House Prices, Endogenous Productivity, and the Effects of Government Spending Shocks*

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Abstract

We present aggregate and regional evidence showing that U.S. house prices increase persistently in response to positive shocks to fiscal spending. In sharp contrast to this, house prices decline in conventional dynamic general equilibrium models, where shocks that have short-lived effects on the shadow value of housing inevitably generate negative comovement between households' marginal utility of consumption and house prices (see Barsky et al., 2007). In response to an increase in government spending, the negative wealth effect exerted by the simultaneous increase in the present-value tax burden increases the marginal utility of consumption. Even overcoming the consumption crowding-out puzzle is not sufficient to resolve this shortcoming. To tackle this problem, we extend an otherwise standard model embedding a lender-borrower relationship with alternative—yet, potentially complementary—propagation channels that leverage the expansion in total factor productivity stemming from a positive shock to fiscal spending, so as to contrast the negative wealth effect of higher taxes. This class of models succeeds in generating a persistent expansion in house prices, although the propagation required to match the data is stronger—in some cases significantly so—than what is typically found in the literature. The positive interplay between house prices and productivity finds support in both aggregate and regional data.

Keywords: Fiscal Policy, House Prices, Endogenous Total Factor Productivity.

JEL classifications: E13, E20, E32, E62.

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1 Introduction

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

The contribution of this paper is twofold. First, we devise a Bayesian Vector Autoregression (BVAR) model of aggregate U.S. data to document that an expansion in fiscal spending generates a persistent increase in house prices alongside a range of other macroeconomic variables. We corroborate this finding using contract data from the Department of Defense (DoD), reporting that a positive change in federal government spending in a given city expands the price of housing relative to other cities. Thus, our second and main contribution consists of devising and validating a set of mechanisms that produce a coherent transmission of shocks to fiscal spending on the variables involved in both our aggregate and regional evidence, with a special focus on house prices. The expansionary impact of fiscal spending on house prices has been documented in existing empirical work based on aggregate data (see Khan and Reza, 2017; Alpanda et al., 2021). From a theoretical viewpoint, however, Khan and Reza (2017) have shown that this finding represents a significant challenge, as it is extremely difficult to reproduce within conventional dynamic stochastic general equilibrium (DSGE) frameworks.

A major shortcoming of these models is that house prices contract in the face of a fiscal expansion. Why is this the case? As originally highlighted by Barsky et al. (2007), housing is a long-lived durable and, as such, it features an approximately constant shadow value. Two key elements lie behind this property. First, the marginal utility of housing depends on the stock of housing, which is weakly affected by changes in the flow. Second, temporary shocks—such as those to government spending—typically exert little influence on the future marginal utility of housing. As a result, in any model where Ricardian households participate in the housing market and ultimately determine property prices, the shadow value of their income will inevitably exhibit a negative conditional comovement with house prices.[1] Following an increase in government spending, the present value of lifetime after-tax income drops, thus raising the shadow value of income, ultimately reducing consumption and house prices. As discussed by Khan and Reza (2017), any conventional remedy proposed within otherwise standard Ricardian settings—such

[1] 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), or by assuming that Ricardian households own a fixed share of the aggregate housing stock (see, e.g., Justiniano et al., 2019).

as restrictions on housing supply, nominal stickiness, deep habits, and complementarity between private and public consumption—proves inadequate in breaking the quasi-constancy property of Ricardian households' shadow value of housing, even when producing consumption crowding-in. Furthermore, considering an environment in which housing is priced by financially constrained agents does not resolve the issue, as long as housing works as an idealized durable good whose stock responds weakly to fiscal shocks. We demonstrate this fact in a tractable theoretical model, where households purchase both nondurables and housing, the latter also serving as collateral for borrowing.

To overcome this structural limitation, we rest on the observation that positive shocks to fiscal spending induce an increase in productivity, along with house prices. This is a robust regularity in both aggregate and regional data. Therefore, we assess the capacity to produce a conditional increase in house prices in a battery of frameworks embedding this comovement through alternative—yet, potentially complementary—supply-side propagation channels. To this end, we lay down a flexible-price framework that incorporates a lender-borrower relationship among households with different attitudes to discount the future, and a government responsible for imposing household (lump-sum) taxation to finance wasteful spending on consumption goods produced by firms operating under perfect competition. In turn, this baseline setting is extended to accommodate four different mechanisms that channel fiscal shocks through endogenous movements in Total Factor Productivity (TFP). Three of these are taken off-the-shelf from the existing literature: productive government spending (Aschauer, 1989; Baxter and King, 1993), learning by doing (Chang et al., 2002; d'Alessandro et al., 2019), and variable technology utilization (Bianchi et al., 2019; Jørgensen and Ravn, 2022). In addition, we devise a fourth model structure that combines endogenous entry and exit with some 'taste for variety', whereby spreading a given amount of capital and labor inputs within a sector over an increasing number of firms expands productivity and total output, thus generating increasing returns to scale in the aggregate.

The propagation of shocks to fiscal spending within the alternative models we consider bears the potential to overcome the negative wealth effect induced by an increase in fiscal spending (financed either through a tax hike or an increase in government debt), so that Ricardian households' shadow value of income drops. How is this possible? To address this question, it is instructive to examine how the labor market equilibrium is attained in a standard neoclassical model without any supply-side mechanisms that may induce positive comovement between public spending and productivity. In this case, an expansion in government spending leads to an increase in labor supply, at given factor prices. Holding productivity fixed typically implies a fall in the real wage, which exacerbates the drop in the present value of disposable income. As a result, households' shadow value of income expands, thus driving the house price down. When productivity instead expands sufficiently—through either of the four mechanisms introduced above—real wages expand too, and so consumption and house prices.

While the first three propagation channels under consideration have previously been employed in the study of government spending shocks, the mechanism we propose deserves further explanation. In the presence of endogenous firm turnover, a fiscal expansion leads to enhanced profit opportunities, determining an increase in the number of intermediate goods producers. This compresses their price markups, and expands TFP and labor demand. This is the so-called *competition effect* studied, e.g., by Lewis and Winkler (2017), who also report an increase in net firm entry after a government spending shock in a structural VAR model using U.S. data.[2] However, in line with the literature we discuss below, this channel may only operate at the expense of setting the steady-state markup and/or the Frisch elasticity of labor supply at implausibly high values. Therefore, we complement this mechanism with taste for variety à la Benassy (1996). This implies that, as the number of intermediate goods producers within a given sector increases, the aggregate sectoral good expands for a given input of intermediate goods; the *variety effect*. This twist enhances the model's capacity to induce a joint increase in nondurable consumption and house prices, for given values of the markup and the Frisch elasticity. Combining the variety and the competition effects within our quantitative setting is capable of generating an increase in TFP, business formation, and the wage rate, in response to a fiscal stimulus. Altogether, this induces households to substitute out of leisure and into both nondurables and housing.

To discipline our battery of models, we match their impulse responses to the empirical ones from the BVAR, which features house prices, output, consumption, TFP, mortgage debt, and the real wage, along with federal government spending and tax revenues. In line with the empirical evidence, each model produces a sustained increase in the house price, alongside output and consumption, spurred by a rise in TFP and the real wage. Each of the models can account for more than half of the cumulative increase in house prices observed in aggregate data during the first six years after the shock. Nevertheless, we find that the estimation procedure tends to select parameter combinations that overestimate the responses of TFP and wages, in order to boost the magnitude of the house-price response. As a reflection of this, the estimated strength of each of the supply-side propagation mechanisms is higher—in some cases significantly so—than what is typically found in existing studies. For this reason, we also estimate versions of each model where some plausible parameter restrictions are imposed. In this case, the house-price responses are significantly dampened in each case, with the model featuring love of variety now producing the most robust increase. Finally, we confirm that letting the house price be determined by the financially constrained household has little or no quantitative impact. Overall, our results indicate that propagation of fiscal shocks through endogenous productivity is crucial for matching the sign

[2] We confirm this finding when including business formation in the BVAR. In the regional data, we find that an increase in federal government spending in a given city expands both the number of establishments and labor productivity, relative to other cities, and that cities characterized by a sharper increase in house prices also display sharper hikes in business formation and labor productivity. This relationship is also supported by Epstein et al. (2023), 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.

of the house-price response observed in the data.

Related literature We contribute to a large literature on the macroeconomic effects of shocks to government spending, as surveyed extensively by Ramey (2016). The response of house prices, however, has received little attention. The two studies most closely related to ours are those of Khan and Reza (2017) and Alpanda et al. (2021), who both report a positive house-price response based on structural VAR models for the U.S. economy. We expand on these studies by documenting that this response coexists with increases in TFP and real wages—which play a key role in our proposed theoretical mechanism—and we corroborate these findings by reporting novel evidence based on regional data. As discussed above, Khan and Reza (2017) consider a range of approaches employed in the literature to obtain a positive response of private consumption to a government spending shock, finding that none of these can produce an increase in house prices. Instead, they show that augmenting the monetary policy rule with a negative reaction to government spending can do the job. While such an assumption may complement our approach, two considerations lead us to search for other mechanisms: First, Khan and Reza (2017) obtain only a very short-lived increase in house prices, in contrast to the persistent and hump-shaped response found in the data; and second, this mechanism cannot account for the observed increase in TFP. While Alpanda et al. (2021) also study the response of house prices, their main focus is on the heterogeneous consumption responses of different types of households, empirically as well as theoretically. Our theoretical apparatus yields predictions consistent with their results.

Our paper also adds to an emerging literature emphasizing the potential supply-side effects associated with government spending, in addition to the more traditional focus on aggregate demand. Several recent papers have documented a positive response of TFP to an increase in government spending, and have offered theoretical explanations based on endogenous technology adoption and/or learning-by-doing (see d'Alessandro et al., 2019; Jørgensen and Ravn, 2022; Ilzetzki, 2024). Others have emphasized the productivity-enhancing effects of government spending on research and development (see Antolin-Diaz and Surico, 2022; Moretti et al., 2022; Elfsbacka Schmöller, 2022). We show how some of these mechanisms may be leveraged to ultimately produce a coherent transmission of fiscal shocks onto house prices and other variables.

Our modelling contribution—the combination of endogenous entry and exit with taste for variety—is related to previous studies of the link between government spending shocks, net firm entry, and consumption crowding-in by Devereux et al. (1996) and Lewis and Winkler (2017), although none of them focus on house prices. In Devereux et al. (1996), firm entry generates increasing returns to specialization. This effect is closely related to the variety effect in our model, but it is distinct from the competition effect.[3] Lewis and Winkler (2017) implicitly allow

[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).

for a separate parameter in the production function of industry goods that indexes the variety effect, but constrain it to be in line with the Dixit and Stiglitz (1977) benchmark, thus choosing to focus on the competition 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 insights (in this respect, see also Bilbiie, 2011).

We also contribute to a recent literature combining endogenous firm entry with taste for variety. This builds in large part on Bilbiie et al. (2012), who show that incorporating these ingredients improves the empirical performance of standard Real Business Cycle (RBC) models in response to productivity shocks. We use a variant of the Constant Elasticity of Substitution (CES) function with generalized love of variety introduced by Benassy (1996). This function disentangles market power from love of variety, so 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. Other authors have 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).

Structure The paper proceeds as follows. Section 2 reports empirical evidence based on aggregate and regional data on the response of U.S. house prices in the face of shocks to fiscal spending. Section 3 frames the structural problem that impairs standard frameworks in generating a plausible reaction of house prices, and devises a stylized model embedding endogenous transmission of government spending shocks to TFP and, thus, house prices. Section 4 outlines the baseline model to be employed in the quantitative analysis, while Section 5 provides an overview of the endogenous-productivity mechanisms adopted to generate a coherent transmission of fiscal shocks. Section 6 describes the calibration and estimation of the model(s). Section 7 discusses the main qualitative and quantitative implications. Section 8 concludes.

2 Empirical evidence

In this section we provide empirical evidence to support the claim that increases in government spending have a positive effect on U.S. house prices. We first look at aggregate data, so as to establish a benchmark for the quantitative assessment of the frameworks we will introduce. To this end, we devise a structural BVAR, following the tradition of most of the empirical literature on the aggregate effects of government spending shocks. While we are mainly interested in the aggregate effects of fiscal policy—and primarily in the resolution of the puzzle concerning the response of house prices within DSGE models—we then seek to corroborate our findings by means of regional data. In this respect, we study how a change in federal government spending in a given city—as compared with other cities—affects relative house price movements. We do so by following the approach of Nakamura and Steinsson (2014) and Auerbach et al. (2020b), where military procurement is used as a source of regional variation in spending. In addition to house prices, both types of analysis focus on the conditional behavior of productivity, which is key to validating the transmission mechanism embodied by the quantitative frameworks we will consider.

2.1 Aggregate evidence

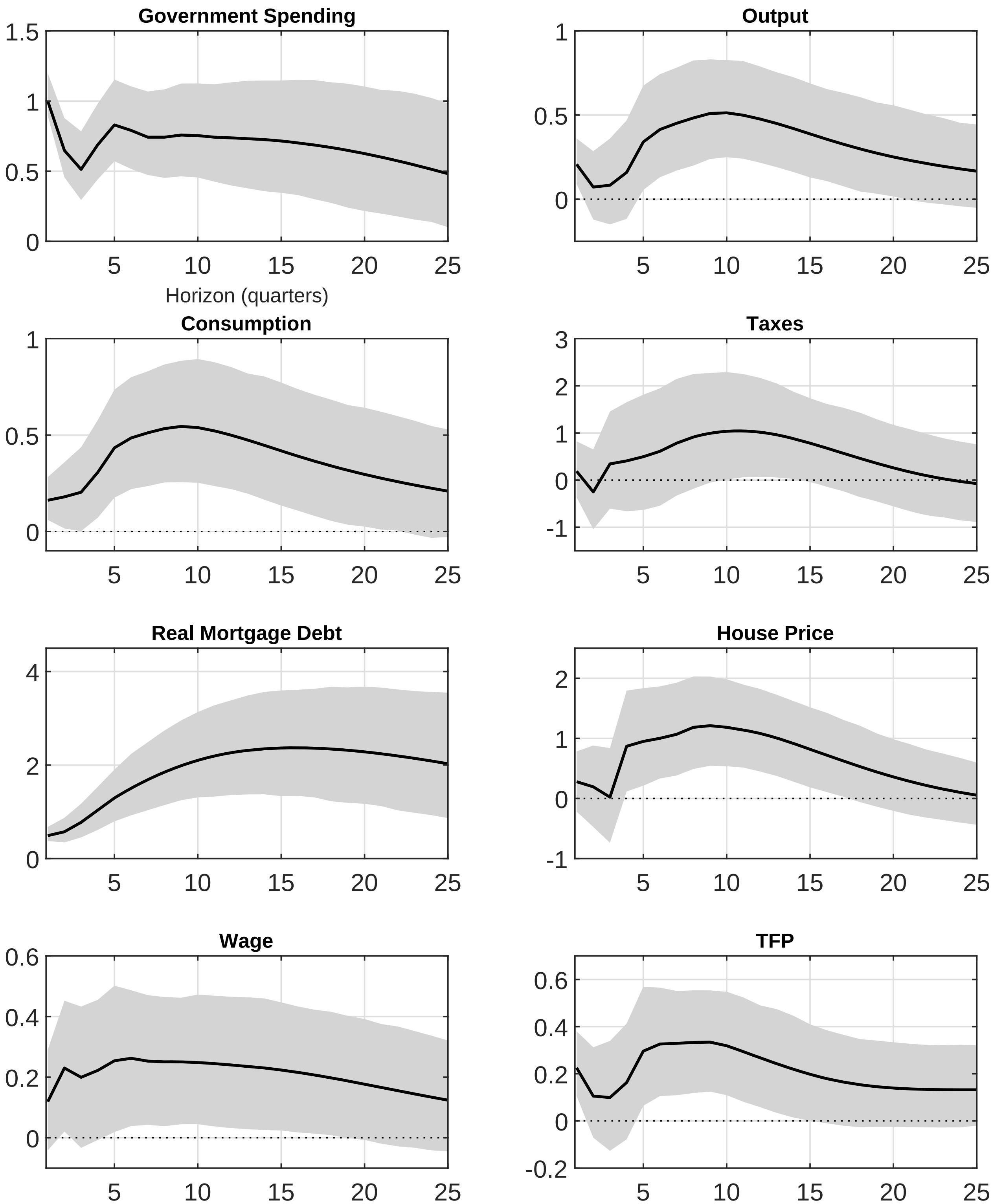
We set up a BVAR model for the U.S. economy. The Bayesian approach is motivated by the fairly large number of variables we consider. In order to identify truly unexpected government spending shocks that do not suffer from potential anticipation effects, we rely on the survey-based forecast errors of the growth rate of government spending (denoted by FE_t) computed by Auerbach and Gorodnichenko (2012). We order this variable first in the BVAR model, thus identifying an unanticipated government spending shock as an innovation to the forecast error of the growth rate of government spending. The BVAR further includes the following variables: Real government consumption and investment (G_t), real gross domestic product (GDP) (Y_t), real private consumption (C_t), real net tax revenues (T_t), real mortgage debt (B_t), the real house price (Q_t), the real wage (W_t), and total factor productivity (TFP_t).[4] 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.[5] The data span the period from 1966:Q4 to 2019:Q4, with the start of the sample dictated by the availability of FE_t from Auerbach and Gorodnichenko (2012), and its end chosen so as to avoid the turbulence of Covid-19. We use the update of the forecast error series of Auerbach and Gorodnichenko (2012) until 2019 constructed by Jørgensen and Ravn (2022). The BVAR includes four lags and a constant, and we detrend all variables except FE_t using a linear and a quadratic trend. We use a standard Minnesota prior, with the tightness of the prior determined by the data using the method of Giannone et al. (2015). Additional details regarding the BVAR and the data are provided in Appendix A.1.

