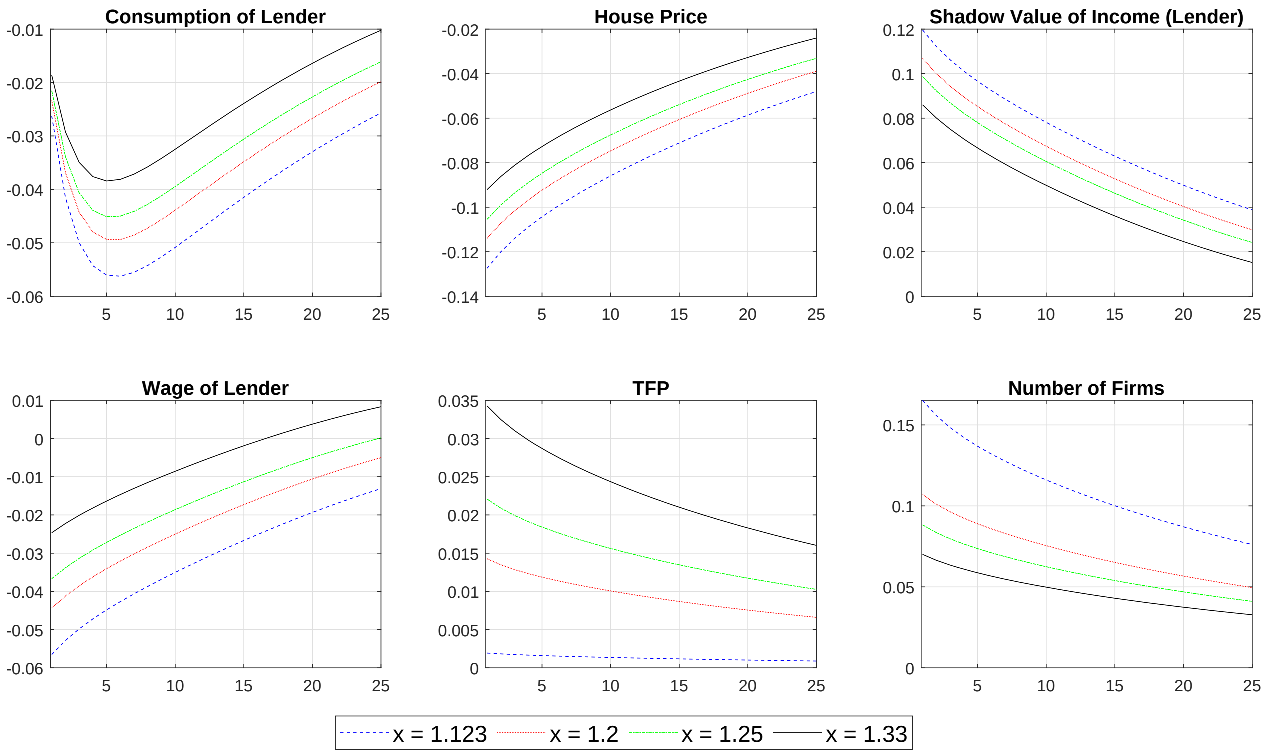


Notes: The figure shows the effects of a shock to government spending for various values of the love-of-variety parameter τ. Dashed line: $\tau = 0$. Dotted line: $\tau = 0.5$. Dashed-dotted line: $\tau = 1$. Crossed line: $\tau = 2.5$. Solid line: $\tau = 4.925$ (estimated value). All other parameters are kept at their baseline values.

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 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

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 line: $x = 1.12$. Dotted line: $x = 1.2$. Dashed-dotted line: $x = 1.25$. Solid line: $x = 1.33$. The love-of-variety parameter τ has been set to zero. All other parameters are kept at their baseline values.

effect, implying that TFP is only affected by a small decline in the markup. These arguments are confirmed by Figure D.4 in Appendix D, which shows that TFP, the number of firms, and the house price increase in tandem when the variety effect is present, and more so the lower is x.

6 Concluding remarks

We report new regional estimates indicating that house prices in the US 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 increases. 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. These features generate increasing returns to scale at the aggregate level that overcome the negative wealth effect induced by an increase in fiscal spending.

We overcome a longstanding limitation of dynamic frameworks featuring Ricardian households that participate to 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. A key feature of our modeling strategy, which consists of disentangling the taste for variety from the degree of market power, is that we may obtain a conditional increase in house prices, while estimating empirically plausible markups.

While our estimated model resolves, at least qualitatively, 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. We leave this challenge for future research.

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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 regional analysis

A.1.1 Data used in the regional analysis

We collected 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 is 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% 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 data set. 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 tracks 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.