In Figure 1 we present the impulse responses from the BVAR to a positive government spending shock normalized to 1 percent, along with 68 percent credible sets. Following such a shock, we observe a persistent and hump-shaped increase in the house price. This finding is consistent with the evidence reported by Khan and Reza (2017) and Alpanda et al. (2021). To get a sense

[4] Our results are almost identical if we use only real government consumption as our measure of G_t.

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

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



Notes: The figure shows the effects of a shock to government spending estimated in the BVAR model (solid black line), with the grey areas representing the 68 percent credible sets (i.e., the 16th and 84th percentiles of the posterior distribution based on 1,000 draws).

of the magnitude of the effect, we can divide the responses by the sample average of the ratio of government spending to GDP (which equals 0.237). This results into a house-price response that peaks at around 5 percent, two years after an increase in government spending equal to 1 percent of GDP. Furthermore, the positive responses observed for output, consumption, and the wage rate are in line with most of the existing literature, e.g. Galí et al. (2007), while the increase in mortgage debt corroborates recent evidence reported by Auerbach et al. (2020a) and Bayer et al. (2023), who both find that government spending has a stimulative effect on credit markets. Finally, the observed increase in TFP is consistent with recent studies by d'Alessandro et al. (2019) and Jørgensen and Ravn (2022), among others. In the regional data we employ in the next subsection, we consider the response of labor productivity instead of TFP, since the latter is not available at the regional level. As documented in Appendix A.1.2, the responses of TFP and labor productivity are very similar at the aggregate level.

As an extension of the analysis, we consider the response of business formation to a fiscal shock, since this will play an important role in the model with endogenous firm entry and love of variety that we propose below. However, extending the BVAR with a measure of business formation is made difficult by the lack of consistent data at the quarterly frequency spanning the entire sample under consideration. In Appendix A.1.3 we consider the conditional responses of two measures of business formation, each of them available only for a shorter subsample. Both display a positive response, consistent with Lewis and Winkler (2017). However, we also document that while the forecast error series of Auerbach and Gorodnichenko (2012) constitutes a strong and relevant instrument for government spending over the full sample, it does not pass conventional weak instruments tests for each of the two subsamples. Thus, while the available evidence does support the prediction of our theoretical model, we wish to treat these results with caution.

2.2 Regional evidence

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

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

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

where $Q_{i,t}$ denotes the house price index, $G_{i,t}$ is DoD spending, $X_{i,t}$ is a vector of controls, and $Y_{i,t}$ is GDP (all at the MSA level). The MSA fixed effect, $\alpha_{i,h}$, controls for MSA-specific trends in house prices, while the time fixed effect, η_{t+h}, controls for common, national variation in house prices.[6] All variables are measured in nominal terms, although we obtain similar results upon transforming $Q_{i,t}$, $G_{i,t}$, and $Y_{i,t}$ to real terms using the MSA-level GDP deflator.[7]

The coefficient of interest is β_h, which measures the growth in house prices from t to $t+h$ relative to other MSAs, as a result of a 1 percent increase in DoD spending from period t to $t+1$, and relative to period-t GDP.[8] However, the OLS estimate of β_h is likely to be biased, since military contracts tend to flow disproportionately more to areas that experience relatively bad economic outcomes, due to political factors influencing the allocation of contracts (Nakamura and Steinsson, 2014).

We deal with the potential bias by resorting to a Bartik (1991) instrument: The change in national DoD spending interacted with the MSA's average share of DoD spending to local GDP over the sample period. This instrument identifies the effect of spending on house prices by relating changes in the MSAs' DoD spending to their persistent and differential exposure to changes in national military spending. That is, when the federal government expands military spending, some MSAs tend to receive more DoD contracts than others, because they are systematically more exposed to changes in military spending. This systematic component of changes in local DoD spending is isolated by the instrument. In this respect, Figure A.4 in Appendix A.2 plots the period-by-period first-stage Kleibergen-Paap F-statistics. The F-statistics are in the range 26-63, which is above the cluster-robust threshold for weak instruments of 23.1 provided by Montiel Olea and Pflueger (2013), confirming the relevance of the instrument.

The identifying assumption behind this approach is that, conditional on controls, there are

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

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

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

no confounding factors affecting local house price growth—not just contemporaneously, but also at different leads and lags—that are correlated with the MSAs' exposure to changes in military spending over the cross section, as well as with changes in national military spending in the time-series dimension:[9]

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

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

The average DoD spending share in a given MSA is likely to be an equilibrium object determined by factors such as industry composition, or by having a military base nearby. For this reason, it is worth stressing that the exogeneity condition is formulated in terms of *changes* in the outcome, rather than *levels*. Even if the level of local house prices was codetermined with the local DoD spending share, equation (2.2) would still hold. However, a potential concern is that the MSAs' exposure to military spending is related to their exposure to the national business cycle—for instance through different industry or housing market composition—which drives the differential house price response across MSAs through correlation between national DoD spending and the business cycle. More formally, this would imply that the exogeneity condition (2.2) does not hold, because the error term $\varepsilon_{i,t+h+j}$ carries a $\gamma_i \xi_{t+h+j}$ structure, where γ_i is correlated with $\bar{G}_i$ over the cross section, and ξ_{t+h+j} is correlated with $\frac{G^{nat}_{t+1} - G^{nat}_t}{Y_{i,t}}$ in the time-series dimension. We address some of these concerns in Appendix A.2.5, through a number of robustness checks, as summarized in the next subsection.

An additional identifying assumption is necessary because of the dynamic effects of government spending: Lagged and leading shocks to local government spending should also be unrelated to the contemporaneous instrument (Stock and Watson, 2018). This assumption does not hold, since the instrument itself is autocorrelated. In that case, the estimate of β_h will not only pick up the contemporaneous effect from the shock, but also the effects from past shocks, so that β_h cannot be interpreted as the effect of an unanticipated shock to DoD spending. For this reason, we follow Ramey and Zubairy (2018) by including two lags of the instrument, $\bar{G}_i \frac{G^{nat}_{t+1} - G^{nat}_t}{Y_{i,t}}$, as well as two lags of the one-year normalized change in local spending, $\frac{G_{i,t+1} - G_{i,t}}{Y_{i,t}}$. Table A.4 illustrates that conditioning on autoregressive dynamics in the instrument and local spending leads to a smaller estimate of the house price response, thus capturing past changes in shocks to local spending. Finally, we add two lags of the one-year growth in the dependent variable to pick

[9] 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).

up momentum in house prices. As we show in Appendix A.2.3, house prices are slightly— yet, significantly—lower in the two years prior to a spending change, which could reflect anticipation effects.[10] However, controlling for this pre-trend—as we do in our baseline regression—makes little difference to our results besides reducing standard errors.

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

2.2.1 The response of house prices to a fiscal spending shock

We estimate h separate (2.1) regressions, for $h = 1, 2, \dots, 13$ horizons, and report the estimates in Figure 2. The OLS and IV estimates are shown in panels (a) and (b), respectively. Standard errors are heteroskedasticity-robust and clustered at the MSA level, so as to account for within-MSA correlation of the error term. Clustering by MSA allows for non-parametric time series dependence in the errors by taking advantage of the cross-sectional dimension of the data.[11] Grey areas indicate 95 percent confidence bands based on the point estimate standard errors.

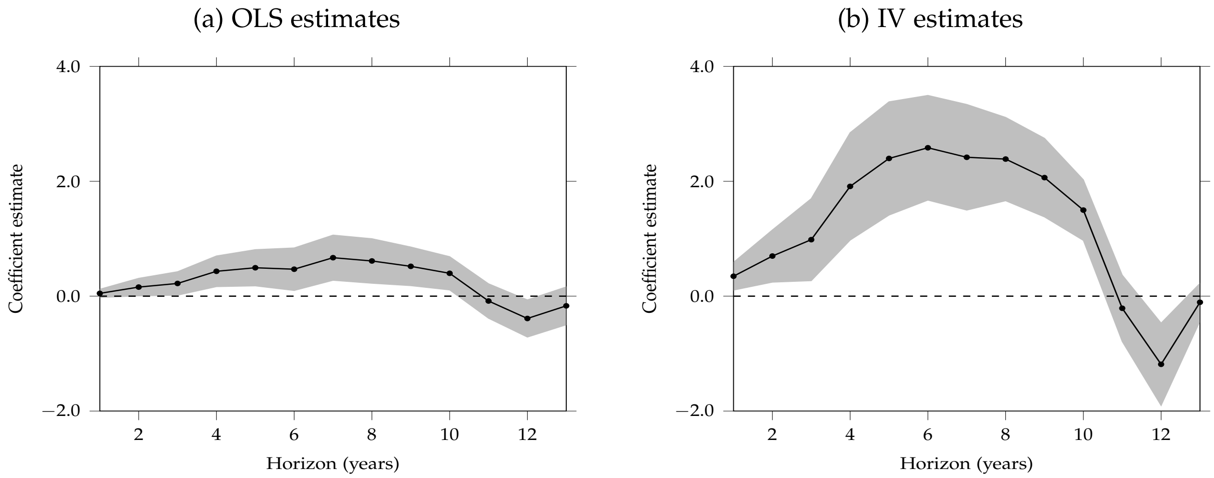
According to the IV estimates, the response of house prices to an expansion in government spending follows a hump-shaped pattern, peaking after six years, thus reverting back to trend. In terms of magnitude, we observe a relative increase in house prices of 2.6 percent after six years, as a result of an increase in spending of 1 percent of GDP in the first year. The OLS estimates also follow a hump-shaped pattern, but are biased toward zero. The difference between the IV and OLS estimates is sizeable, with non-overlapping confidence bands until the 11-year horizon. Formally, period-by-period endogeneity tests firmly reject the null of the IV estimate being equal to the OLS estimate until $h = 11$, as shown in Figure A.6 in Appendix A.2.

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

[10] As Auerbach et al. (2024) argue, the instrument does not filter out all anticipation effects since anticipated changes in national military spending are not removed by the instrument.

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

Figure 2: Regional house price responses to military spending



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

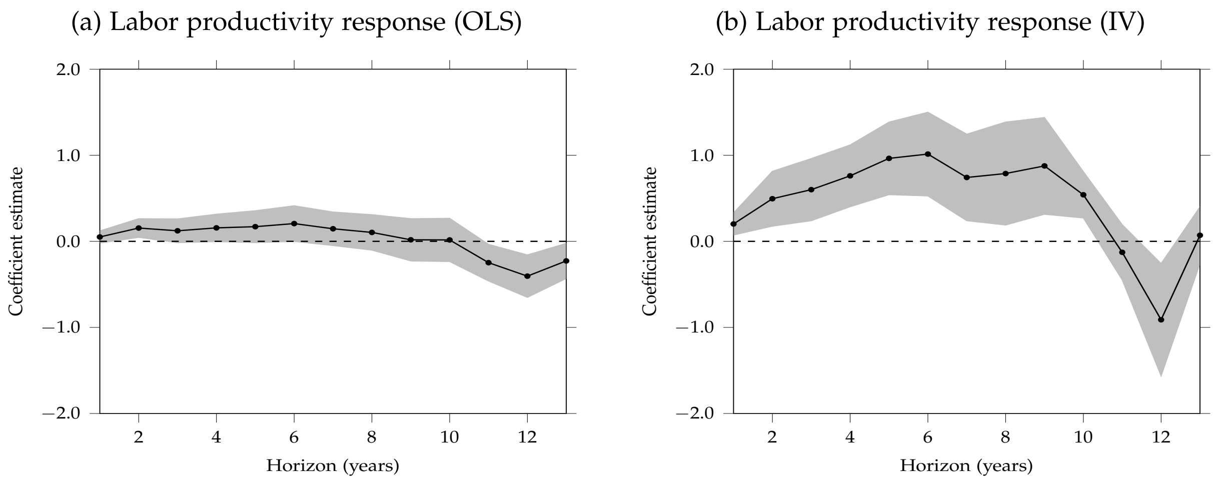
our results to outliers and different sets of controls in the baseline regression. There are some key takeaways from the robustness exercises, which support a causal interpretation of the IV estimates. We show that our results are robust to controlling for three well-known measures of house-price cycle exposure interacted with year dummies.[12] This addresses the concern that our results could reflect a differential exposure to the 2000s boom-bust cycle in U.S. house prices. Moreover, differences in industry composition do not drive our results, since controlling for two-digit industry employment shares multiplied by year dummies does not alter the estimates.

2.2.2 The conditional behavior of productivity

The procyclical behavior of productivity represents the common trait of the mechanisms we will consider to frame the expansionary effect of fiscal spending on house prices. To validate the role of this essential feature, we return to regression (2.1), this time using (the growth rate of) local labor productivity as the dependent variable. Labor productivity is measured as local GDP divided by local employment using data from the BEA. Figure 3 presents the estimates of the productivity response to a change in DoD spending. The IV estimates are clearly in line with our model's predictions and the BVAR evidence presented above, as well as the evidence reported by Auerbach et al. (2024). Figure 4 conveys further insights into the empirical connection between labor productivity and house prices. We report the scatter plots of the IV estimates of the impulse-responses (at a six-year horizon), based on 1,000 iterations from a cluster bootstrap drawing

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

Figure 3: The regional response of labor productivity to military spending

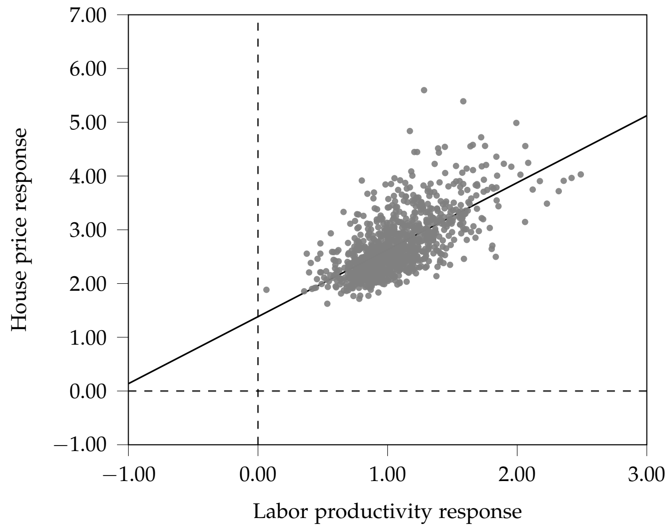


Notes: The figure shows the estimates of β_h from regression (2.1) based on an annual panel of 380 MSAs covering the period 2001-2019, with labor productivity as the dependent variable. The OLS and IV estimates are plotted in panels (a) and (b), respectively. The regressions include as controls two lags of the one-year growth in labor productivity, 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.

sets of MSAs from our sample. There is a clear positive relationship across MSAs between the magnitudes of the responses of house prices and labor productivity to a change in DoD spending. That is, for the sets of MSAs where the estimated house price response is larger, so is the response of labor productivity.

We end this section with a brief comment on the conditional response of business formation in the regional data. In Appendix A.2.6, we document an increase in the number of establishments in response to an increase in local government spending. We also find that, for the sets of MSAs where the estimated responses of the house price and labor productivity are larger, so is the response of establishments. The link between these three variables is relatively tight, with pairwise correlation coefficients ranging between 0.7 and 0.8.

Figure 4: The joint distribution of house price and labor productivity responses



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

3 Framing the problem

In standard neoclassical economies populated by Ricardian households, 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 induced by the 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 response of the price of housing, in virtually any model where Ricardian agents benefit from housing services and, *de facto*, price the asset. To see why, consider a standard Ricardian household's Euler equation for housing, which may be solved forward to yield an expression for their shadow value of housing:

$$q_t \lambda_t = \Upsilon H_t^{-\sigma_h} + \beta E_t \left\{ \lambda_{t+1} q_{t+1} \right\} = \Upsilon \mathrm{E}_t \left\{ \sum_{i=0}^{\infty} \beta^i H_{t+i}^{-\sigma_h} \right\} \equiv \Lambda_t, \tag{3.1}$$

where q_t, λ_t, and H_t denote, respectively, the price of housing (relative to that of the numeraire, typically consumption goods), households' shadow value of income, and their stock of housing, while σ_h, β, and Υ measure, respectively, the curvature of CRRA utility with respect to housing, households' discount factor, and the weight of housing in household utility. Since housing does not depreciate, or depreciates very slowly, H_t is effectively an "idealized durable" according to Barsky et al. (2007): This means that the intertemporal elasticity of substitution in housing demand is close to infinite. As a result, short-term movements in H_t—such as those generated by a temporary shock to fiscal spending—will affect the right-hand side of (3.1) relatively little, given that β is close to one. Hence, it is possible to approximate

$$q_t \lambda_t = \Lambda_t \approx \Lambda. \tag{3.2}$$

Accordingly, movements in the price of housing are forced to mirror changes in households' shadow value of income.[13] In light of this, any model where a Ricardian household "dictates" the price in the housing market 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. 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—represents the only viable option.[14]

As we detail in the next subsection, our point consists of considering supply-side mechanisms that, through a suitable connection between government spending and TFP, generate a sizeable increase in the real wage in the face of a fiscal expansion, thus expanding the shadow value of income of the households that ultimately price housing.

3.1 A tractable model with credit frictions

Supported by our empirical evidence, we posit that a positive response of house prices to an expansion in government spending hinges on the conditional expansion of productivity. We will now present a parsimonious framework that illustrates this point. The critical element of the model lies in the strength of the productivity increase that is prompted by an expansion in government spending. If this elasticity is sufficiently high, house prices will increase. To make the analysis as general as possible, we will contemplate a baseline scenario where households are subject to financial frictions, a situation that nests the Ricardian benchmark discussed above as a special case.

To preserve tractability, we consider a small open economy with free capital mobility, although the main insights do not depend on this. Time is discrete, $t = 1, 2, ..., \infty$. The economy is inhabited by a continuum of homogeneous households of size one. The same assumption extends to the population of firms, which operate competitively to produce a perishable consumption good by hiring workers, whose labor hours are remunerated at the rate w_t. Households' intertemporal utility is additively separable in all its components, which are: two logarithmic terms—one defined over the consumption of perishable goods (C_t, assumed to be the numerarire), and one over

[13]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).

[14]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. For instance, this is the case when imposing complementarity between private and public spending, as in Bouakez and Rebei (2007).

services emanating from the holdings of housing (available in unit supply, so that we preserve the property of housing as an idealized durable)—as well as a disutility term from supplying labor hours (N_t). Households can borrow internationally at the constant gross real interest rate $R > 1$, subject to the collateral constraint $B_t \leq mq_tH_t$, where B_t denotes the real stock of debt, and $m > 0$ is the loan-to-value ratio. We assume that households are less patient than their foreign counterparts. Hence, their discount factor, $0 < \beta < 1$, satisfies $R\beta < 1$. Firms utilize a production technology that is linear in labor: $Y_t = A_tN_t$, where A_t denotes Total Factor Productivity (TFP). Henceforth, we report the key equations from the model in log-linear form, where lowercase-hat variables denote log-deviations from steady state.

Households' maximization of intertemporal utility gives rise to the following equations:

$$\hat{w}_t = \psi\hat{N}_t + \hat{C}_t, \tag{3.3}$$

$$\hat{C}_t = \beta R E_t\left\{\hat{C}_{t+1}\right\} - (1-\beta R)\hat{\mu}_t, \tag{3.4}$$

$$\hat{q}_t - \hat{C}_t = m(1-\beta R)(\hat{q}_t + \hat{\mu}_t) + \beta E_t\left\{\hat{q}_{t+1} - \hat{C}_{t+1}\right\}. \tag{3.5}$$

These conditions describe, respectively, households' labor supply schedule (where $\psi \geq 0$ denotes the inverse Frisch elasticity of labor supply), the Euler equation governing households' intertemporal consumption-saving decisions, and the Euler equation associated with the consumption of housing services (where $\hat{\mu}_t$ denotes the shadow value of borrowing, in percentage deviations from its steady-state value).

In line with the distinctive property of the structural frameworks that we will employ in our quantitative analysis, we assume that productivity is endogenous, and responds positively to an increase in government spending, $\hat{G}_t$. We capture this through a reduced-form relationship: $\hat{A}_t = \alpha\hat{G}_t$, where $\alpha > 0$. Therefore, by virtue of firms' profit maximization, the labor-demand schedule can be written as $\hat{w}_t = \alpha\hat{G}_t$. We further assume $\hat{G}_t = \rho\hat{G}_{t-1} + \varepsilon_{g,t}$, where $\epsilon_{g,t} \sim N\left(0, \sigma_g^2\right)$ and $\rho \in [0,1)$. The model is closed by goods market clearing—$\hat{Y}_t = (1-\theta)\hat{C}_t + \theta\hat{G}_t$, where $\theta \in [0,1]$ denotes the ratio of government spending to total output—and the government's budget constraint, $\hat{G}_t = \hat{\tau}_t$, where $\hat{\tau}_t$ denotes the revenue from lump-sum taxes.

Notice that, in the absence of a binding collateral constraint, the frictionless counterpart of (3.5) can be solved forward to obtain $\hat{q}_t = \hat{C}_t, \forall t$, so that the analysis above applies. In the baseline case where the borrowing constraint is binding, we may combine (3.4) and (3.5) to cancel out $\hat{\mu}_t$, obtaining (after setting $m = 1$, without loss of generality):

$$\hat{q}_t = \frac{1}{R}E_t\left\{\hat{q}_{t+1}\right\} + \frac{R-1}{R}E_t\left\{\hat{C}_{t+1}\right\}. \tag{3.6}$$

This condition implies that, as long as consumption is a function of the fiscal shock, the house-price response will be proportional to the response of consumption. The only difference from the frictionless case lies in the emergence of a scaling factor. In fact, by conjecturing $\hat{C}_t = b\hat{G}_t$, we can readily show that:

$$\hat{q}_t = \frac{R-1}{R-\rho}\rho \underbrace{b\hat{G}_t}_{=\hat{C}_t}, \tag{3.7}$$

where $\frac{R-1}{R-\rho}\rho < 1$ represents the above-mentioned scaling factor. From (3.7), it is easy to infer that obtaining a procyclical response of the house price to a fiscal shock amounts to ensuring that $b > 0$. This is remarkable, as it shows that, no matter whether housing is priced by financially constrained or Ricardian agents, crowding-in of house prices is ultimately a factor that depends on the impact of fiscal spending on these agents' nondurable consumption, while comovement between the latter and house prices remains unaffected, from a qualitative viewpoint. The intuition behind this result is that the household, while financially constrained in the access to credit, is effectively using housing as a vehicle to smooth consumption. When faced with a fiscal shock, financially constrained households experience a similar drop in lifetime after-tax income as purely Ricardian households, and therefore display a qualitatively similar response.

The final step consists of verifying our guess, combining it with the aggregate resource constraint, the production technology, and the labor-market relationships, to obtain:

$$\hat{C}_t = \underbrace{\frac{\alpha(1+\psi) - \psi\theta}{1+(1-\theta)\psi}}_{=b} \hat{G}_t, \tag{3.8}$$

so that we need $\alpha > \frac{\psi\theta}{1+\psi}$ to ensure consumption crowding-in and, ultimately, a procyclical response of house prices. In other words, TFP needs to be sufficiently responsive to the fiscal shock. In contrast, $\alpha = 0$ implies that $b < 0$, showing that in the absence of an endogenous response of TFP, consumption and house prices will decline. The larger the elasticity of labor supply (and the lower the share of government spending to total output), the more likely it is that a given shock will induce consumption and house prices to increase after a shock to government spending. Despite the introduction of physical capital, this feature is in place in the quantitative frameworks we analyze in the next sections.

4 Baseline quantitative framework

This section first lays down the common ground for the endogenous-productivity mechanisms to operate in the transmission of shocks to fiscal spending. To this end, we envisage a real business cycle economy populated by two types of households, differentiated by their discount factors: Impatient households have a lower discount factor than patient households, and can borrow up to a share of the present value of their housing stock. This implies that patient households act as lenders. Both household types work, consume nondurables, and accumulate housing. Patient households also accumulate capital that is rented to firms. The motivation for including impatient households is twofold: First, this allows us to account for movements in mortgage debt alongside those in house prices, as these two variables are closely related in the data. Second, it allows us to analyze whether the dynamics of the house price is affected by which type of agent is determining the house price in equilibrium.

4.1 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_t^j, where $j \in \{b, l\}$ indexes impatient and patient household-specific variables, respectively. Each type of household maximizes the following life-time utility function:

$$E_0 \left\{ \sum_{t=0}^{\infty} \left(\beta^j\right)^t \left[\frac{\left(C_t^j - h^j C_{t-1}^j\right)^{1-\sigma_c}}{1-\sigma_c} + \Upsilon^j \frac{\left(H_t^j\right)^{1-\sigma_h}}{1-\sigma_h} - \Psi^j \frac{\left(N_t^j\right)^{1+\psi}}{1+\psi} \right] \right\}, \tag{4.1}$$

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

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

$$C_t^b + q_t H_t^b + R_{t-1} B_{t-1}^b = w_t^b N_t^b + B_t^b + q_t H_{t-1}^b - \tau_t^b, \tag{4.2}$$

$$B_t^b \leq \gamma B_{t-1}^b + (1-\gamma) m \frac{E_t \left\{ q_{t+1} H_t^b \right\}}{R_t}, \tag{4.3}$$

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

The borrowing constraint in equation (4.3) states that impatient households can borrow up to a fraction $m \in [0, 1]$ of the present value of their housing stock at the beginning of the next period, as in Kiyotaki and Moore (1997). Following Guerrieri and Iacoviello (2017), we allow for inertia in the dynamics of mortgage debt, as measured by $\gamma \in [0, 1)$, to account for the slow-moving nature of the stock of debt in the data. We assume that shocks to the economy are sufficiently small that the borrowing constraint invariantly holds with equality in the neighborhood of the steady state.

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

$$C_t^l + q_t H_t^l + I_t + B_t^l + B_t^g = w_t^l N_t^l + q_t H_{t-1}^l + R_{t-1} B_{t-1}^l + R_{t-1} B_{t-1}^g + r_t^k K_{t-1} - \tau_t^l, \tag{4.4}$$

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:

$$K_t = K_{t-1}(1-\delta) + I_t(1-\Phi_t), \tag{4.5}$$

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

4.2 Production

Four different production settings will be considered. As will become clear, each of these implies that aggregate production can be written as:

$$Y_t = TFP_t K_{t-1}^{\mu} \left[\left(N_t^b\right)^{\alpha} \left(N_t^l\right)^{1-\alpha}\right]^{1-\mu}, \qquad \alpha, \mu \in (0,1), \tag{4.6}$$

where the appropriate expression for TFP_t will be different across models, in each case depending endogenously on some model-specific variables that will ultimately be affected by shocks to fiscal spending.[15]

4.3 Fiscal policy

Government spending follows an autoregressive process (in logs):

$$\log G_t = (1-\gamma_g)\log \bar{G} + \gamma_g \log G_{t-1} + \epsilon_{g,t}, \qquad \epsilon_{g,t} \sim N\left(0, \sigma_g^2\right). \tag{4.7}$$

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:

$$\tau_t^b = \alpha \tau_t^{TOT}, \tag{4.8}$$

$$\tau_t^l = (1-\alpha)\tau_t^{TOT}. \tag{4.9}$$

[15]Our definition of TFP as the share of output not accounted for by capital and labor inputs (i.e., the Solow residual) follows conventional definitions in the literature. From such a standpoint, all the model variations we will be considering are consistent with standard measures of TFP, such as the one computed by Fernald (2014), with the difference that fluctuations in TFP will be due to some endogenous propagation channels.

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

$$R_{t-1}B^g_{t-1} + G_t = \tau^{TOT}_t + B^g_t. \tag{4.10}$$

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:

$$\tau^{TOT}_t = \left(\tau^{TOT}_{t-1}\right)^{\rho_\tau} \left(\frac{B^g_{t-1}}{Y_{t-1}}\right)^{(1-\rho_\tau)\gamma_\tau}, \tag{4.11}$$

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.

4.4 Market clearing

The market clearing conditions are:

$$Y_t = C_t + G_t + I_t, \tag{4.12}$$

$$C_t = C^b_t + C^l_t, \tag{4.13}$$

$$H = H^l_t + H^b_t, \tag{4.14}$$

where H is a fixed stock of housing in the economy. Lastly, the mortgage market clears when patient households' lending equals impatient households' borrowing:

$$B^l_t = B^b_t. \tag{4.15}$$

5 Propagation channels

We now turn to a set of supply-side mechanisms that allow shocks to government spending to have an endogenous effect on TFP. We take the first three of these mechanisms off the shelf from existing literature. Since they have been studied in previous work, we keep their description short and concise. In each case, we relegate the first-order conditions for optimal choices of factor inputs to Appendix C. We devote more attention to the fourth mechanism—which combines love of variety with endogenous firm turnover—since it draws on various contributions and is therefore less studied in the existing business-cycle literature, and since it entails a more elaborate production structure.

5.1 Productive government spending (PGS)

We first consider the case in which government spending has productivity-augmenting effects, as considered by Aschauer (1989), Baxter and King (1993), and several subsequent studies. In this case, the production function takes the following form:

$$Y_t = \left(K^g_{t-1}\right)^{\mu_g} K^{\mu}_{t-1} \left[\left(N^b_t\right)^{\alpha} \left(N^l_t\right)^{1-\alpha} \right]^{1-\mu}, \tag{5.1}$$

where K^g_t denotes the stock of public capital, and $\mu_g \geq 0$ measures its productivity, so that $TFP_t = \left(K^g_{t-1}\right)^{\mu_g}$. We maintain the assumption of constant returns to scale over private-sector inputs, as in Baxter and King (1993), giving rise to increasing returns to scale when considering all input factors. Public capital accummulates according to a standard law of motion:

$$K^g_t = \left(1 - \delta_g\right) K^g_{t-1} + I^g_t, \tag{5.2}$$

where $\delta_g \in [0, 1]$ is the depreciation rate of public capital, and I^g_t denotes public investment, which we equate to government spending, $I^g_t = G_t$. In practice, this formulation of the problem is likely to be on the optimistic side, as we are implicitly assuming that *all* public spending is productive, and that the accummulation of public capital is subject neither to investment adjustment costs, nor to time-to-build delays. In this model version, the two key parameters are the productivity and the depreciation rate of public capital, μ_g and δ_g.

5.2 Learning by doing (LBD)

The next mechanism we consider is learning by doing, as in Chang et al. (2002) and d'Alessandro et al. (2019). According to this theory, labor productivity is an increasing function of past hours worked, reflecting that workers learn on-the-job. Specifically, the production function is now given by:

$$Y_t = K^{\mu}_{t-1} \left[\left(X^b_t N^b_t\right)^{\alpha} \left(X^l_t N^l_t\right)^{1-\alpha} \right]^{1-\mu}, \tag{5.3}$$

where X^b_t and X^l_t denote the accummulated skills of borrowers and lenders, respectively. In turn, each of these evolves according to

$$X^j_t = \left(X^j_{t-1}\right)^{\rho_x} \left(N^j_{t-1}\right)^{\theta_n}, \qquad j \in \{b, l\}, \tag{5.4}$$

where $\theta_n \geq 0$ measures the sensitivity of accummulated skills with respect to past hours worked, and $\rho_x \in [0, 1)$ captures the persistence of these skills. This formalizes the assumption that workers accummulate skills through (past) employment. We assume that this process does not depend on whether households are borrowers or lenders, i.e., that the two key parameters, θ_n

and ρ_x, are the same for both types of household. In this setting, $TFP_t = \left[\left(X_t^b\right)^{\alpha}\left(X_t^l\right)^{1-\alpha}\right]^{1-\mu}$.

5.3 Variable technology utilization (VTU)

As a further alternative, we allow for endogenous variations in the rate at which firms utilize the available level of technology in the economy, following Bianchi et al. (2019) and Jørgensen and Ravn (2022). That is, while the aggregate technology level remains exogenous (and fixed, in our context), firms can now choose the technology utilization rate optimally, capturing the fact that firms may not always choose to (or be able to) adopt new inventions and technologies into their specific production setup as soon as these become available. The production function then becomes:

$$Y_t = u_t A_Z K_{t-1}^{\mu}\left[\left(N_t^b\right)^{\alpha}\left(N_t^l\right)^{1-\alpha}\right]^{1-\mu}, \tag{5.5}$$

where A_Z is the exogenous technology level, which we normalize to 1, and u_t denotes the technology utilization rate, which also corresponds to TFP_t, in this setting. To make it costly for firms to adopt new technologies as they arrive, we introduce a convex cost of changing the utilization rate, denoted by $z(u_t)$. Specifically, this takes the following form:

$$z(u_t) = \chi_1(u_t - u) + \frac{\chi_2}{2}(u_t - u)^2, \tag{5.6}$$

with $\chi_1, \chi_2 > 0$. Note that this function satisfies $z(u) = 0$—i.e., adjustment costs are zero in steady state, where the utilization rate equals 1—as well as $z''(\cdot) = \chi_2$, ensuring that the function is convex. The same functional form was used by Jørgensen and Ravn (2022), as well as by Christiano et al. (2005) in the context of variable capital utilization.[16] The key parameter is the curvature of the adjustment cost function associated with changes in the utilization rate of technology, as given by $\frac{\chi_2}{\chi_1}$.

5.4 Love of variety and endogenous firm turnover (LOV)

We finally turn to the model with love of variety and endogenous entry and exit of firms. To this end, we assume that the production of nondurables and investment goods occurs in a two-layer production sector, in the vein of Rotemberg and Woodford (1992), Jaimovich (2007), and Jaimovich and Floetotto (2008), among others. 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

[16] As discussed by Jørgensen and Ravn (2022), whereas variable technology utilization may give rise to increasing returns, variable utilization of capital (or labor) would not. The reason is that firms would pay rental rates on *utilized* capital or labor, implying that an increase in the utilization rate of either of these would entail an increase in firms' marginal costs, in addition to the cost of changing the utilization rate. In contrast, technology itself is free to use.

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. The first-order conditions of the firms are presented in Appendix C.4.