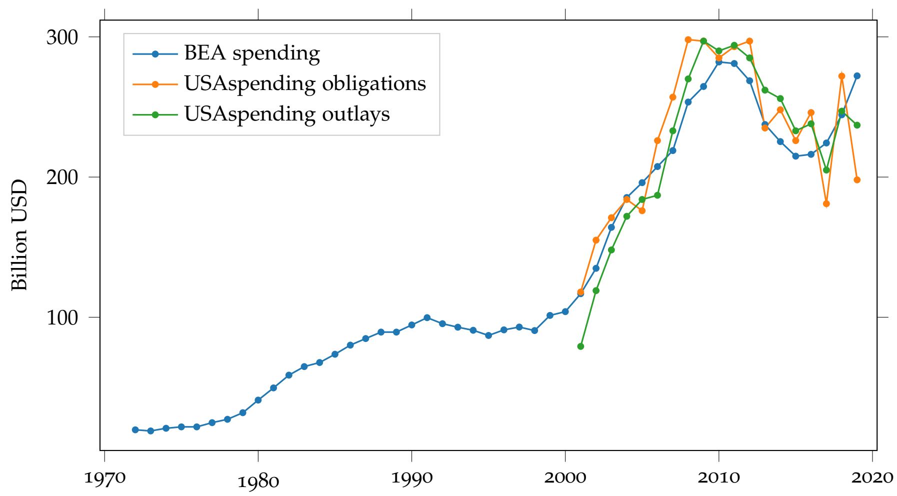
A.1.2 First-stage estimates

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

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

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 regression with

Figure A.1: Military spending according to USAspending 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.

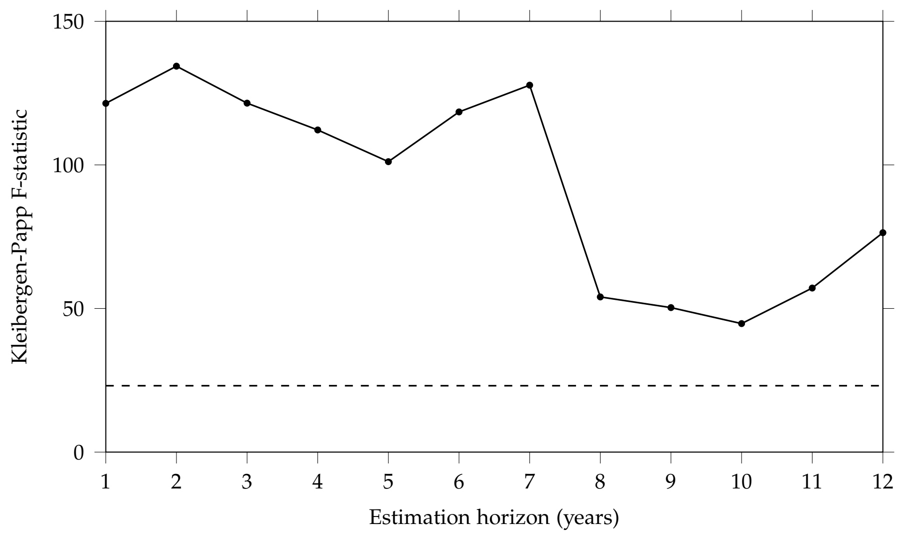
establishment growth in that it has two lags of house price growth as controls, rather than two lags of establishment growth. Although not shown, the *F*-statistics from these two sets of first-stage regressions are almost identical.

A.1.3 Robustness of the regional estimates

This section analyzes the robustness of our regional estimates 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 1 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.1.1 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 and establishment 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. (2018). Lastly, column (8) adds state $\times$ year fixed effects to control for state-specific house price

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



Notes: The figure shows the Kleibergen-Papp *F*-statistics from the first-stage regression (A.1) 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% significant level.

and establishment growth fluctuations.

For the case of the housing response, we 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 by around a half but the estimates are still significant and display a hump-shaped pattern. When using the proxy for outlays instead of obligations, the estimates become substantially larger, which is also the case for the response of establishments.

Turning to the robustness checks to the response of establishments, the estimates that differ from the baseline are those from the specification that normalizes spending by population, and the specification controlling for housing exposure. The latter result seems to be driven by the model being estimated on the subsample of MSAs on which we have data on exposure to aggregate house prices fluctuations. If we estimate the model on this subsample but do not control for this differential exposure, the estimates are broadly similar to those in column (7) but with slightly larger standard errors.

Next, we show the sensitivity of our estimates to the outliers in Table A.2. 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. The estimates from all three specifications display the same hump-shape as the baseline estimates and within the range of the baseline estimates at conventional confidence levels. However, the establishment response is less statistically significant when removing MSAs in the bottom and top of DoD spending distribution, while the standard errors become larger when winsorized observations are removed.

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 lower the estimates across the entire horizon as seen in column (2), which suggests that the lag exogeneity condition of the instrument with itself is not met unless we condition on past values of spending growth and the instrument. 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 dependent variables to the regression with lagged spending and instruments, however, reduces the standard errors (see column (4)), thus improving the efficiency of the estimators. This is especially pronounced for the estimates of the house price response, since there is substantial within-MSA autocorrelation in house price growth.

A.2 Appendix to the BVAR analysis

We estimate a BVAR model with four lags and a constant. The ordering of the variables is the following:

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

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: Government consumption expenditure 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).[28] All series are seasonally adjusted. We convert from nominal to real terms using the GDP deflator (GDPDEF).

D_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.

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 obtain the series directly from Auerbach and Gorodnichenko (2012).