5.4.1 Final good production

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

$$Y_t = \left[\int_0^1 Q_t(j)^{\omega} dj\right]^{\frac{1}{\omega}}, \qquad \omega \in (0,1). \tag{5.7}$$

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):

$$Q_t(j) = F_t(j)^{\tau+\frac{\rho-1}{\rho}} \left[\sum_{i=1}^{F_t(j)} m_t(j,i)^{\rho}\right]^{\frac{1}{\rho}}, \qquad \rho \in (0,1), \tag{5.8}$$

where $m_t(j,i)$ is the output of firm i in sector j. The production function in equation (5.8) is a generalization of the Dixit and Stiglitz (1977) CES aggregation function that disentangles the variety effect from the elasticity of substitution across inputs, $1/(1-\rho)$.[17] The variety effect is measured by $\tau \geq 0$, and implies that, as the number of intermediate firms within a sector increases, the sectoral aggregate good expands for a given input of intermediate goods. If $\tau = (1-\rho)/\rho$, the function reduces to the Dixit-Stiglitz function, in which the variety effect is tied to the elasticity of substitution, while $\tau = 0$ represents the case of no love for variety.[18]

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

$$Q_t(j) = \left(\frac{p_t(j)}{P_t}\right)^{\frac{1}{\omega-1}} Y_t, \tag{5.9}$$

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 mini-

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

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

mization problem, and is given by

$$m_t(j,i) = \left(\frac{p_t(j,i)}{p_t(j)}\right)^{\frac{1}{\rho-1}} \left(\frac{p_t(j)}{P_t}\right)^{\frac{1}{\omega-1}} \frac{Y_t}{\left(F_t(j)^{\tau+\frac{\rho-1}{\rho}}\right)^{\frac{\rho}{\rho-1}}}, \tag{5.10}$$

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

$$p_t(j) = \frac{1}{F_t(j)^{\tau+\frac{\rho-1}{\rho}}} \left[\sum_{i=1}^{F_t(j)} p_t(j,i)^{\frac{\rho}{\rho-1}}\right]^{\frac{\rho-1}{\rho}}. \tag{5.11}$$

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

5.4.2 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:

$$m_t(j,i) = k_{t-1}(j,i)^{\mu} \left[\left(n_t^b(j,i)\right)^{\alpha} \left(n_t^l(j,i)\right)^{1-\alpha}\right]^{1-\mu} - \varphi, \tag{5.12}$$

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

Firms sell their intermediate goods to the final good firm in a monopolistically competitive market within each sector. In doing so, they account for the effect they exert on the sectoral price index, $p_t(j)$, but not on the final good price, P_t, as in Jaimovich (2007). Thus, the elasticity of demand for the intermediate goods firm, according to the demand curve (5.10) and the price index (5.11), is

$$\varepsilon_{m_t(j,i)} = \frac{1}{\rho-1} + \left(\frac{1}{\omega-1} - \frac{1}{\rho-1}\right) \left(\frac{p_t(j,i)}{p_t(j)F_t(j)^{\tau}}\right)^{\frac{\rho}{\rho-1}} \frac{1}{F_t(j)}. \tag{5.13}$$

We assume that the elasticity of substitution within sectors is higher than the elasticity of substitution across sectors, $\frac{1}{1-\omega} < \frac{1}{1-\rho}$.[19] This implies that if an individual firm increases its price, $p_t(j,i)$, relative to the sectoral price index adjusted for the variety effect, $p_t(j)F_t(j)^{\tau}$, the elasticity of demand increases, since the demand for the aggregate sectoral good falls through the firm's effect on the sectoral price index.

The elasticity of demand in equation (5.13) results in firms setting prices at the following

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

markup over the marginal cost:

$$x_t(j,i) = \frac{\varepsilon_{m_t(j,i)}}{1+\varepsilon_{m_t(j,i)}}, \tag{5.14}$$

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

Firms' cost minimization results in the following cost function:

$$C_t(j,i) = A\left(r_t^k\right)^{\mu}\left(w_t^b\right)^{\alpha(1-\mu)}\left(w_t^l\right)^{(1-\alpha)(1-\mu)}\left(m_t(j,i)+\varphi\right), \tag{5.15}$$

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:

$$\frac{p_t(j,i)}{P_t} m_t(j,i) = C_t(j,i). \tag{5.16}$$

Finally, combining the free entry condition with the cost function in equation (5.15) 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:

$$m_t(j,i)\left(x_t(j,i)-1\right) = \varphi. \tag{5.17}$$

5.4.3 Symmetric firm equilibrium

Imposing symmetry and combining the aggregate price index with the sectoral price index now allows us to express the price of an intermediate good relative to that of the final good as a function of the number of firms, whenever $\tau > 0$:

$$\frac{p_t}{P_t} = F_t^{\tau}. \tag{5.18}$$

Moreover, we obtain from (5.8):

$$Y_t = F_t^{1+\tau} m_t. \tag{5.19}$$

Equations (5.18) and (5.19) yield two key insights about the *variety effect* and its interplay with the *competition effect*. First, p_t/P_t increases in the number of firms, for $\tau > 0$. Increasing the taste for variety lowers the marginal cost of the final good firm, thereby lowering the price of the final good relative to that of the intermediate goods. Second, a larger number of intermediate firms increases final output more than one-for-one, for given intermediate firm-level production. Thus, there are increasing returns to the number of firms, while the production technology at the intermediate-firm level features constant returns to scale.

Analogous considerations about the role of the two effects at play in the model can be made with respect to their impact on TFP. To this end, we may combine (5.12), (5.14), (5.17), and (5.19), together with market clearing in the factor market, to write aggregate output as in (4.6). However, now

$$TFP_t = F_t^{\tau}/x_t, \tag{5.20}$$

implying that the entry and exit of firms results in endogenous procyclical TFP variations through the competition and the variety effects. As emphasized by Jaimovich and Floetotto (2008), the competition effect stimulates TFP through the impact that changes in the number of firms exert on the markup. To see this, consider an increase in the number of firms fostered by a fiscal expansion, which lowers the markup through more intense competition. In turn, this induces firms to increase production to cover their fixed costs, thus driving up TFP. 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.

To illustrate how the variety and competition effects formally combine, we consider a simplified version of the model in Appendix D, which we solve analytically. This facilitates an analytical characterization of the condition under which the competition effect alone is capable of generating an increase in house prices following a fiscal expansion. In line with the insights of Devereux et al. (1996) and Bilbiie (2011), this is only possible if the steady-state markup and/or the Frisch elasticity of labor supply are set at implausibly high values. We then proceed to show that, once both channels are active, the model can produce an increase in consumption and the house price if the value of τ is above a given threshold.

6 Calibration and estimation

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

6.1 Calibration

The vector $\omega_1 = \{\alpha, \beta^b, \beta^l, \delta, \mu, m, \theta, \Xi\}$ contains the parameters—common across models—that we choose to calibrate. We set the borrowers' share of labor income to $\alpha = 0.21$, in line with the estimate of Iacoviello and Neri (2010). The discount factors of borrowers and lenders are set to $\beta^b = 0.97$ and $\beta^l = 0.99$, respectively, as in Jensen et al. (2018). The depreciation rate of capital is set at $\delta = 0.025$, while the income share of capital is set to $\mu = 0.25$. These values imply ratios of investment to output and of capital to output of 0.18 and 1.8, respectively, both of which are broadly in line with the corresponding average values for the U.S. economy. We set the loan-to-value ratio m to 0.85, as in Iacoviello and Neri (2010). The share of government spending to output, denoted by θ, is set to 0.24, while the ratio of public debt to output, denoted by Ξ, is set to 0.7. Both of these numbers are closely in line with the average values for the U.S. over the past decades. We collect the calibrated parameters in Table 1.[20]

We then turn to the productivity-augmenting mechanisms under consideration. The only parameters we choose to calibrate are those governing the elasticity of substitution within and across sectors, ρ and ω, which are specific to the model with endogenous entry and love of variety. We set $\rho = 0.9$ and $\omega = 0.75$, in order to obtain elasticities of substitution of 10 (within sectors) and 4 (across sectors), respectively. The latter value is in line with Bilbiie et al. (2019), who use an elasticity of substitution of 3.8, while the former is chosen to reflect that varieties are increasingly substitutable as product categories are disaggregated, as found by Broda and Weinstein (2006), who estimate elasticities of substitution ranging from 1.2 to 17. Note that the values of ρ and ω encompass the common practice in the New Keynesian literature of setting the elasticity of substitution in one-sector models to 6; see, e.g., Rotemberg and Woodford (1992).

6.2 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—for each of the four different production structures outlined above—to the empirical responses from our BVAR in Section 2.1. We collect in $\omega_2 = \left\{\sigma_c, \sigma_h, h^l, h^b, \psi, \phi, \gamma, \gamma_\tau, \rho_\tau, \gamma_G, \sigma_g, \Sigma^k\right\}$ the parameters to be estimated. Here, Σ^k denotes the parameters specific to each of the supply-side mechanisms under consideration, which are given by: $\Sigma^{LOV} = \{\tau, x\}$, $\Sigma^{PGS} = \left\{\mu_g, \delta_g\right\}$, $\Sigma^{LBD} = \{\theta_n, \rho_x\}$, and $\Sigma^{VTU} = \{\chi_2\}$.

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:

$$\widehat{\omega}_2 = \arg\min_{\omega_2} \left(\Gamma(\omega_2) - \widehat{\Gamma}\right)' W \left(\Gamma(\omega_2) - \widehat{\Gamma}\right). \tag{6.1}$$

The weighting matrix W is diagonal, with the inverse of the sample variances of the BVAR-based impulse responses as entries. Effectively, this means that we are attaching higher weights to those impulse responses that are estimated most precisely. For each model, we match the impulse responses of all variables in our BVAR (except the forecast error series, FE_t) for the first

[20] 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.

Table 1: Calibrated parameter values

Parameter	Description	Value
Panel A: Calibrated parameters of the common block		
α	Income share of impatient households	0.21
β^b	Discount factor, borrowers	0.97
β^l	Discount factor, lenders	0.99
δ	Capital depreciation rate	0.025
μ	Capital share of production	0.25
m	Loan-to-value ratio of borrowers	0.85
θ	Ratio of government spending to output	0.24
Ξ	Ratio of government debt to output	0.7
Panel B: Calibrated model-specific parameters		
ρ	Substitution parameter within sectors (LOV model)	0.9
ω	Substitution parameter across sectors (LOV model)	0.75

25 quarters after the shock.[21] To obtain credible sets for the parameter estimates, we employ the entire distribution of impulse responses used to construct the credible sets in Figure 1; that is, we run 1,000 estimations of each DSGE model.

6.3 Estimation results

We report the median parameter estimates in Table 2, as well as the associated credible sets based on the 16th and 84th percentiles of the distribution of the estimates.[22] Regarding the set of parameters common across models, we note that these generally take on values that are in line with the existing literature. The parameters of households' utility function display limited variation across models, with the exception of the PGS model, which features no habit formation for borrowers and a very low Frisch elasticity of labor supply.[23] The degree of inertia in mortgage debt is typically close to the estimate of Guerrieri and Iacoviello (2017) of 0.7. The estimated capital adjustment costs are somewhat different, though this parameter displays substantial variation in the existing literature. Finally, the parameters governing the tax rule and the process of government spending itself are relatively similar in all cases.

We then turn to the parameters governing the strength of the productivity-augmenting propa-

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

[23] In the estimation, we impose a lower bound of 0.1 on the Frisch elasticity, and an upper bound of 4. While the latter is well above microeconometric estimates, it allows traditional RBC models to match business-cycle data (see the discussion by Chetty et al., 2011).

Table 2: Estimated parameter values

Parameter	Description	LOV model	PGS model	LBD model	VTU model
σ_c	Curvature in utility of C	1.210 [0.711−2.741]	2.335 [1.769−2.666]	1.331 [1.102−1.523]	1.383 [1.230−3.407]
σ_h	Curvature in utility of H	0.293 [0.105−1.202]	0.100 [0.100−0.581]	0.101 [0.100−2.356]	0.138 [0.104−0.822]
h^l	Habit formation, lenders	0.380 [0.156−0.536]	0.629 [0.527−0.634]	0.381 [0.364−0.535]	0.350 [0.171−0.502]
h^b	Habit formation, borrowers	0.614 [0.405−0.775]	0.000 [0.000−0.662]	0.684 [0.323−0.749]	0.574 [0.506−0.749]
ψ	Inverse Frisch elasticity	0.309 [0.258−0.534]	10.000 [2.428−10.000]	0.251 [0.250−0.347]	1.877 [0.267−2.890]
ϕ	Capital adjustment costs	9.826 [3.994−10.187]	10.144 [6.023−14.511]	22.428 [3.857−24.980]	10.238 [8.159−13.598]
γ	Inertia of mortgage debt	0.742 [0.356−0.871]	0.575 [0.300−0.892]	0.832 [0.648−0.912]	0.589 [0.462−0.838]
γ_τ	Tax response to gov't debt	0.529 [0.045−0.778]	0.877 [0.480−0.900]	0.899 [0.051−0.900]	0.348 [0.045−0.580]
ρ_τ	Inertia of tax rate	0.485 [0.383−0.804]	0.501 [0.281−0.750]	0.611 [0.329−0.838]	0.484 [0.431−0.825]
γ_G	Persistence of G shock	0.942 [0.922−0.952]	0.939 [0.871−0.981]	0.955 [0.920−0.977]	0.938 [0.917−0.951]
σ_g	Std. dev. of G shock	0.097 [0.083−0.118]	0.090 [0.086−0.096]	0.099 [0.081−0.112]	0.097 [0.083−0.116]
τ	Love for variety	4.196 [3.494−4.557]	*N/A*	*N/A*	*N/A*
x	Steady-state value of markup	1.139 [1.127−1.156]	*N/A*	*N/A*	*N/A*
μ_g	Productivity of public capital	*N/A*	0.939 [0.466−0.985]	*N/A*	*N/A*
δ_g	Depreciation of public capital	*N/A*	0.378 [0.211−0.545]	*N/A*	*N/A*
θ_n	Skill acquisition from working	*N/A*	*N/A*	1.727 [0.995−2.596]	*N/A*
ρ_x	Persistence of acquired skills	*N/A*	*N/A*	0.990 [0.392−0.990]	*N/A*
χ_2	Tech. utilization adj. cost	*N/A*	*N/A*	*N/A*	0.022 [0.006−0.297]

Notes: The table reports the median parameter estimates from each model, with the 68 percent credible sets (i.e., the 16th and 84th percentiles) reported in brackets.

gation mechanisms. In general, these are found to be larger—in some cases significantly so—than the values typically considered in the existing literature. As discussed in the next section, this reflects that each model requires a relatively strong supply-side response in order to match the magnitude of the increase in house prices and consumption observed in the data. Based on these observations, our analysis in the next section will also consider a version of each model where some realistic parameter bounds have been imposed in the estimation procedure.

In the case of the LOV model (column 1), the data appear to emphasize the role of the variety effect more than that of the competition effect. In fact, the steady-state markup, x, is estimated relatively close to the lower bound set by $\frac{1}{\rho} = 1.11$. Under these circumstances—given the modest estimate of ψ—the condition for consumption crowding-in boils down to obtaining a relatively high value of τ, as shown analytically in Appendix D. With a low steady-state markup, fixed costs are relatively small, and there are relatively many firms with little market power within each sector. As a result, the markup is relatively insensitive to fiscal shocks. Thus, to produce sizeable upward changes in TFP—which are key to bringing about an equilibrium increase in patient households' nondurable consumption and, thus, house prices—a relatively high τ is necessary, so as to amplify the effect of F_t on TFP. These arguments explain why we obtain an estimate of $\tau = 4.196$, which is somewhat higher than what the literature has typically contemplated; for example, 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).[24]

Regarding the PGS model (column 2), the estimated parameters are much higher than existing estimates. Regarding the productivity parameter of public capital (μ_g), we obtain an estimate of 0.939. In comparison, Baxter and King (1993) consider values up to 0.4, corresponding to the highest estimate obtained in the empirical analysis of Aschauer (1989), while Leeper et al. (2010) and Ramey (2020) consider values ranging from 0.05 to 0.12, based on more recent empirical evidence. The estimated depreciation rate of public capital is above one third per quarter, which is clearly at odds with the data. For example, the calibration exercise of Ramey (2020) yields a value of 0.01. While the value of μ_g directly governs the productivity of public spending, a high depreciation rate of public capital drives up the sensitivity of TFP with respect to the *flow* of government spending, thus enhancing the strength of the mechanism.[25]

A similar pattern emerges for the LBD model (third column). The estimated elasticity of skills with respect to past hours worked (θ_n) is 1.727, which is more than an order of magnitude higher than the micro-based estimate of Chang et al. (2002), 0.111, a value also employed by

[24]Lewis and Poilly (2012) estimate a DSGE model featuring love of variety using impulse-response matching to monetary policy shocks, and find that the love of variety parameter is poorly identified.

[25]In this respect, the specification used by Lewis and Winkler (2017) implicitly assumes that the public capital stock depreciates entirely each period, i.e., that $\delta_g = 1$. We have chosen to follow the traditional approach in the literature—e.g., Baxter and King (1993)—in assuming that what matters for production is the *stock* of public capital, not the *flow* of government spending/investment.