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

Table A.1: Robustness of regional 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
	Dependent variable: House price growth							
1-year horizon	0.18**	0.12	0.44***	0.15**	0.0029*	0.16	0.22***	0.13**
	(0.08)	(0.10)	(0.11)	(0.07)	(0.00)	(0.11)	(0.06)	(0.05)
2-year horizon	0.45***	0.35**	0.95***	0.39***	0.0079***	0.43**	0.52***	0.34***
	(0.15)	(0.17)	(0.21)	(0.12)	(0.00)	(0.20)	(0.12)	(0.11)
4-year horizon	1.18***	1.16***	2.36***	0.95***	0.023***	1.30***	1.14***	0.76***
	(0.27)	(0.30)	(0.45)	(0.21)	(0.01)	(0.41)	(0.30)	(0.23)
6-year horizon	1.59***	1.52***	3.01***	1.27***	0.032***	1.88***	1.32***	0.87***
	(0.33)	(0.33)	(0.62)	(0.27)	(0.01)	(0.50)	(0.44)	(0.27)
10-year horizon	1.00***	0.64***	2.67***	0.74***	0.015***	1.03***	0.79***	0.58***
	(0.21)	(0.20)	(0.59)	(0.17)	(0.00)	(0.27)	(0.30)	(0.18)
	Dependent variable: Establishment growth							
1-year horizon	-0.018	-0.018	-0.015	-0.010	-0.0002	-0.019	0.012	-0.019
	(0.03)	(0.03)	(0.06)	(0.03)	(0.00)	(0.03)	(0.04)	(0.03)
2-year horizon	0.00058	0.0026	0.047	0.0085	0.000085	0.016	0.041	-0.0095
	(0.06)	(0.06)	(0.10)	(0.05)	(0.00)	(0.06)	(0.05)	(0.04)
4-year horizon	0.13*	0.14*	0.29**	0.11*	0.0021	0.20**	0.089	0.14**
	(0.07)	(0.08)	(0.11)	(0.06)	(0.00)	(0.09)	(0.06)	(0.07)
6-year horizon	0.20**	0.21**	0.31**	0.15**	0.0032	0.29**	0.13	0.16*
	(0.10)	(0.10)	(0.15)	(0.08)	(0.00)	(0.12)	(0.09)	(0.09)
10-year horizon	0.039	0.039	0.36***	0.032	-0.0002	0.11	0.13**	0.023
	(0.07)	(0.07)	(0.12)	(0.06)	(0.00)	(0.07)	(0.05)	(0.06)
MSAs	380	380	380	380	380	380	255	373

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.1.1. Column (4) normalizes DoD spending by the BEA's measure of personal income. Column (5) normalizes DoD spending by 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' 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. (2018) 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. 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.2: Robustness of regional estimates (outliers)

	(1) Baseline	(2) Remove extreme DoD shares	(3) Non-winsorized	(4) Remove winsorized
	Dependent variable: House price growth			
1-year horizon	0.18^{**}	0.17^{*}	0.18^{**}	0.23
	(0.08)	(0.09)	(0.07)	(0.15)
2-year horizon	0.45^{***}	0.40^{**}	0.43^{***}	0.59^{**}
	(0.15)	(0.16)	(0.12)	(0.27)
4-year horizon	1.18^{***}	1.24^{***}	1.11^{***}	1.40^{**}
	(0.27)	(0.33)	(0.25)	(0.55)
6-year horizon	1.59^{***}	1.68^{***}	1.43^{***}	1.92^{***}
	(0.33)	(0.39)	(0.34)	(0.66)
10-year horizon	1.00^{***}	1.11^{***}	0.96^{***}	1.85^{***}
	(0.21)	(0.23)	(0.26)	(0.48)
	Dependent variable: Establishment growth			
1-year estimate	-0.018	-0.053^{**}	-0.012	-0.030
	(0.03)	(0.02)	(0.03)	(0.07)
2-year estimate	0.00058	-0.058	0.0072	0.0091
	(0.06)	(0.05)	(0.05)	(0.10)
4-year estimate	0.13^{*}	0.11	0.13^{*}	0.074
	(0.07)	(0.09)	(0.07)	(0.15)
6-year estimate	0.20^{**}	0.16	0.17^{*}	0.30
	(0.10)	(0.11)	(0.09)	(0.19)
10-year estimate	0.039	0.0031	0.050	0.16
	(0.07)	(0.08)	(0.06)	(0.18)
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 estimates (controls)

	(1)	(2)	(3)	(4)
	Dependent variable: House price growth			
1-year estimate	0.25***	0.24*	0.13**	0.18**
	(0.07)	(0.14)	(0.06)	(0.08)
2-year estimate	0.73***	0.49**	0.56***	0.45***
	(0.14)	(0.24)	(0.12)	(0.15)
4-year estimate	2.50***	1.01**	2.41***	1.18***
	(0.44)	(0.44)	(0.49)	(0.27)
6-year estimate	2.93***	1.22**	3.09***	1.59***
	(0.62)	(0.49)	(0.64)	(0.33)
10-year estimate	2.45***	0.82**	2.42***	1.00***
	(0.55)	(0.41)	(0.50)	(0.21)
	Dependent variable: Establishment growth			
1-year estimate	0.017	-0.016	0.017	-0.018
	(0.02)	(0.03)	(0.02)	(0.03)
2-year estimate	0.033	-0.003	0.038	0.00058
	(0.03)	(0.05)	(0.03)	(0.06)
4-year estimate	0.28***	0.12	0.28***	0.13*
	(0.10)	(0.08)	(0.10)	(0.07)
6-year estimate	0.32**	0.17	0.32**	0.20**
	(0.13)	(0.10)	(0.14)	(0.10)
10-year estimate	0.22**	0.054	0.20**	0.039
	(0.10)	(0.08)	(0.10)	(0.07)
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/establishments. 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:

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

Marginal costs are derived by minimizing the following:

$$w_t^l n_t^l(j,i) + w_t^b n_t^b(j,i) + r_t^k k_{t-1}(j,i) + \lambda_t(j,i)\left(m_t(j,i) - k_{t-1}(j,i)^\mu \left[n_t^b(j,i)^\alpha n_t^l(j,i)^{1-\alpha}\right]^{1-\mu} + \varphi\right), \tag{B.31}$$

where $\lambda_t(j,i)$ is the multiplier on the production function.

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

$$r_t^k - \lambda_t(j,i)\mu \frac{k_{t-1}(j,i)^\mu \left[n_t^b(j,i)^\alpha n_t^l(j,i)^{1-\alpha}\right]^{1-\mu}}{k_t(j,i)} = 0, \tag{B.32}$$

$$w_t^b - \lambda_t(j,i)(1-\mu)\alpha \frac{k_{t-1}(j,i)^\mu \left[n_t^b(j,i)^\alpha n_t^l(j,i)^{1-\alpha}\right]^{1-\mu}}{n_t^b(j,i)} = 0, \tag{B.33}$$

$$w_t^l - \lambda_t(j,i)(1-\mu)(1-\alpha) \frac{k_{t-1}(j,i)^\mu \left[n_t^b(j,i)^\alpha n_t^l(j,i)^{1-\alpha}\right]^{1-\mu}}{n_t^l(j,i)} = 0. \tag{B.34}$$

Substituting these into the production function leads to the following:

$$m_t(j,i) = k_{t-1}(j,i)\left(\frac{r_t^k}{w_t^b}\frac{\alpha(1-\mu)}{\mu}\left(\frac{w_t^b}{w_t^l}\frac{1-\alpha}{\alpha}\right)^{1-\alpha}\right)^{1-\mu} - \varphi, \tag{B.35}$$

$$m_t(j,i) = n_t^b(j,i)\left(\frac{\mu}{(1-\mu)\alpha}\frac{w^b}{r^k}\right)^\mu \left(\frac{w_t^b}{w_t^l}\frac{1-\alpha}{\alpha}\right)^{(1-\alpha)(1-\mu)} - \varphi, \tag{B.36}$$

$$m_t(j,i) = n_t^l(j,i)\left(\frac{\mu}{(1-\mu)(1-\alpha)}\frac{w^l}{r^k}\right)^\mu \left(\frac{w_t^l}{w_t^b}\frac{\alpha}{1-\alpha}\right)^{\alpha(1-\mu)} - \varphi. \tag{B.37}$$

which can be inserting into the cost function $C_t(j,i) = w_t^l n_t^l(j,i) + w_t^b n_t^b(j,i) + r_t^k k_{t-1}(j,i)$ to get an expression for costs:

$$C_t(j,i) = A\left(r_t^k\right)^\mu \left(w_t^b\right)^{\alpha(1-\mu)} \left(w_t^l\right)^{(1-\alpha)(1-\mu)} \left(m_t(j,i) + \varphi\right), \tag{B.38}$$

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

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):

$$R = \frac{1}{\beta^l}, \tag{B.39}$$

$$r^k = \frac{1}{\beta^l} - (1-\delta). \tag{B.40}$$

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:

$$\frac{C}{Y} = 1 - \bar{G} - \frac{\delta\mu}{\frac{1}{\beta^l} - (1-\delta)}. \tag{B.41}$$

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

$$\frac{r^k K}{Y} = \mu, \tag{B.42}$$

$$\frac{w^b N^b}{Y} = (1-\mu)\alpha, \tag{B.43}$$

$$\frac{w^l N^l}{Y} = (1-\mu)(1-\alpha). \tag{B.44}$$

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:

$$\frac{\tau^{TOT}}{Y} = \left(\frac{1}{\beta^l} - 1\right)\frac{B^g}{Y} + \frac{G}{Y}, \tag{B.45}$$

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.39) are given by

$$q\lambda^b = \Upsilon^b \left(H^b\right)^{-\sigma_h} + \beta^b\lambda^b q + \mu^b m\beta^l q(1-\gamma) \tag{B.46}$$

$$\mu^b = \lambda^b \frac{1 - \frac{\beta^b}{\beta^l}}{1-\beta^b\gamma} \tag{B.47}$$

Substituting the latter equation into the former yields

$$\Upsilon^b \left(H^b\right)^{-\sigma_h} \quad = \quad q\lambda^b \left[1 - \beta^b - \frac{\beta^l - \beta^b}{1-\beta^b\gamma} m(1-\gamma)\right]. \tag{B.48}$$

The housing demand equation for the patient households is given by

$$\Upsilon^l \left(H^l\right)^{-\sigma_h} = q\lambda^l \left(1 - \beta^l\right) \tag{B.49}$$

The budget constraint multipliers follow from (B.1) and (B.5):

$$\lambda^b = \left(1 - h^b \beta^b\right)\left(\left(1 - h^b\right) C^b\right)^{-\sigma_c}$$
$$\lambda^l = \left(1 - h^l \beta^l\right)\left(\left(1 - h^l\right) C^l\right)^{-\sigma_c}$$

Dividing (B.51) by (B.48) and inserting the steady state expressions for the budget constraint multipliers 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:

$$\begin{aligned}
\frac{Y^l\left(H^l\right)^{-\sigma_h}}{Y^b\left(H^b\right)^{-\sigma_h}} &= \frac{q\lambda^l\left(1-\beta^l\right)}{q\lambda^b\left[1-\beta^b-\frac{\beta^l-\beta^b}{1-\beta^b\gamma}m(1-\gamma)\right]} \\
\left(\frac{H}{H^b}-1\right)^{-\sigma_h} &= \frac{Y^b}{Y^l}\frac{1-\beta^l}{1-\beta^b-\frac{\beta^l-\beta^b}{1-\beta^b\gamma}m(1-\gamma)}\frac{\lambda^l}{\lambda^b} \\
\left(\frac{H}{H^b}-1\right)^{-\sigma_h} &= \frac{Y^b}{Y^l}\frac{1-\beta^l}{1-\beta^b-\frac{\beta^l-\beta^b}{1-\beta^b\gamma}m(1-\gamma)}\frac{\left(1-h^l\beta^l\right)\left(\left(1-h^l\right)\left(C-C^b\right)\right)^{-\sigma_c}}{\left(1-h^b\beta^b\right)\left(\left(1-h^b\right)C^b\right)^{-\sigma_c}} \\
\left(\frac{H}{H^b}-1\right)^{-\sigma_h} &= \frac{Y^b}{Y^l}\frac{1-\beta^l}{1-\beta^b-\frac{\beta^l-\beta^b}{1-\beta^b\gamma}m(1-\gamma)}\frac{1-\beta^l h^l}{1-\beta^b h^b}\left(\frac{1-h^l}{1-h^b}\right)^{-\sigma_c}\left(\frac{C}{C^b}-1\right)^{-\sigma_c}. \qquad \text{(B.50)}
\end{aligned}$$

Similarly, the housing demand equation for the patient household in equation (B.6) can be combined with the budget constraint multiplier from equation (B.5) to get

$$\phi_H^l\left(H^l\right)^{-\frac{1}{\vartheta}} = q\phi_C^l\left(1-\beta^l\right)\left(C^l\right)^{-\frac{1}{\vartheta}}. \qquad \text{(B.51)}$$

Dividing (B.51) by (B.48) and inserting 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:

$$\begin{aligned}
\frac{\phi_H^l}{\phi_H^b}\left(\frac{H^l}{H^b}\right)^{-\frac{1}{\vartheta}} &= \frac{\phi_C^l}{\phi_C^b}\left(\frac{C^l}{C^b}\right)^{-\frac{1}{\vartheta}}\frac{1-\beta^l}{1-\beta^b-\frac{\beta^l-\beta^b}{1-\beta^b\gamma}m\left(1-\gamma\right)} \iff \\
\left(\frac{H}{H^b}-1\right)^{\frac{1}{\vartheta}} &= \frac{\phi_C^b\phi_H^l}{\phi_C^l\phi_H^b}\frac{1-\beta^b-\frac{\beta^l-\beta^b}{1-\beta^b\gamma}m\left(1-\gamma\right)}{1-\beta^l}\left(\frac{C}{C^b}-1\right)^{\frac{1}{\vartheta}}. \qquad \text{(B.52)}
\end{aligned}$$

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.39), the labor income share (B.43), and the lump-sum tax payment (3.22):

$$\frac{C^b}{C} = \frac{Y}{C}\left[\left(\beta^l-1\right)m\frac{qH}{Y}\frac{H^b}{H}+\alpha\left(1-\mu-\frac{\tau^{TOT}}{Y}\right)\right]. \qquad \text{(B.53)}$$

The housing wealth-to-output ratio, $\frac{qH}{Y}$, is calibrated, while the consumption share, $\frac{C}{Y}$, follows from (B.41), so (B.52) and (B.53) 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:

$$\hat{\lambda}_t^b = -\frac{\sigma_c^b}{(1-\beta^b h^b)(1-h^b)}\left(\hat{C}_t^b - h^b \hat{C}_{t-1}^b - \beta^l h^b E_t\left\{\hat{C}_{t+1}^b - h^b \hat{C}_t^b\right\}\right) \tag{B.54}$$

$$\hat{w}_t^b + \lambda_t^b = \psi \hat{N}_t^b \tag{B.55}$$

$$\widehat{B}_t = \gamma \widehat{B}_{t-1} + (1-\gamma)\left(\mathrm{E}_t \widehat{q}_{t+1} + \widehat{H}_t^b - \widehat{R}_t\right) \tag{B.56}$$