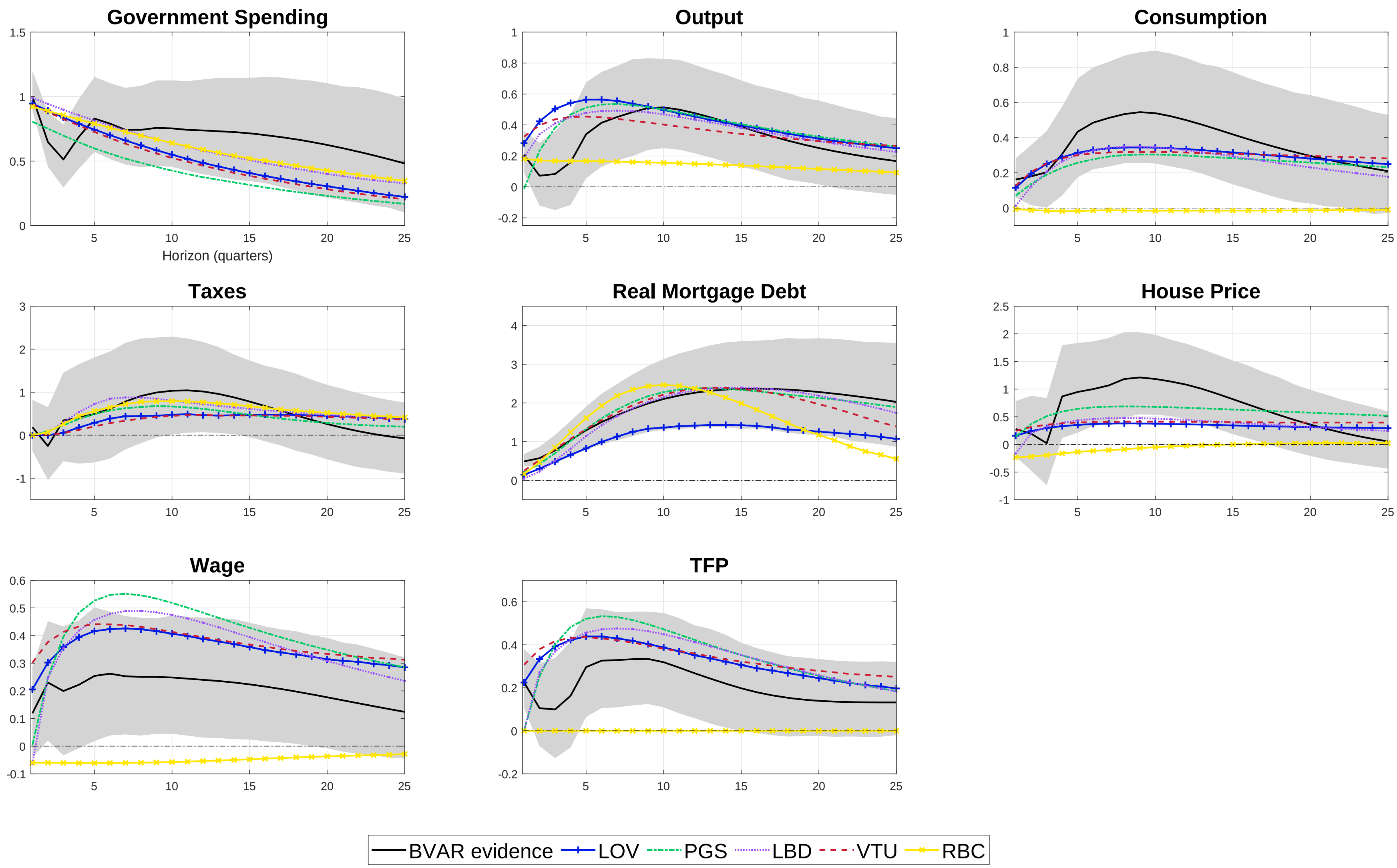
d'Alessandro et al. (2019). This reflects again the inclination of the data to select parameters that entail a large increase in TFP. The persistence of acquired skills is driven to the upper bound of 0.99 imposed in the estimation, enabling the model to match the persistent empirical response of TFP. The persistence estimated by Chang et al. (2002) is somewhat lower, 0.797.

Turning finally to the VTU model (column 4), the estimate of the cost of adjusting the rate of technology utilization is *per se* hard to interpret. One option is to consider the magnitude of the increase in the utilization rate of technology relative to the increase in output, as proposed by Jørgensen and Ravn (2022). In this case, the utilization rate increases roughly one-for-one with output, which is somewhat more than in the estimated model of Jørgensen and Ravn (2022), where the utilization rate increases roughly by half the increase in output. More concerning is the fact that our baseline estimation of the VTU model implies a unit root—i.e., the model is not stationary—and the effects of a government spending shock on output and other variables are therefore permanent. As discussed by Jørgensen and Ravn (2022), the VTU mechanism requires that the adjustment cost displays a certain degree of convexity, for the cost to increase sufficiently with changes in the utilization rate, so as to curb these. In this case, instead, the data prefer a parameterization that gives rise to a large increase in the utilization rate in the short run (recall that our matching exercise considers only the first 25 quarters, so that deviations beyond that horizon are not costly). Thus, generating a sizeable house-price increase again comes at the cost of a less realistic parameterization of the model.

7 Model dynamics

In Figure 5, we report the estimated impulse response functions from the four models, in each case computed as the pointwise median of the distribution of impulse responses we obtain, following Antolín-Díaz and Rubio-Ramírez (2018). We also report the empirical counterparts from the BVAR model and a baseline RBC model with no endogenous TFP. As the figure shows, each model produces responses that largely match the sign and shape of the responses of all variables, including the hump-shaped dynamics exhibited by most of these. In terms of magnitudes, the model-implied responses mostly remain inside the credible regions from the BVAR model. However, all four models overestimate the responses of TFP and the wage rate. This confirms the impression from the parameter estimates that our estimation scheme tends to push the supply-side propagation mechanisms as far as possible, in order to produce a sizeable increase in the house price. In response to the increase in government spending, the positive response of TFP—driven by either of the mechanisms considered above—is key to overcoming the usual negative wealth effect, as discussed in the introduction. Higher productivity drives up the marginal products of labor and capital, and thus the wage rate and the rental rate on capital. In turn, this paves

Figure 5: Estimated effects of a government spending shock



Notes: The figure shows the effects of a shock to government spending. Solid black line: BVAR model. Grey areas: 68 percent credible sets from BVAR model. Solid blue line (with +'s): Estimated model with love of variety (LOV). Dashed-dotted green line: Estimated model with productive government spending (PGS). Dotted purple line: Estimated model with learning by doing (LBD). Dashed red line: Estimated model with variable technology utilization (VTU). Solid yellow line (with x's): RBC model without endogenous productivity mechanisms (RBC).

the way for an increase in consumption, as well as the house price.[26] Yet, in each model, the magnitude of the latter falls short of that observed in the data. Of the models considered, the PGS model generates the largest increase in the house price, but also produces the largest "overshooting" of the responses of TFP and the wage rate. Otherwise, the models with endogenous productivity display only modest differences along other dimensions.

As for the baseline RBC model—whose parameter estimates are relegated to Appendix E—we observe a flat response of TFP, by construction. Consequently, the model produces a decline in the wage rate, consumption (albeit small), and the house price. This also implies a counterfactual negative comovement between house prices and mortgage debt. Overall, these results highlight the necessity of extending the RBC framework in order to account for the empirical evidence.

To gauge the quantitative performance of the model in producing a conditional increase in house prices, Table 3 reports the present-value cumulative fiscal multipliers for the house price that we obtain from the BVAR, alongside those from each of the estimated DSGE models, at three different horizons (on impact, after two years, and at the end of the 25-quarter horizon considered in Figure 5). The table confirms that the PGS model generates the largest response of the house price, which even exceeds the empirical response when measured over the entire 25 quarters (both because the PGS model produces a very persistent house-price response, and because the cumulative increase in government spending, which appears in the denominator of the multiplier, is significantly smaller than in the data). Again, as discussed above, recall that the results from the PGS model are obtained at the cost of less realistic parameter estimates. The other three endogenous productivity models all account for between half and two thirds of the increase in house prices observed in the data—except for the LBD model, on impact—with multipliers well within the credible sets from the BVAR. In contrast, the multipliers of the RBC model with constant TFP always remain negative and outside the estimated credible sets.

Estimation with parameter restrictions The results presented so far have indicated that the endogenous productivity mechanisms we consider can provide a reasonable account of the empirical evidence, qualitatively as well as quantitatively. However, as already mentioned, some of the estimated parameter values are well above conventional values, and in a few cases simply unrealistic. To confront this issue, we now impose a set of plausible parameter restrictions, and then re-estimate each of the models. Specifically, in the LOV model, we set an upper bound of $\tau \leq 2$, corresponding to the highest value employed in existing studies (see Corsetti et al., 2007). In the PGS model, we set an upper bound of $\mu_g \leq 0.4$ to match the highest value considered by Baxter and King (1993), although we note that even this value is higher than suggested by

[26] In each of the four models, the increase in aggregate consumption reflects a positive response of consumption of lenders—which is crucial for the response of the house price—as well as borrowers. The latter is in line with the empirical evidence of Alpanda et al. (2021), based on the U.S. Consumer Expenditure Survey. The estimated RBC model without any of the productivity-enhancing mechanisms features a decline in the consumption of both agents.

Table 3: Present-value house price fiscal multipliers

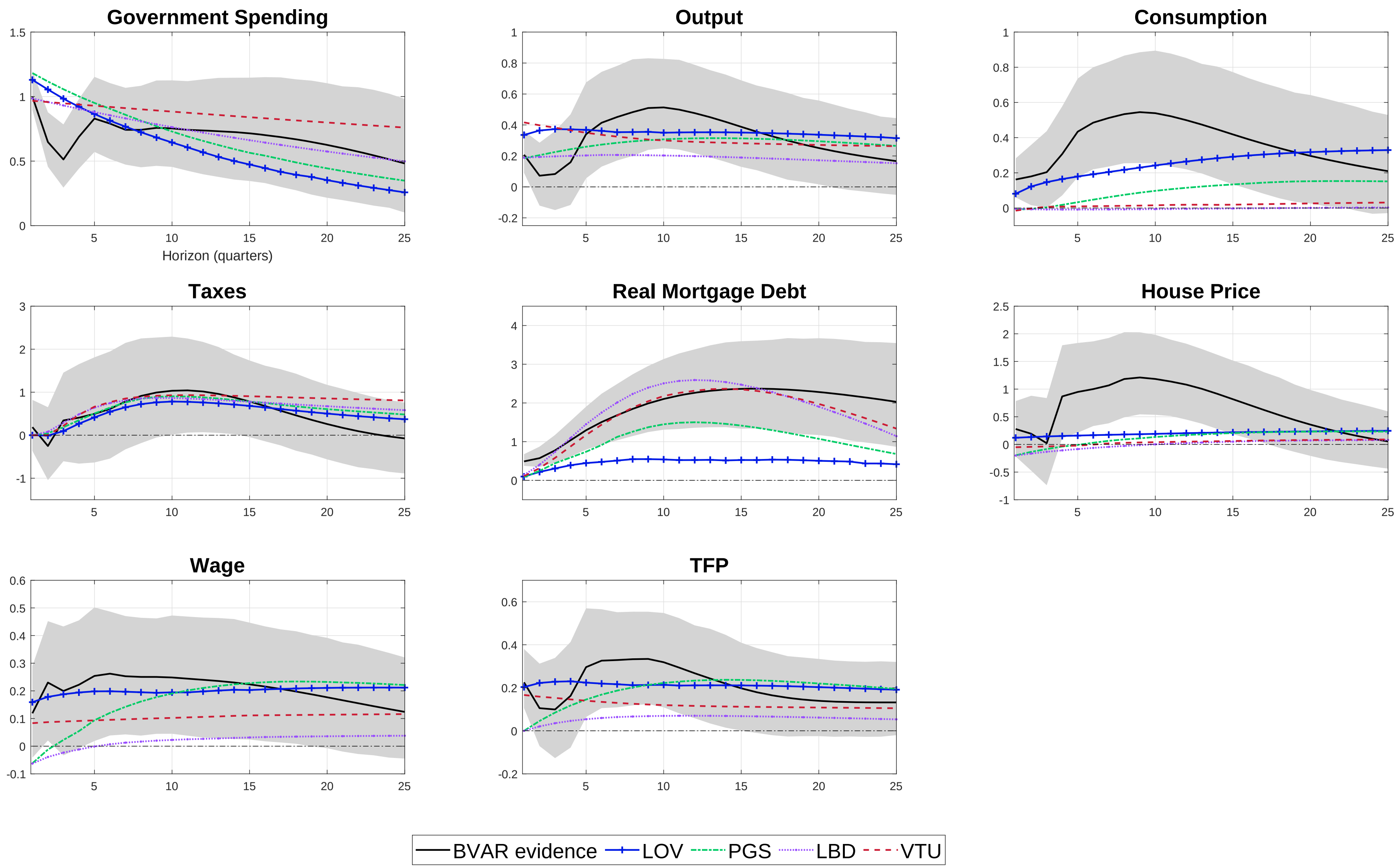
	Impact	8 quarters	25 quarters
Data (BVAR)	0.28 [−0.24;0.65]	0.88 [−0.04;1.41]	0.98 [0.14;1.33]
LOV model	0.17	0.39	0.56
PGS model	0.18	0.81	1.25
LBD model	−0.17	0.36	0.53
VTU model	0.26	0.47	0.69
RBC model	−0.25	−0.20	−0.12

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

more recent studies. Furthermore, we impose an upper bound of $\delta_g \leq 0.1$, representing an upper bound of the depreciation rates of public (as well as private) capital typically considered in the literature. For the LBD model, we use the estimates provided by Chang et al. (2002) and used by d'Alessandro et al. (2019) as the upper bounds, that is, we set $\theta_n \leq 0.111$ and $\rho_x \leq 0.797$. The VTU model is slightly harder to discipline, since the estimated value of χ_2 is hard to interpret, *per se*. Therefore, we choose to impose as a restriction that the model cannot display a unit root. We have found that implementing a lower bound of $\chi_2 \geq 1.5$ is sufficient to meet this requirement.

We report the parameter estimates obtained from this exercise in Table E.2 in Appendix E. Unsurprisingly, the parameters governing the endogenous productivity mechanisms are generally driven (very close) to the bounds we now impose. This is true across all four models, with the possible exception of the LOV model, where we now obtain $\tau = 1.867$, i.e., a value somewhat below the upper bound of 2. The implied impulse responses, reported in Figure 6, indicate that the ability of the models to match the data is severely impaired. Of the four models, only the LOV model produces an increase in the house price on impact and throughout the first year after the shock. Likewise, the house-price fiscal multipliers for this set of models (reported in Table 4) indicate that only the LOV model remains within the credible set from the BVAR at all horizons considered. As for the responses of the remaining variables, only the LOV and PGS models produce a clear increase in consumption, while also providing a reasonable match of the responses of wages and TFP. The LOV model, on the other hand, implies an increase in mortgage debt that falls short of the observed response. All in all, given plausible parameter restrictions, the quantitative bite of the mechanisms we have considered is reduced considerably.

Figure 6: Estimated effects of a government spending shock with parameter restrictions



Notes: The figure shows the effects of a shock to government spending. Solid black line: BVAR model. Grey areas: 68 percent credible sets from BVAR model. Solid blue line (with +'s): Estimated model with love of variety (LOV). Dashed-dotted green line: Estimated model with productive government spending (PGS). Dotted purple line: Estimated model with learning by doing (LBD). Dashed red line: Estimated model with variable technology utilization (VTU). All models are estimated with the parameter restrictions described in the main text.

Table 4: Present-value house price fiscal multipliers with parameter restrictions

	Impact	8 quarters	25 quarters
Data (BVAR)	0.28 [−0.24;0.65]	0.88 [−0.04;1.41]	0.98 [0.14;1.33]
LOV model (restricted)	0.11	0.16	0.28
PGS model (restricted)	−0.17	−0.04	0.11
LBD model (restricted)	−0.20	−0.12	−0.03
VTU model (restricted)	−0.05	−0.02	0.03

Notes: The table reports the present-value cumulative fiscal multipliers for the house price obtained from the BVAR (including 68 percent credible sets in square brackets), along with those from the estimated DSGE models with parameter restrictions. See also notes to Table 3.

7.1 More on the LOV model

We find it worthwhile to shed some more light on the mechanics of the model with love of variety and endogenous entry and exit, since the effects of a government spending shock in such an environment have not previously been studied in the literature. To this end, we first consider the effects of reducing the value of τ. As discussed in Section 5.4, 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 $\tau = 0$. When τ is sufficiently high, 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 (as well as the rental rate of capital). As we show in Figure F.1 in Appendix F, a value of $\tau = 1$ is sufficient for this to occur, holding all other parameters fixed. Instead, if τ is null or very small, the contraction in labor supply dominates, and the real wage drops. Figure F.1 thus confirms the message from the stylized version of the LOV model in Appendix D: When the taste for variety is sufficiently strong, the model produces a positive response of Ricardian agents' consumption, a decline in their shadow value of income, and thus an increase in the house price.

Likewise, Figure F.2 confirms that in the absence of the variety effect (i.e., when $\tau = 0$), the competition effect is insufficient to generate an increase in the house price, even as we vary the level of the steady-state markup, x, which scales the strength of the competition effect. A relatively high value of x corresponds to a rather strong competition effect, since the marginal entrant has higher chances of producing a sizable impact on the markup—and thus on $TFP_t = 1/x_t$—in a relatively concentrated market. Yet, for the values of x we consider, this produces an insufficient increase in TFP, and thus a drop in the house price.

7.2 The (limited) importance of credit constraints

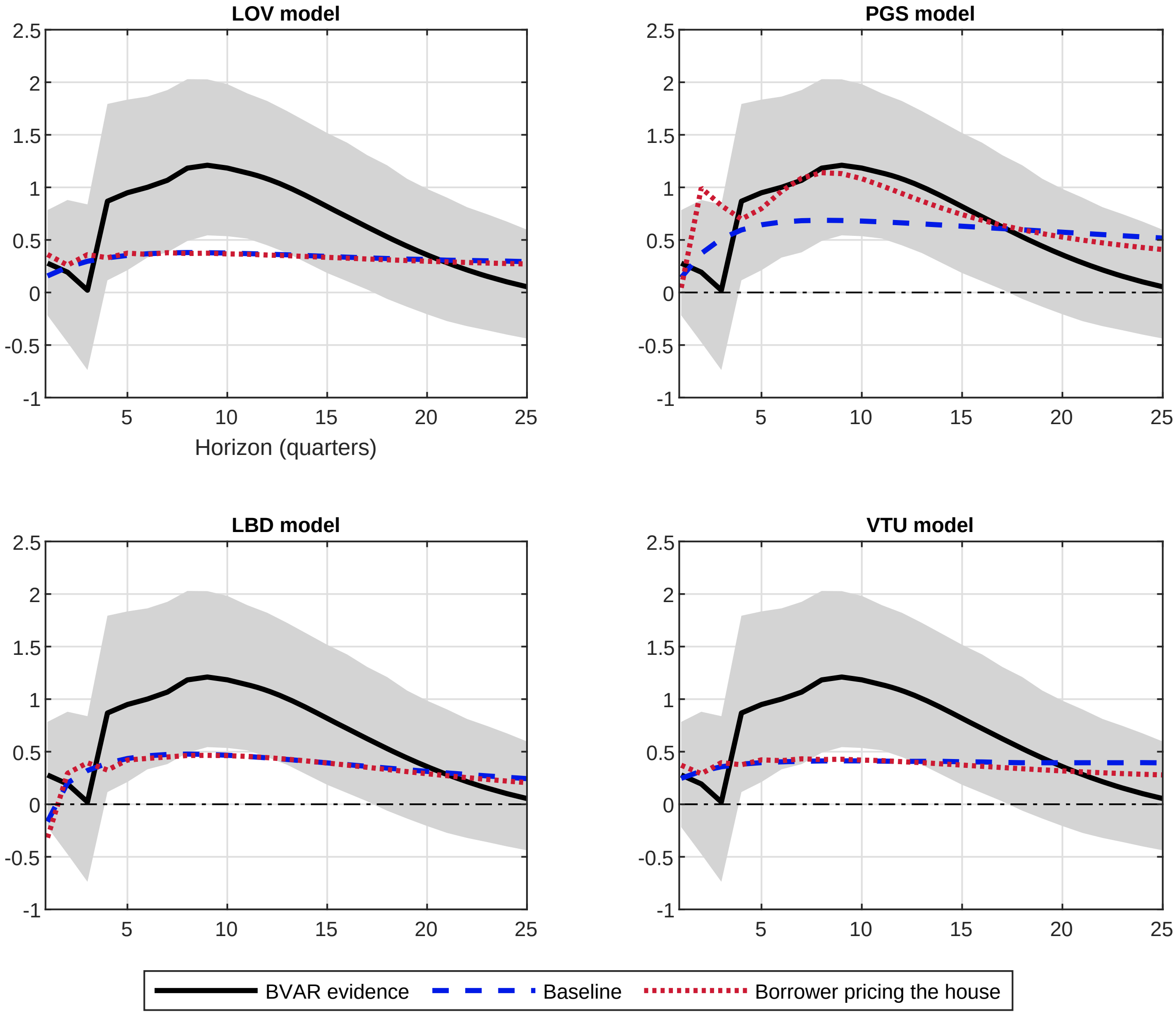
We now return to the role of credit frictions. In Section 3.1, we showed analytically that letting the house price be determined by credit-constrained households did not alter the fundamental insights obtained from a model without financial frictions regarding the effects of a fiscal expansion. To provide a quantitative treatment of this issue, we follow Justiniano et al. (2019) and assume that lenders own a constant share of the aggregate housing stock. This amounts to replacing the lenders' Euler equation for durables with $H_t^l = \bar{H}^l$. Effectively, this implies that the equilibrium house price will be determined by the borrowers, and will therefore be affected by credit frictions, as in Section 3.1.

We then reestimate each of the four models under consideration (without the parameter restrictions imposed above) upon implementing this change. In each panel of Figure 7, we plot the impulse response of the house price from this exercise (dashed red lines) against those obtained from the baseline version of the corresponding model (dotted blue lines), i.e., the same as reported in Figure 5. While some differences can be observed, the figure confirms that the response of house prices to a government spending shock is in the same ballpark, irrespective of whether this price is determined by financially constrained or unconstrained households. The model with productive government spending is the only case where the alternative assumption about house pricing appears to have some effect. However, in this case, we obtain estimates of $\mu_g = 0.957$ and $\delta_g = 0.290$, indicating that the close match of the empirical response comes at the cost of unrealistic parameter estimates. In general, the parameter estimates governing the supply-side propagation mechanisms are similar to those obtained from the baseline set of models: The LOV model produces an estimate of $\tau = 4.905$; the LBD model returns $\theta_n = 1.554$ and $\rho_x = 0.914$; and the VTU model implies a value of $\chi_3 = 0.110$ and an increase in the technology utilization rate very similar to the baseline model. Altogether, this leads us to conclude that letting the house price be determined by credit-constrained households does not lead to any significant improvement in the models' ability to match the data, nor to a reduction in the strength of the supply-side mechanisms required to do so.[27]

We end with a word of caution regarding the interpretation of the results in this subsection. In recent work, Greenwald and Guren (2021) have found that credit conditions may exert a significant effect on house prices, in equilibrium. While our results may appear, at a first glance, to be in contrast with their findings, we want to stress that this is not the case, for two reasons. *First*, we focus on shocks to fiscal spending, and find that the *presence* of credit frictions does not alter the effects of such shocks on house prices, whereas Greenwald and Guren (2021) study whether *changes* in (or shocks to) credit conditions can move house prices. In the context of our model,

[27] The responses of the remaining variables are also quite similar, with the exception of mortgage debt, which now displays a significantly smaller increase, since borrowers are unable to acquire more housing in this case (see Figure E.1 in Appendix E).

Figure 7: Effects of a government spending shock on house prices: The role of credit frictions



Notes: The figure shows the house-price response to a government spending shock under different assumptions regarding which agent is determining the house price in equilibrium. Solid black line: BVAR model. Grey areas: 68 percent credible sets from BVAR model. Dashed blue line: Baseline model with lender pricing the house. Dotted red line: Alternative model with constrained borrower pricing the house.

this could be implemented through stochastic variations in the LTV ratio, m, which would alter the analysis in Section 3.1. *Second*, while the type of credit friction considered in our analysis—collateralized borrowing—has been employed in a wide range of studies, our results do not imply irrelevance for *any* type of credit friction, regarding the effects of fiscal shocks on house prices.

8 Concluding remarks

We report new evidence for the U.S. economy indicating that house prices increase following an unanticipated expansion in fiscal spending. We add to the existing time-series evidence pointing to this fact by documenting a positive response of house prices also at the regional level. Additionally, we observe a positive response of productivity to a fiscal expansion, both at the aggregate and the regional level. This feature plays a key role in accounting for the impact of fiscal spending on house prices, according to a class of dynamic general equilibrium economies featuring an endogenous response of TFP to a fiscal expansion. We consider a range of structural propagation mechanisms capable of generating such a response.

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

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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 analysis, as well as some robustness checks.

A.1 Appendix to the BVAR analysis

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

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

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

$$\mathbf{X}_t = \left[\begin{array}{ccccccccc} FE_t & G_t & Y_t & C_t & T_t & B_t & Q_t & W_t & TFP_t \end{array} \right]'.$$

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.

A.1.1 Data description

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

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

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

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

T_t: Government current tax receipts (W054RC1Q027SBEA) + Government income receipts on assets (W058RC1Q027SBEA) + Government current transfer receipts (W060RC1Q027SBEA) - Government current transfer payments (A084RC1Q027SBEA) - Government interest payments

(A180RC1Q027SBEA) - Government subsidies (GDISUBS).[28] All series are seasonally adjusted. We convert from nominal to real terms using the GDP deflator (GDPDEF).

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

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

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

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

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

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

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

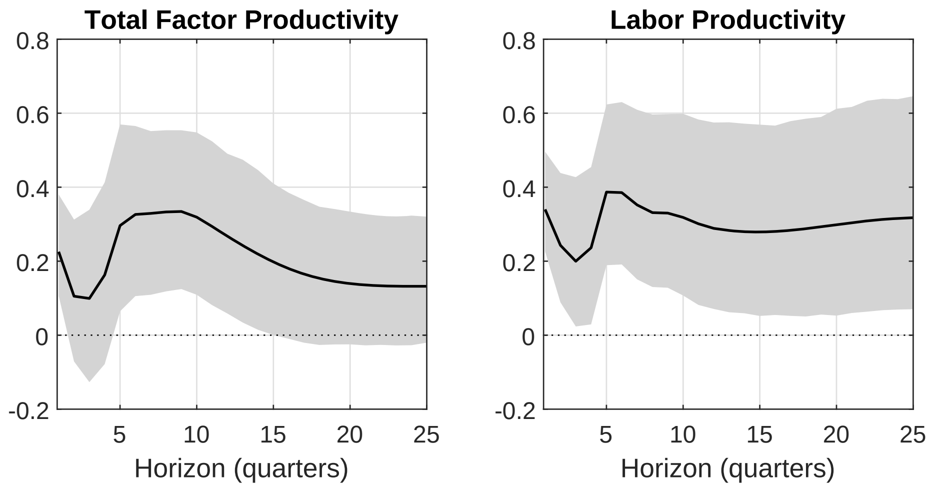
A.1.2 Response of labor productivity

In our BVAR based on aggregate data, we included a measure of TFP. In our regional analysis, we considered instead the response of labor productivity, as measured by output per worker, since no measure of TFP at the MSA level exists. While we cannot verify the response of TFP at the regional level, we can instead study the response of labor productivity at the aggregate level. To this end, we return to our baseline BVAR and replace the TFP series with Labor Productivity (Output per Hour) for All Workers in the Nonfarm Business Sector, obtained from the BLS.[29] We report the response of labor productivity in the right panel of Figure A.1, while the left panel reproduces the response of TFP from our baseline BVAR of Section 2.1 for comparison. As the figure illustrates, the responses of the two series are very similar, both qualitatively and quantitatively, with labor productivity displaying a slightly larger and more persistent response. Thus, at least at the aggregate level, the conditional responses of the two measures of productivity are similar.

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

[29] This series is available in the FRED database (FRED name: OPHNFB, seasonally adjusted, 2012=100). We take logs and detrend the resulting series, as described for the other macro variables in Appendix A.1.1.

Figure A.1: Response of different productivity measures to a government spending shock



Notes: The figure shows the effects of a shock to government spending estimated in the BVAR model (solid black line), with the grey areas representing the 68 percent credible sets (i.e., the 16th and 84th percentiles of the posterior distribution based on 1,000 draws). *Left panel*: Response of Total Factor Productivity obtained from Fernald (2014) (reproduced from Figure 1). *Right panel*: Response of Labor Productivity (Output per Hour) obtained from the BLS.

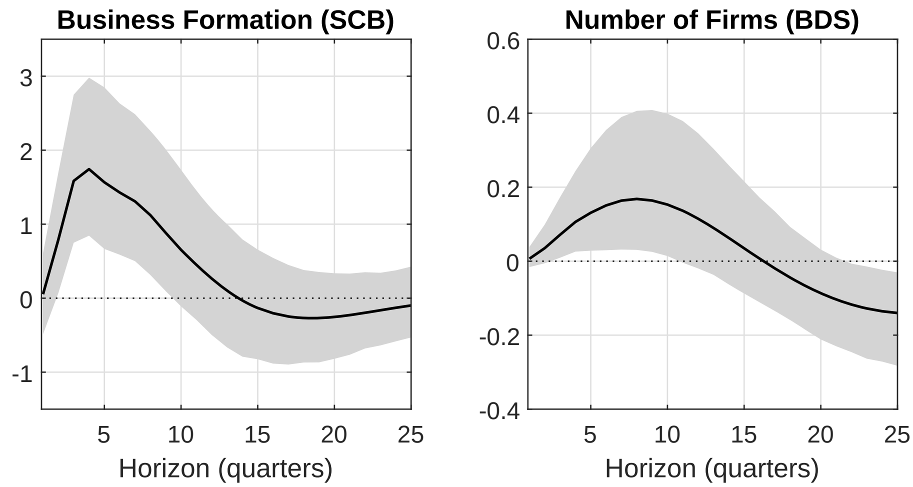
A.1.3 Response of business formation

As discussed in the main text, including a measure of business formation in the BVAR is challenging due to data limitations. We now try to shed some light on the response of business formation by considering two measures of it, each of which is available for a shorter subsample than considered above. The first is the Index of Net Business Formation contained in the Survey of Current Business (SCB) from the BEA.[30] This series was employed by Lewis and Winkler (2017), but was discontinued after 1994. We obtain the quarterly average of monthly observations, and then take logs. The left panel of Figure A.2 reports the response of the SCB index of business formation to an identified government spending shock in the BVAR model for the period 1966:Q4-1994:Q4. We observe a positive and significant response, in line with the evidence reported by Lewis and Winkler (2017).

As an alternative, we also consider the Number of Firms in the United States, obtained from the Business Dynamics Statistics (BDS) produced by the Census Bureau.[31] This series is available at the annual frequency, with the first observation in 1978, although we only use the data from 1979 onwards, following Tian (2018). Since the data is annual, we interpolate the series using a spline in order to obtain a quarterly measure, and then take logs. The right panel of Figure A.2 reports the response of this variable to an identified government spending shock in the BVAR model for the period 1979:Q1-2019:Q4. As seen from the figure, the number of firms also displays a significantly positive response to this type of shock.

[30]The data can be found at https://apps.bea.gov/scb/issues/1996/scb-1996-january-february.pdf (see page C-8).
[31]The data can be found at https://www.census.gov/data/tables/time-series/econ/bds/bds-tables.html.

Figure A.2: Estimated response of business formation to a government spending shock (BVAR)



Notes: The figure shows the effects of a shock to government spending estimated in the BVAR model (solid black line), with the grey areas representing the 68 percent credible sets (i.e., the 16th and 84th percentiles of the posterior distribution based on 1,000 draws). *Left panel*: Response of Index of Net Business Formation from the Survey of Current Business (1966:Q4-1994:Q4). *Right panel*: Response of Number of Firms from the Business Dynamics Statistics (1979:Q1-2019:Q4).

While the responses reported in Figure A.2 are thus—at least qualitatively—consistent with the model featuring love of variety and endogenous firm entry/exit, we want to be cautious in interpreting them. The reason is that the forecast error series of Auerbach and Gorodnichenko (2012) (updated to 2019) turns out to be a much weaker instrument for each of these two subsamples than for our full sample. To arrive at this result, we consider the first-stage regression of government spending on four lags of FE_t and all other variables in the BVAR model (including business formation when relevant), along with a linear and a quadratic trend, as in the BVAR. We do this for our baseline sample period as well as for each of the two subsamples considered when including business formation. The results are reported in Table A.1. As seen from the table, the F-statistic for the full sample (34.1) is well above the threshold for weak instruments of 23.1 of Montiel Olea and Pflueger (2013), while it is clearly below this threshold (i.e., has much lower relevance) for each of the two subsamples considered (17.4 and 13.4, respectively). See Montiel Olea et al. (2021) for a discussion of how to diagnose weak instruments in the context of VAR models.[32]

[32]We have considered splicing the two variables, which would circumvent the weak instruments problem. We have opted against this option, for two reasons: First, the two series are quite different between them: The series obtained from the SCB displays a lot of variation at medium and high frequencies, and a modest upward trend over time. The BDS measure, instead, displays a much stronger upward trend, but very limited short-term fluctuations (reflecting, in part, that the quarterly series is interpolated). Second, for the sample period for which both measures are available (1979:Q1-1994:Q4), they actually display a clear negative correlation (-0.34). Altogether, we therefore do not deem it appropriate to splice the two series.

Table A.1: First-stage F-statistics

	1966:Q4-2019:Q4	1966:Q4-1994:Q4	1979:Q1-2019:Q4
Kleibergen-Paap F-statistic	34.1	17.4	13.4

Notes: The table reports the Kleibergen-Paap F-statistics from the first-stage regression for the three different subsamples considered in the analysis based on aggregate data.

A.2 Appendix to the regional analysis

We now present some details regarding the regional evidence.

A.2.1 Data used in the regional analysis

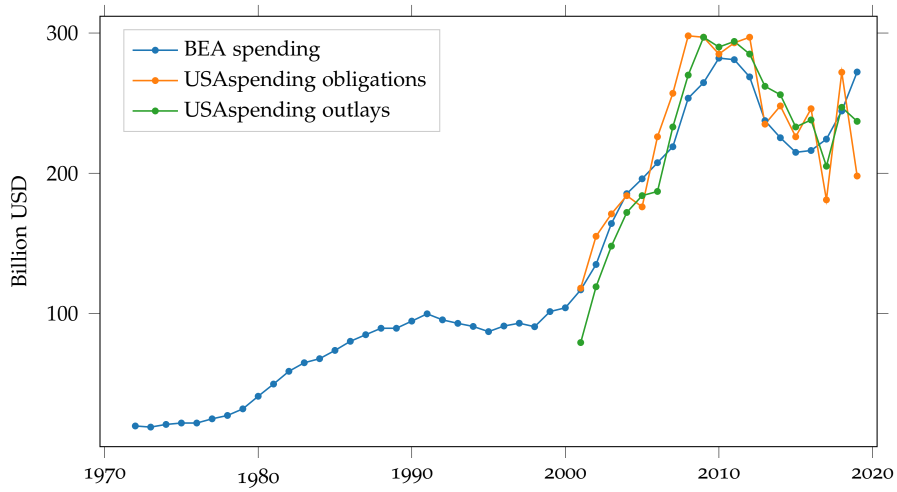
We collect data on contracts signed by firms with the Department of Defense from USAspending.gov to construct the data used in Section 2. The data cover all DoD prime contracts signed from 2001 through 2019, including terminated contracts. The dataset does not contain information on the timing of actual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

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

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

Our data track official data on national military spending from the BEA well in terms of both magnitude and movements. This is seen in Figure A.3, which plots national obligations and our proxy for outlays according to the data from USAspending.gov, together with intermediate goods and services purchased for national defense from the BEA's NIPA tables.