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.39), (B.48) and (B.47):

$$\begin{aligned}
q\lambda^b\left(\hat{q}_t + \hat{\lambda}_t^b\right) &= -\sigma_h q \lambda^b\left[1 - \beta^b - \frac{\beta^l - \beta^b}{1-\beta^b\gamma} m(1-\gamma)\right]\hat{H}_t^b + \beta^b \lambda^b q E_t\left\{\hat{\lambda}_{t+1}^b + \hat{q}_{t+1}\right\} \\
&\quad + \lambda^b \frac{1 - \frac{\beta^b}{\beta^l}}{1-\beta^b\gamma} m(1-\gamma)\frac{q}{R} E_t\left\{\hat{\mu}_t^b + \hat{q}_{t+1} - \hat{R}_t\right\}, \\
\hat{q}_t + \hat{\lambda}_t^b &= -\sigma_h\left[1 - \beta^b - \frac{\beta^l - \beta^b}{1-\beta^b\gamma} m(1-\gamma)\right]\hat{H}_t^b + \beta^b E_t\left\{\hat{\lambda}_{t+1}^b + \hat{q}_{t+1}\right\} \\
&\quad + \frac{\beta^l - \beta^b}{1-\beta^b\gamma} m(1-\gamma) E_t\left\{\hat{\mu}_t^b + \hat{q}_{t+1} - \hat{R}_t\right\}.
\end{aligned} \tag{B.57}$$

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.47) results in

$$\widehat{\lambda}_t^b + \beta^b \gamma \frac{1 - \frac{\beta^b}{\beta^l}}{1-\beta^b\gamma} E_t\left\{\widehat{\mu}_{t+1}^b\right\} = \frac{1 - \frac{\beta^b}{\beta^l}}{1-\beta^b\gamma}\widehat{\mu}_t^b + \frac{\beta^b}{\beta^l} E_t\left\{\widehat{\lambda}_{t+1}^b + \widehat{R}_t\right\}. \tag{B.58}$$

The log-linearized budget constraint becomes

$$\begin{aligned}
&\frac{C^b}{C}\frac{C}{Y}\hat{C}_t^b + \frac{qH}{Y}\frac{H^b}{H}\left(\hat{H}_t^b - \hat{H}_{t-1}^b\right) + m\frac{qH}{Y}\frac{H^b}{H}\left(\hat{R}_{t-1} + \hat{B}_{t-1}\right) = \\
&(1-\mu)\alpha\left(\hat{w}_t^b + \hat{N}_t^b\right) + m\frac{qH}{Y}\frac{H^b}{H}\beta^l \hat{B}_t - \alpha\frac{\tau^{TOT}}{Y}\widehat{\tau}_t^{TOT}.
\end{aligned} \tag{B.59}$$

B.5.2 Optimality conditions of the patient households

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

$$\hat{\lambda}_t^l = -\frac{\sigma_c^l}{(1-\beta^l h^l)(1-h^l)}\left(\hat{C}_t^l - h^l \hat{C}_{t-1}^l - \beta^l h^l E_t\left\{\hat{C}_{t+1}^l - h^l \hat{C}_t^l\right\}\right), \tag{B.60}$$

$$\hat{\lambda}_t^l = E_t\left\{\hat{\lambda}_{t+1}^l\right\} + \hat{R}_t, \tag{B.61}$$

$$\hat{w}_t^l + \hat{\lambda}_t^l = \psi \hat{N}_t^l, \tag{B.62}$$

$$\hat{K}_t = (1-\delta)\hat{K}_{t-1} + \delta \hat{I}_t. \tag{B.63}$$

Similar to the derivation for the impatient households, log-linearization of (B.6) in combination with (B.51) becomes

$$\hat{q}_t + \hat{\lambda}_t^l = -\sigma_h\left(1-\beta^l\right)\hat{H}_t^l + \beta^l E_t\left\{\hat{q}_{t+1} + \hat{\lambda}_{t+1}^l\right\}. \tag{B.64}$$

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.63) results in

$$\phi\left(\hat{K}_t - \hat{K}_{t-1}\right) + \hat{\lambda}_t^l = E_t\left\{\hat{\lambda}_{t+1}^l + \beta^l r^k \hat{r}_{t+1}^k + \beta^l\phi\left(\hat{K}_{t+1} - \hat{K}_t\right)\right\}. \tag{B.65}$$

Lastly, the budget constraint (3.4) becomes

$$\begin{aligned}&\frac{C^l}{C}\frac{C}{Y}\hat{C}_t^l + \frac{qH}{Y}\frac{H^l}{H}\left(\hat{H}_t^l - \hat{H}_{t-1}^l\right) + \frac{\delta K}{Y}\hat{I}_t + m\frac{qH}{Y}\frac{H^b}{H}\beta^l\hat{B}_t + \Xi\hat{B}_t^g = \\ &(1-\mu)(1-\alpha)\left(\hat{w}_t^l + \hat{N}_t^l\right) + m\frac{qH}{Y}\frac{H^b}{H}\left(\hat{R}_{t-1} + \hat{B}_{t-1}\right) + \mu\left(\hat{r}_t^k + \hat{K}_{t-1}\right) \\ &+ \frac{1}{\beta^l}\Xi\left(\hat{R}_{t-1} + \hat{B}_{t-1}^g\right) - (1-\alpha)\frac{\tau_t^{TOT}}{Y}\hat{\tau}_t^{TOT}.\end{aligned} \tag{B.66}$$