Figure A.3: Military spending according to USAspending.gov and BEA data



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

A.2.2 First-stage estimates

Figure A.4 shows the Kleibergen-Paap F-statistics over different estimation horizons from the first-stage regression

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

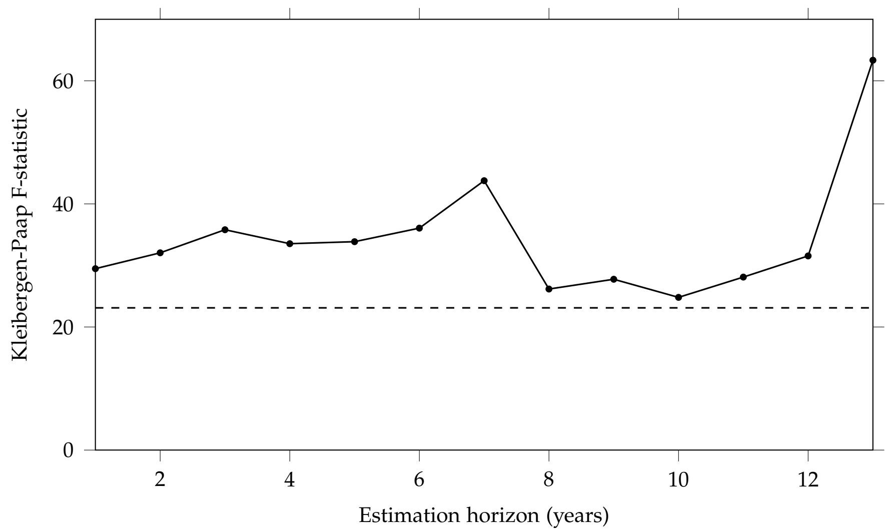
We only show the F-statistics from the first stage to the regression with house price growth as dependent variable. This regression only differs from the first stage to the regressions with establishments or labor productivity growth in that it has two lags of house price growth as controls, rather than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

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

This section presents a pre-trend analysis of our regional estimates of the house price response in Section 2. The blue line in Figure A.5 shows the estimates of the house price response from a modification of our baseline regression in which we do not control for lags of house price growth and extend the estimation horizon to four years prior to the spending change (that is, for horizons $h = -4, -3, \ldots, 13$). The red line plots the estimates from our baseline regression.

The blue line indicates that house prices are slightly, albeit significantly, lower in the two years before a change in military spending. Reassuringly, we see that the estimates of the house price response after a spending change are similar for both regressions. Hence, whether or not we

Figure A.4: F-statistics from first-stage regression



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

control for the relatively small pre-trend in house prices matters little for our results.

A.2.4 Endogeneity analysis of the estimates

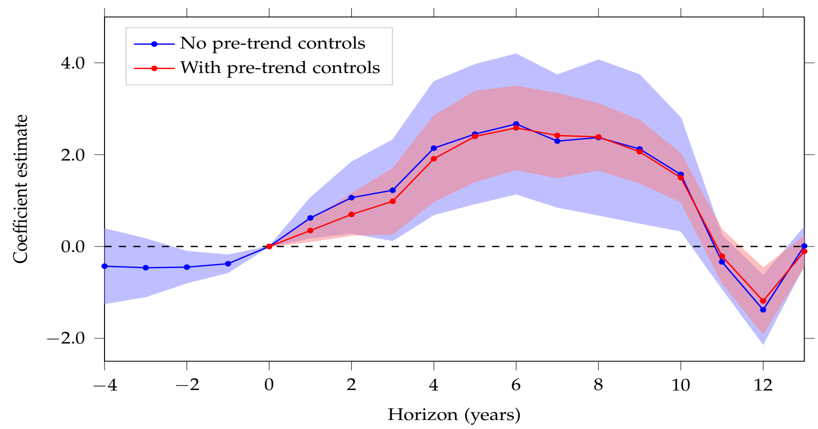
This section presents an endogeneity analysis of our regional estimates of the house price response in Section 2. We test the null hypothesis of the endogenous regressor in the IV regression being exogenous using a Hausman test (i.e. testing that the IV and OLS estimates are identical). The resulting p-values are very close to zero up until year 11 as shown in Figure A.6, supporting the need for using an IV approach.

A.2.5 Robustness of the regional house price estimates

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

Table A.2 shows the IV estimates from alternative specifications of regression (2.1). Estimates from the baseline specification also shown in Figure 2 are presented in column (1). Column (2) shows the estimates from a regression in which house prices, DoD spending, and GDP have been deflated by the MSA-level GDP deflator. Column (3) reports estimates from a regresion in which we use the proxy for outlays described in Appendix A.2.1 to measure DoD spending. Columns (4) and (5) present estimates with alternative normalizations of DoD spending (by personal income and population in thousand persons, respectively). Column (6) controls for house price

Figure A.5: Pre-trend analysis of house price responses to military spending



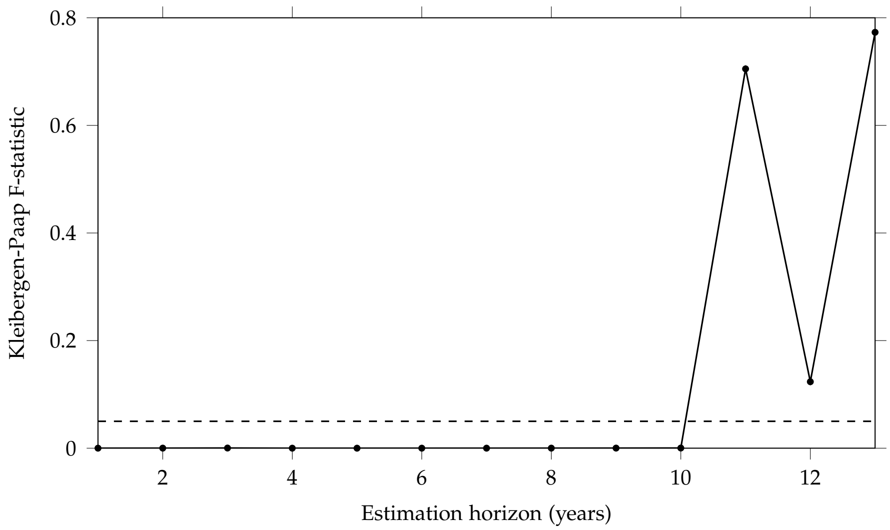
Notes: The figure shows the IV estimates of β_h from regression (2.1) based on an annual panel of 380 MSAs covering the period 2001-2019. The red line shows the baseline specification, which includes as controls two lags of the one-year growth in house prices, two lags of the instrument, and two lags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

movements associated with industry composition, by adding to the regression 2-digit industry employment shares multiplied by year dummies. Column (7) controls for differential exposure to aggregate house price movements by adding to the regression three time-invariant controls multiplied by year dummies: the Wharton Regulation Index, the Saiz (2010) instrument, and the Bartik-like instrument for sensitivity to regional house price movements by Guren et al. (2021). Column (8) adds state × year fixed effects to control for state-specific house price growth fluctuations. Finally, column (9) reports estimates from a version of the regression in which the instrument is constructed using the average DoD spending share over 2001 and 2002. By using beginning-of-sample spending shares, we avoid endogeneity issues that could potentially stem from the within-sample DoD spending shares being correlated with economic fluctuations.

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

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

Figure A.6: p values from endogeneity test



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

Using non-winsorized changes in local spending has little impact on the estimates, while removing MSAs in the top and bottom DoD spending share distribution or winsorized observations amplifies the house price response.

Finally, we show the sensitivity of the estimates to the inclusion of control variables in Table A.4. Column (1) shows estimates from regressions without any controls. Adding lagged spending and instruments lowers the estimates across the entire horizon, as seen in column (2). This suggests that the lag exogeneity condition is not met unless we condition on past values of both spending growth and the instrument itself. When lags of the dependent variable are added as controls, this has limited impact on the estimates, as shown in column (3). Adding lags of the dependent variables to the regression with lagged spending and instruments—thereby retrieving our baseline results—reduces the standard errors (see column (4)), thus improving the efficiency of the estimators.

A.2.6 The response of establishments at the regional level

To study the response of business formation to a change in local government spending, we return to regression (2.1), this time using (the growth rate of) the local number of firms as the dependent variable. The number of firms is measured as the number of establishments within the MSA. Data on establishments are from the County Business Patterns from the U.S. Census Bureau, which contains information on the stock of establishments at the county level. We aggregate these data to get an MSA-level series on establishment counts.

Figure A.7 presents the estimates for the response of establishments. As in the case of produc-

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

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

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

tivity, the IV estimates for the responses of establishments to military spending are in line with the predictions of the model featuring love of variety and endogenous firm turnover. The number of establishments increases following a rise in government spending, and eventually reverts back to its local trend, thus validating the central feature of our model. This finding is consistent with the evidence reported by Auerbach et al. (2024). Furthermore, Figure A.8 reports the scatter plot of the IV estimates of the impulse responses of house prices and establishments (at a six-year horizon), based on 1,000 iterations from a cluster bootstrap drawing sets of MSAs from our sample. For the sets of MSAs where the estimated house price response is larger, so is the response of establishments.

A.2.7 Robustness of the regional labor productivity and establishments estimates

This section reports some robustness exercises regarding the responses of labor productivity and the number of establishments. Table A.5 reports the IV estimates from the same alternative specifications as those reported in Table A.2. The estimates that differ from the baseline are those

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

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

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

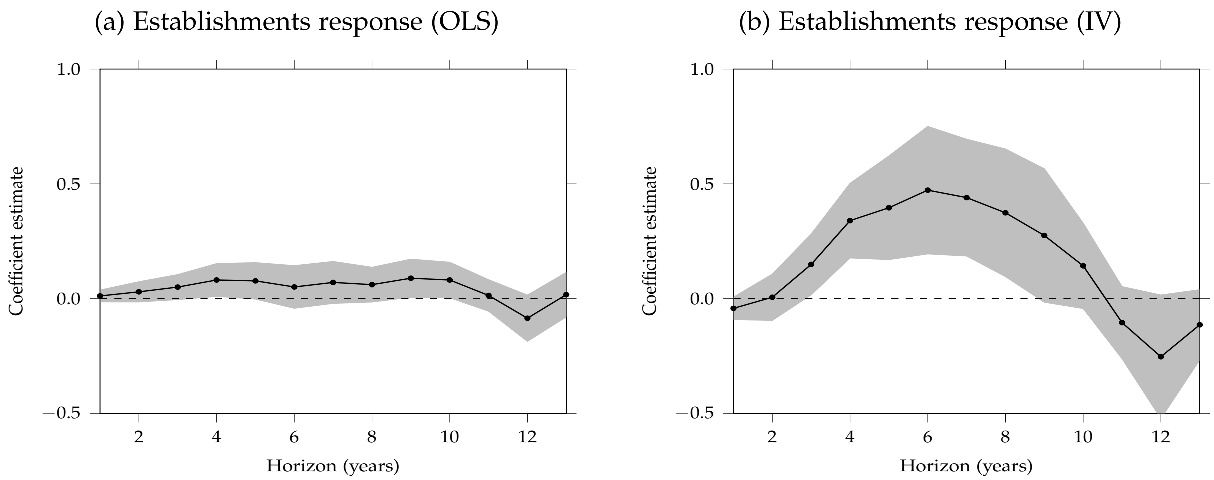
Table A.4: Robustness of regional house price estimates (controls)

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

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

from the specification controlling for housing exposure, which are smaller than the baseline and in many cases insignificant; see column (7). For the labor productivity response, this seems to be driven by the model being estimated on the subsample of MSAs for which we have data on exposure to aggregate house price fluctuations. If we estimate the model on this subsample but do not control for differential exposure, the results are broadly similar to those in column

Figure A.7: The regional response of establishments to military spending



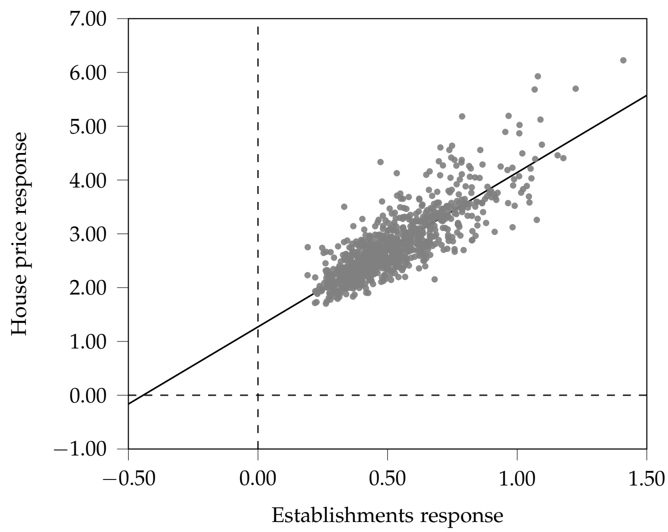
Notes: The figure shows the estimates of β_h from regression (2.1) based on an annual panel of 380 MSAs covering the period 2001-2019, with the number of establishments as the dependent variable. The OLS and IV estimates are plotted in panels (a) and (b), respectively. The regressions include as controls two lags of the one-year growth in 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.

(7). For the establishments response, the results reflect a combination of the reduced number of MSAs in the regression and the fact that, unconditionally, local establishment growth is positively correlated with local house price growth (see also Epstein et al., 2023), which in turn is closely related to the measures of house price exposure that we control for. In contrast, there are no statistically significant pairwise correlations between the house price exposure measures and our instrument, which rules out potential endogeneity concerns stemming from differential exposure to aggregate house price fluctuations.

Turning to the sensitivity to outliers, Table A.6 shows the sensitivity of the labor productivity and establishments response estimates to outliers using the same checks as those in Table A.3. The results are similar to those from the house price response estimates.

Table A.7 reports the sensitivity of our baseline estimates to the control variables. The main take-aways from this exercise are in line with those from Table A.4. While controlling for lags of spending growth and of the instrument generally lowers the response estimates, adding lags of the dependent variables tends to reduce the standard errors.

Figure A.8: The joint distribution of house price and establishments responses



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

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

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

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

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

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

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

Table A.7: Robustness of regional establishments and labor productivity estimates (controls)

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

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

B Model appendix: Common block

We now present some additional details of the model, starting with the common model elements outlined in Section 4.

B.1 Households' first-order conditions

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

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

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

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

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

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

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

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

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

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

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

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

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 Steady state

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

Observe first that, irrespective of the assumed production structure, the steady-state factor income shares can be derived from (4.6) and firms' cost-minimization conditions, and expressed as follows:

$$\frac{r^k K}{Y} = \mu, \tag{B.11}$$

$$\frac{w^b N^b}{Y} = (1-\mu)\alpha, \tag{B.12}$$

$$\frac{w^l N^l}{Y} = (1-\mu)(1-\alpha). \tag{B.13}$$

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

$$R = \frac{1}{\beta^l}, \tag{B.14}$$

$$r^k = \frac{1}{\beta^l} - (1-\delta)\,. \tag{B.15}$$

The capital-to-output ratio is derived from (B.11) and (B.15):

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

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 (4.5) and the aggregate resource constraint (4.12) gives us the consumption-to-output ratio:

$$\frac{C}{Y} = 1 - \bar{G} - \frac{\delta\mu}{\frac{1}{\beta^l} - (1-\delta)}. \tag{B.16}$$

The steady-state version of the government budget constraint (4.10) implies:

$$\frac{\tau^{TOT}}{Y} = \left(\frac{1}{\beta^l} - 1\right)\frac{B^g}{Y} + \frac{G}{Y}, \tag{B.17}$$

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.14) are given by

$$q\lambda^b = \Upsilon^b \left(H^b\right)^{-\sigma_h} + \beta^b \lambda^b q + \mu^b m \beta^l q(1-\gamma), \tag{B.18}$$

$$\mu^b = \lambda^b \frac{1 - \frac{\beta^b}{\beta^l}}{1 - \beta^b \gamma}. \tag{B.19}$$

Substituting the latter equation into the former yields

$$Y^b \left(H^b\right)^{-\sigma_h} = q\lambda^b \left[1 - \beta^b - \frac{\beta^l - \beta^b}{1 - \beta^b \gamma} m(1-\gamma)\right]. \tag{B.20}$$

The housing demand equation for the patient households is given by

$$Y^l \left(H^l\right)^{-\sigma_h} = q\lambda^l \left(1 - \beta^l\right). \tag{B.21}$$

Dividing (B.21) by (B.20) and inserting the steady-state expressions for the budget constraint multipliers—which follow directly from (B.1) and (B.5)—together with the consumption and housing market clearing conditions (4.13) and (4.14) gives us an expression for the housing and consumption shares of the impatient households:

$$\begin{aligned}
\frac{Y^l \left(H^l\right)^{-\sigma_h}}{Y^b \left(H^b\right)^{-\sigma_h}} &= \frac{q\lambda^l \left(1-\beta^l\right)}{q\lambda^b \left[1 - \beta^b - \frac{\beta^l - \beta^b}{1-\beta^b\gamma} m(1-\gamma)\right]} \\
\left(\frac{H}{H^b} - 1\right)^{-\sigma_h} &= \frac{Y^b}{Y^l} \frac{1-\beta^l}{1 - \beta^b - \frac{\beta^l - \beta^b}{1-\beta^b\gamma} m(1-\gamma)} \frac{\lambda^l}{\lambda^b} \\
\left(\frac{H}{H^b} - 1\right)^{-\sigma_h} &= \frac{Y^b}{Y^l} \frac{1-\beta^l}{1 - \beta^b - \frac{\beta^l - \beta^b}{1-\beta^b\gamma} m(1-\gamma)} \frac{\left(1 - h^l\beta^l\right)\left(\left(1-h^l\right)\left(C - C^b\right)\right)^{-\sigma_c}}{\left(1 - h^b\beta^b\right)\left(\left(1-h^b\right)C^b\right)^{-\sigma_c}} \\
\left(\frac{H}{H^b} - 1\right)^{-\sigma_h} &= \frac{Y^b}{Y^l} \frac{1-\beta^l}{1 - \beta^b - \frac{\beta^l - \beta^b}{1-\beta^b\gamma} m(1-\gamma)} \frac{1-\beta^l h^l}{1-\beta^b h^b} \left(\frac{1-h^l}{1-h^b}\right)^{-\sigma_c} \left(\frac{C}{C^b} - 1\right)^{-\sigma_c}.
\end{aligned} \tag{B.22}$$