B.5.3 Symmetric firm equilibrium conditions

The log-linearized factor prices read as

$$\hat{r}_t^k = (1+\tau)\left(\left(\left(\mu - \frac{1}{1+\tau}\right)\hat{K}_{t-1} + (1-\mu)\left(\alpha\hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t, \tag{B.67}$$

$$\hat{w}_t^b = (1+\tau)\left(\left(\mu\hat{K}_{t-1} + (1-\mu)\left(\left(\alpha - \frac{1}{(1+\tau)(1-\mu)}\right)\hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t, \tag{B.68}$$

$$\hat{w}_t^l = (1+\tau)\left(\left(\mu\hat{K}_{t-1} + (1-\mu)\left(\alpha\hat{N}_t^b + \left(1-\alpha-\frac{1}{(1+\tau)(1-\mu)}\right)\hat{N}_t^l\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t. \tag{B.69}$$

while log-linearization of (3.20) results in

$$\hat{Y}_t = (1+\tau)\left(\left(\mu\hat{K}_{t-1} + (1-\mu)\left(\alpha\hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t. \tag{B.70}$$

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

$$F_t = \frac{x_t - 1}{x_t\varphi}K_{t-1}^{\mu}\left[\left(N_t^b\right)^{\alpha}\left(N_t^l\right)^{1-\alpha}\right]^{1-\mu}.$$

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

$$\hat{F}_t = \frac{1}{1+\tau}\left(\hat{Y}_t + \frac{x}{x-1}\hat{x}_t\right). \tag{B.71}$$

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

$$\hat{F}_t = \frac{x}{x-1}\frac{\rho-1}{\rho x-1}\hat{x}_t. \tag{B.72}$$

The log-linearized expression for TFP is

$$T\hat{F}P_t = \tau\hat{F}_t - \hat{x}_t. \tag{B.73}$$

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

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

$$\frac{1}{\beta^l}\Xi Y\left(\hat{R}_{t-1} + \hat{B}_{t-1}^g\right) + \theta\hat{G}_t = \frac{\tau^{TOT}}{Y}\hat{\tau}_t^{TOT} + \Xi\hat{B}_t^g, \tag{B.74}$$

while the adjustment rule for the tax level is given by

$$\hat{\tau}_t^{TOT} = \rho_\tau\hat{\tau}_{t-1}^{TOT} + (1-\rho_\tau)\gamma_\tau\left(\hat{B}_{t-1}^g - \hat{Y}_{t-1}\right). \tag{B.75}$$

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

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

$$\hat{C}_t = \frac{C^b}{C}\hat{C}_t^b + \frac{C^l}{C}\hat{C}_t^l, \tag{B.76}$$

$$0 = \frac{H^b}{H}\hat{H}_t^b + \frac{H^l}{H}\hat{H}_t^l. \tag{B.77}$$

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

$$\hat{G}_t = \gamma_g\hat{G}_t + \epsilon_{g,t}. \tag{B.78}$$

The 25 equations (B.54) to (B.78) define the log-linearized model for the 25 endogeneous variables $\hat{\lambda}_t^b$, $\hat{\lambda}_t^l$, $\hat{C}_t^b$, $\hat{C}_t^l$, $\hat{w}_t^b$, $\hat{w}_t^l$, $\hat{N}_t^b$, $\hat{N}_t^l$, $\hat{H}_t^b$, $\hat{H}_t^l$, $\hat{B}_t$, $\hat{q}_t$, $\hat{R}_t$, $\hat{\mu}_t^b$, $\hat{G}_t$, $\hat{\tau}_t^{TOT}$, $\hat{B}_t^g$, $\hat{K}_t$, $\hat{I}_t$, $\hat{r}_t^k$, $\hat{x}_t$, $\hat{F}_t$, $\hat{Y}_t$, $\hat{C}_t$, 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.76), (B.62), (B.69), (B.70), (B.71), and (B.72), 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

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

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

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

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 estimations or simulations of the quantitative model presented in Section 5. 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 $\tau = 0$

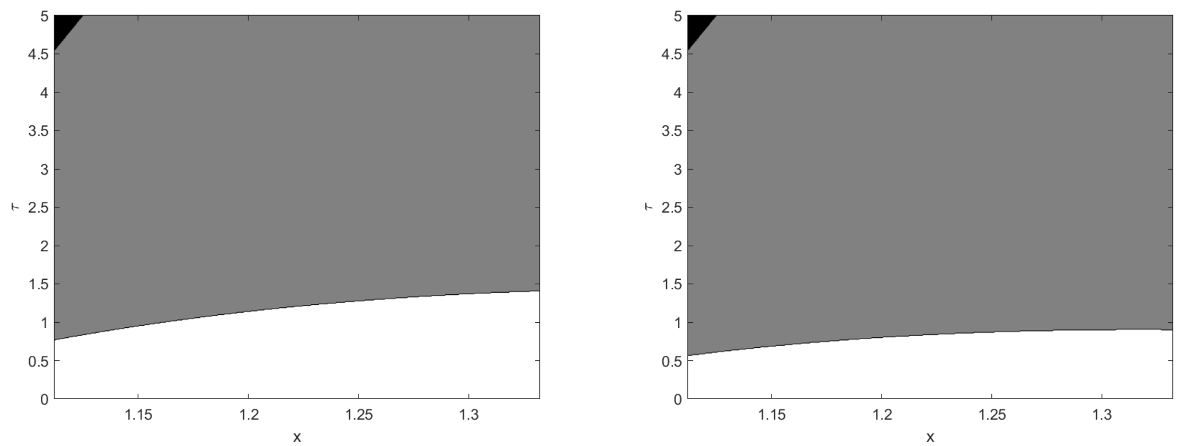
Parameter	Description	Value
σ_c	Curvature in utility of consumption	4.929 [0.000−10.448]
σ_h	Curvature in utility of housing	1.939 [0.000−4.703]
h^l	Habit formation, lenders	0.659 [0.000−1.578]
h^b	Habit formation, borrowers	0.896 [0.187−1.605]
ψ	Inverse Frisch elasticity	0.262 [0.250−1.652]
ϕ	Capital adjustment cost parameter	22.755 [0.000−55.491]
γ	Inertia of mortgage debt	0.946 [0.830−0.950]
x	Steady-state value of markup	1.324 [1.120−1.330]
γ_τ	Tax response to government debt	0.870 [0.000−0.900]
ρ_τ	Inertia of tax level	0.879 [0.686−0.900]
γ_G	Persistence of government spending shock	0.951 [0.938−0.964]
σ_g	Std. dev. of government spending shock	0.098 [0.093−0.103]

Note: 68 percent confidence bands for the estimated parameters are reported in brackets.

We then return to the baseline model characterized by the parameter estimates reported in Section 5.2.1. We rely on Figure D.3 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.3: House price response for different parameter combinations

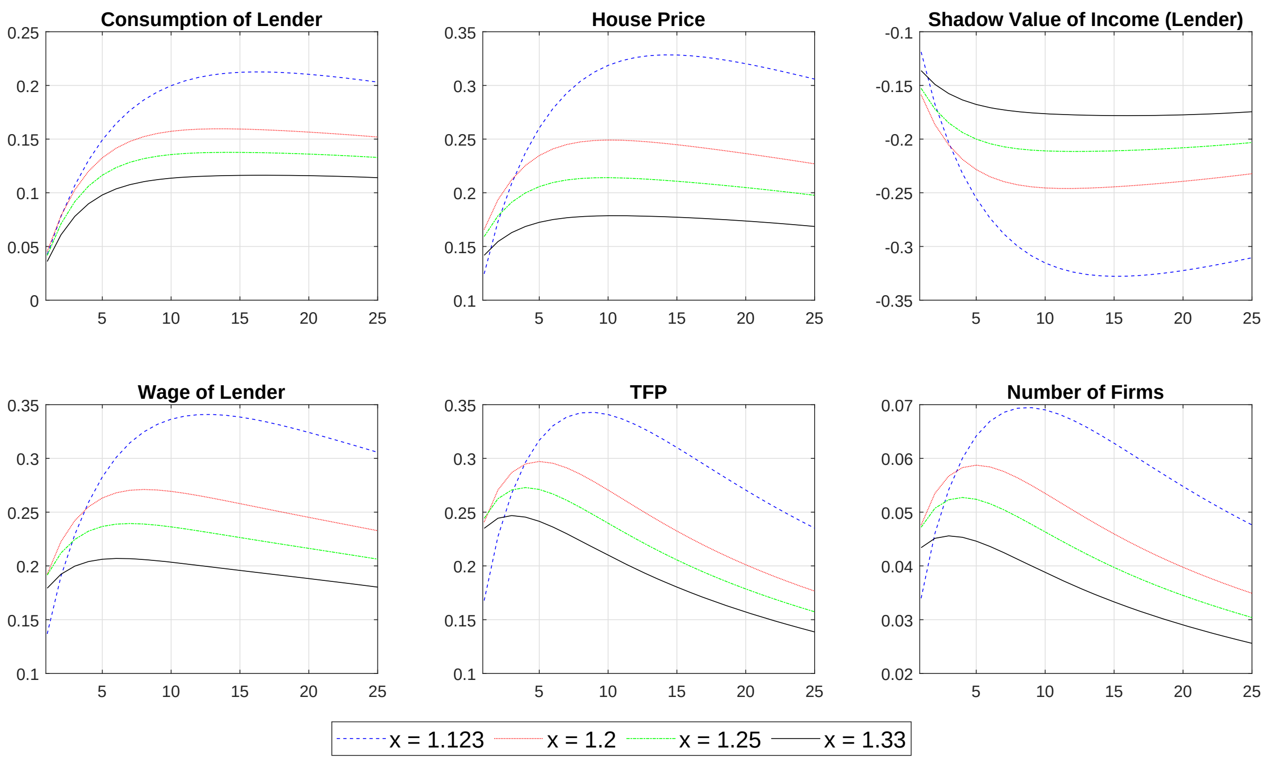
(a) Impact response (b) Cumulative response



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 black area indicates combinations for which the model does not have a unique and determinate solution. 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.4: 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 line: $x = 1.123$ (estimated value). Dotted line: $x = 1.2$. Dashed-dotted line: $x = 1.25$. Solid line: $x = 1.33$. All other parameters are kept at their baseline values.