Similarly, we derive an additional expression for the housing and consumption shares of the impatient households by inserting the borrowing constraint (4.3) into their budget constraint (4.2), and using the interest rate (B.14), the labor income share (B.12), and the lump-sum tax payment (4.8):

$$\frac{C^b}{C} = \frac{Y}{C} \left[\left(\beta^l - 1\right) m \frac{qH}{Y} \frac{H^b}{H} + \alpha \left(1 - \mu - \frac{\tau^{TOT}}{Y}\right)\right]. \tag{B.23}$$

The housing wealth-to-output ratio, $\frac{qH}{Y}$, is calibrated, while the consumption share, $\frac{C}{Y}$, follows from (B.16), so (B.22) and (B.23) 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.3 Log-linearized equations

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.3.1 Optimality conditions of the impatient households

Log-linearization of (B.1), (B.3), and (4.3) gives us the following:

$$\hat{\lambda}_t^b = -\frac{\sigma_c^b}{(1-\beta^b h^b)(1-h^b)} \left(\hat{C}_t^b - h^b \hat{C}_{t-1}^b - \beta^l h^b E_t \left\{ \hat{C}_{t+1}^b - h^b \hat{C}_t^b \right\} \right), \tag{B.24}$$

$$\hat{w}_t^b + \lambda_t^b = \psi \hat{N}_t^b, \tag{B.25}$$

$$\widehat{B}_t = \gamma \widehat{B}_{t-1} + (1-\gamma) \left(\mathrm{E}_t \widehat{q}_{t+1} + \widehat{H}_t^b - \widehat{R}_t \right). \tag{B.26}$$

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.14), (B.19), and (B.20):

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

$$\begin{aligned} \hat{q}_t + \hat{\lambda}_t^b &= -\sigma_h \left[1 - \beta^b - \frac{\beta^l - \beta^b}{1-\beta^b \gamma} m (1-\gamma) \right] \hat{H}_t^b + \beta^b E_t \left\{ \hat{\lambda}_{t+1}^b + \hat{q}_{t+1} \right\} \\ & \quad + \frac{\beta^l - \beta^b}{1-\beta^b \gamma} m (1-\gamma) E_t \left\{ \hat{\mu}_t^b + \hat{q}_{t+1} - \hat{R}_t \right\}. \end{aligned} \tag{B.27}$$

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.19) results in

$$\widehat{\lambda}_t^b + \beta^b \gamma \frac{1 - \frac{\beta^b}{\beta^l}}{1-\beta^b \gamma} E_t \left\{ \widehat{\mu}_{t+1}^b \right\} = \frac{1 - \frac{\beta^b}{\beta^l}}{1-\beta^b \gamma} \widehat{\mu}_t^b + \frac{\beta^b}{\beta^l} E_t \left\{ \widehat{\lambda}_{t+1}^b + \widehat{R}_t \right\}. \tag{B.28}$$

The log-linearized budget constraint becomes

$$\begin{aligned} & \frac{C^b}{C} \frac{C}{Y} \hat{C}_t^b + \frac{qH}{Y} \frac{H^b}{H} \left(\hat{H}_t^b - \hat{H}_{t-1}^b \right) + m \frac{qH}{Y} \frac{H^b}{H} \left(\hat{R}_{t-1} + \hat{B}_{t-1} \right) = \\ & (1-\mu) \alpha \left(\hat{w}_t^b + \hat{N}_t^b \right) + m \frac{qH}{Y} \frac{H^b}{H} \beta^l \hat{B}_t - \alpha \frac{\tau^{TOT}}{Y} \hat{\tau}_t^{TOT}. \end{aligned} \tag{B.29}$$

B.3.2 Optimality conditions of the patient households

Log-linearization of (B.5), (B.7), (B.8), and (4.5) results in

$$\hat{\lambda}_t^l = -\frac{\sigma_c^l}{(1-\beta^l h^l)(1-h^l)}\left(\hat{C}_t^l - h^l\hat{C}_{t-1}^l - \beta^l h^l E_t\left\{\hat{C}_{t+1}^l - h^l\hat{C}_t^l\right\}\right), \tag{B.30}$$

$$\hat{\lambda}_t^l = E_t\left\{\hat{\lambda}_{t+1}^l\right\} + \hat{R}_t, \tag{B.31}$$

$$\hat{w}_t^l + \hat{\lambda}_t^l = \psi\hat{N}_t^l, \tag{B.32}$$

$$\hat{K}_t = (1-\delta)\hat{K}_{t-1} + \delta\hat{I}_t. \tag{B.33}$$

Similar to the derivation for the impatient households, log-linearization of (B.6) in combination with (B.21) becomes

$$\hat{q}_t + \hat{\lambda}_t^l = -\sigma_h\left(1-\beta^l\right)\hat{H}_t^l + \beta^l E_t\left\{\hat{q}_{t+1} + \hat{\lambda}_{t+1}^l\right\}. \tag{B.34}$$

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.33) results in

$$\phi\left(\hat{K}_t - \hat{K}_{t-1}\right) + \hat{\lambda}_t^l = E_t\left\{\hat{\lambda}_{t+1}^l + \beta^l r^k\hat{r}_{t+1}^k + \beta^l\phi\left(\hat{K}_{t+1} - \hat{K}_t\right)\right\}. \tag{B.35}$$

Lastly, the budget constraint (4.4) becomes

$$\begin{aligned}&\frac{C^l}{C}\frac{C}{Y}\hat{C}_t^l + \frac{qH}{Y}\frac{H^l}{H}\left(\hat{H}_t^l - \hat{H}_{t-1}^l\right) + \frac{\delta K}{Y}\hat{I}_t + m\frac{qH}{Y}\frac{H^b}{H}\beta^l\hat{B}_t + \Xi\hat{B}_t^g = \\ &(1-\mu)(1-\alpha)\left(\hat{w}_t^l + \hat{N}_t^l\right) + m\frac{qH}{Y}\frac{H^b}{H}\left(\hat{R}_{t-1} + \hat{B}_{t-1}\right) + \mu\left(\hat{r}_t^k + \hat{K}_{t-1}\right) \\ &+ \frac{1}{\beta^l}\Xi\left(\hat{R}_{t-1} + \hat{B}_{t-1}^g\right) - (1-\alpha)\frac{\tau_t^{TOT}}{Y}\hat{\tau}_t^{TOT}.\end{aligned} \tag{B.36}$$

B.3.3 Fiscal policy, market clearing conditions, and shock processes

The log-linearized version of the government's budget constraint (4.10) is

$$\frac{1}{\beta^l}\Xi Y\left(\hat{R}_{t-1} + \hat{B}_{t-1}^g\right) + \theta\hat{G}_t = \frac{\tau^{TOT}}{Y}\hat{\tau}_t^{TOT} + \Xi\hat{B}_t^g, \tag{B.37}$$

while the adjustment rule for the tax level is given by

$$\hat{\tau}_t^{TOT} = \rho_\tau \hat{\tau}_{t-1}^{TOT} + (1-\rho_\tau)\,\gamma_\tau \left(\hat{B}_{t-1}^g - \hat{Y}_{t-1}\right). \tag{B.38}$$

The market clearing conditions (4.12), (4.13), and (4.14) become

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

$$\hat{C}_t = \frac{C^b}{C}\hat{C}_t^b + \frac{C^l}{C}\hat{C}_t^l, \tag{B.39}$$

$$0 = \frac{H^b}{H}\hat{H}_t^b + \frac{H^l}{H}\hat{H}_t^l. \tag{B.40}$$

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 (4.7) is

$$\hat{G}_t = \gamma_g \hat{G}_t + \epsilon_{g,t}. \tag{B.41}$$

C Model appendix: Supply-side mechanisms

We now turn to the details of the production structure, considering each mechanism in turn.

C.1 Productive government spending

The first-order conditions for the demand for input factors are the following:

$$r_t^K = \mu K_{t-1}^{\mu-1}\left[\left(N_t^b\right)^\alpha \left(N_t^l\right)^{1-\alpha}\right]^{1-\mu}\left[K_{t-1}^g\right]^{\mu_g}, \tag{C.1}$$

$$w_t^b = \alpha\,(1-\mu)\,K_{t-1}^{\mu}\left(N_t^b\right)^{\alpha(1-\mu)-1}\left(N_t^l\right)^{(1-\alpha)(1-\mu)}\left[K_{t-1}^g\right]^{\mu_g}, \tag{C.2}$$

$$w_t^l = (1-\alpha)\,(1-\mu)\,K_{t-1}^{\mu}\left(N_t^b\right)^{\alpha(1-\mu)}\left(N_t^l\right)^{(1-\alpha)(1-\mu)-1}\left[K_{t-1}^g\right]^{\mu_g}, \tag{C.3}$$

since firms take the public capital stock as given. The steady-state level of the public capital stock follows directly from the steady-state version of the law of motion for public capital (5.2), which reads as $\delta_g K^g = I^g$, where I^g is given by the (exogenous) level of public spending.

The log-linearized equilibrium conditions characterizing the economy's supply-side in this model are the following:

$$\hat{Y}_t = \mu \hat{K}_{t-1} + (1-\mu)\left(\alpha \hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right) + \mu_g \hat{K}_{t-1}^g, \tag{C.4}$$

$$\hat{r}_t^k = (\mu - 1)\hat{K}_{t-1} + (1-\mu)\left(\alpha \hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right) + \mu_g \hat{K}_{t-1}^g, \tag{C.5}$$

$$\hat{w}_t^b = \mu \hat{K}_{t-1} + (1-\mu)\left(\left(\alpha - \frac{1}{(1-\mu)}\right)\hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right) + \mu_g \hat{K}_{t-1}^g, \tag{C.6}$$

$$\hat{w}_t^l = \mu \hat{K}_{t-1} + (1-\mu)\left(\alpha \hat{N}_t^b + \left(1 - \alpha - \frac{1}{(1-\mu)}\right)\hat{N}_t^l\right) + \mu_g \hat{K}_{t-1}^g, \tag{C.7}$$

$$\hat{K}_t^g = (1-\delta_g)\hat{K}_{t-1}^g + \delta_g \hat{G}_t. \tag{C.8}$$

Finally, the model-specific measure of total factor productivity is given by $TFP_t = \left(K_{t-1}^g\right)^{\mu_g}$, which on linearized form can be written as:

$$T\hat{F}P_t = \mu_g \hat{K}_{t-1}^g. \tag{C.9}$$

C.2 Learning by doing

In the model with learning by doing, firms take the stock of accummulated skills as exogenously given, following the previous literature. Thus, we can write the first-order conditions as:

$$r_t^K = \mu K_{t-1}^{\mu-1}\left[\left(X_t^b N_t^b\right)^{\alpha}\left(X_t^l N_t^l\right)^{1-\alpha}\right]^{1-\mu}, \tag{C.10}$$

$$w_t^b = \alpha(1-\mu) K_{t-1}^{\mu}\left(X_t^b N_t^b\right)^{\alpha(1-\mu)-1} X_t^b \left(X_t^l N_t^l\right)^{(1-\alpha)(1-\mu)}, \tag{C.11}$$

$$w_t^l = (1-\alpha)(1-\mu) K_{t-1}^{\mu}\left(X_t^b N_t^b\right)^{\alpha(1-\mu)}\left(X_t^l N_t^l\right)^{(1-\alpha)(1-\mu)-1} X_t^l. \tag{C.12}$$

The steady-state level of skills for each type of agent is determined from the evolution of skills, which in steady state becomes: $X^j = \left(N^j\right)^{\frac{\theta_n}{1-\rho_x}}$. For this model version, the log-linearized equilibrium conditions are as follows:

$$\hat{Y}_t = \mu \hat{K}_{t-1} + (1-\mu)\left[\alpha\left(\hat{N}_t^b + \hat{X}_t^b\right) + (1-\alpha)\left(\hat{N}_t^l + \hat{X}_t^l\right)\right], \tag{C.13}$$

$$\hat{r}_t^k = (\mu - 1)\hat{K}_{t-1} + (1-\mu)\left[\alpha\left(\hat{N}_t^b + \hat{X}_t^b\right) + (1-\alpha)\left(\hat{N}_t^l + \hat{X}_t^l\right)\right], \tag{C.14}$$

$$\hat{w}_t^b = \mu \hat{K}_{t-1} + (1-\mu)\left[\left(\alpha - \frac{1}{(1-\mu)}\right)\hat{N}_t^b + \alpha \hat{X}_t^b + (1-\alpha)\left(\hat{N}_t^l + \hat{X}_t^l\right)\right], \tag{C.15}$$

$$\hat{w}_t^l = \mu \hat{K}_{t-1} + (1-\mu)\left[\alpha\left(\hat{N}_t^b + \hat{X}_t^b\right) + \left(1 - \alpha - \frac{1}{(1-\mu)}\right)\hat{N}_t^l + (1-\alpha)\hat{X}_t^l\right], \tag{C.16}$$

$$\hat{X}_t^b = \rho_x \hat{X}_{t-1}^b + \theta_n \hat{N}_{t-1}^b, \tag{C.17}$$

$$\hat{X}_t^l = \rho_x \hat{X}_{t-1}^l + \theta_n \hat{N}_{t-1}^l, \tag{C.18}$$

where the last two equations describe the dynamics of skill acquisition of each type of agent. The expression for total factor productivity is now given by $TFP_t = \left(X_t^b\right)^{\alpha(1-\mu)} \left(X_t^l\right)^{(1-\alpha)(1-\mu)}$, or, in linearized form:

$$T\hat{F}P_t = \alpha(1-\mu)\hat{X}_t^b + (1-\alpha)(1-\mu)\hat{X}_t^l. \tag{C.19}$$

C.3 Variable technology utilization

In this version of the model, the first-order conditions characterizing the optimal choices of labor and capital inputs are given by:

$$r_t^K = \mu u_t K_{t-1}^{\mu-1} \left[\left(N_t^b\right)^{\alpha} \left(N_t^l\right)^{1-\alpha} \right]^{1-\mu}, \tag{C.20}$$

$$w_t^b = \alpha\left(1-\mu\right) u_t K_{t-1}^{\mu} \left(N_t^b\right)^{\alpha(1-\mu)-1} \left(N_t^l\right)^{(1-\alpha)(1-\mu)}, \tag{C.21}$$

$$w_t^l = \left(1-\alpha\right)\left(1-\mu\right) u_t K_{t-1}^{\mu} \left(N_t^b\right)^{\alpha(1-\mu)} \left(N_t^l\right)^{(1-\alpha)(1-\mu)-1}. \tag{C.22}$$

In addition, firms now have an additional choice variable, as they can optimally choose the rate of technology utilization. This gives rise to an additional first-order condition, which takes the following form:

$$z'\left(u_t\right) = K_{t-1}^{\mu} \left[\left(N_t^b\right)^{\alpha} \left(N_t^l\right)^{1-\alpha} \right]^{1-\mu},$$

or simply

$$z'\left(u_t\right) = \frac{Y_t}{u_t}. \tag{C.23}$$

This condition equates the marginal gain from raising the rate of technology utilization (given by the additional amount produced; the right-hand side of the equation) to the marginal cost of doing so, which is determined by the derivative of the cost function of adjusting the utilization rate. In steady state, where the rate of utilization equals one, this simplifies to $z'\left(u\right) = Y$, which can be rewritten—imposing the functional form of $z\left(\cdot\right)$ presented in the main text—as $\chi_1 = Y$. This pins down the parameter χ_1, leaving only χ_2 to be estimated.

The log-linearized equilibrium conditions for this model can then be summarized as:

$$\hat{Y}_t = \hat{u}_t + \mu \hat{K}_{t-1} + (1-\mu)\left(\alpha \hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right), \tag{C.24}$$

$$\hat{r}_t^k = \hat{u}_t + (\mu - 1)\,\hat{K}_{t-1} + (1-\mu)\left(\alpha \hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right), \tag{C.25}$$

$$\hat{w}_t^b = \hat{u}_t + \mu \hat{K}_{t-1} + (1-\mu)\left(\left(\alpha - \frac{1}{(1-\mu)}\right)\hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right), \tag{C.26}$$

$$\hat{w}_t^l = \hat{u}_t + \mu \hat{K}_{t-1} + (1-\mu)\left(\alpha \hat{N}_t^b + \left(1 - \alpha - \frac{1}{(1-\mu)}\right)\hat{N}_t^l\right), \tag{C.27}$$

$$\frac{\chi_2}{\chi_1}\hat{u}_t = \hat{Y}_t - \hat{u}_t, \tag{C.28}$$

where the last equation is the log-linearized version of the optimality condition for u_t. Total factor productivity in this model is simply given by the utilization rate of technology:

$$T\hat{F}P_t = \hat{u}_t. \tag{C.29}$$

C.4 Love of variety and endogenous firm turnover

For this model, we consider each layer of production in turn.

C.4.1 Final good firms

The representative final good firm maximizes profits:

$$P_t Y_t - \int_0^1 Q_t(j) p_t(j) dj, \tag{C.30}$$

subject to the production technologies

$$Y_t = \left[\int_0^1 Q_t(j)^{\omega} dj\right]^{\frac{1}{\omega}}, \tag{C.31}$$

$$Q_t(j) = F_t(j)^{\tau + \frac{\rho-1}{\rho}} \left[\sum_{i=1}^{F_t(j)} m_t(j,i)^{\rho}\right]^{\frac{1}{\rho}}. \tag{C.32}$$

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

$$\min_{\{Q_t(j)\}_{j=0}^1} \int_0^1 Q_t(j) p_t(j) dj \qquad \text{subject to } Y_t = \left[\int_0^1 Q_t(j)^{\omega} dj\right]^{\frac{1}{\omega}}. \tag{C.33}$$

This leads to the standard demand function and price index:

$$Q_t(j) = \left(\frac{p_t(j)}{P_t}\right)^{\frac{1}{\omega-1}} Y_t, \tag{C.34}$$

$$P_t = \left[\int_0^1 p_t(j)^{\frac{\omega}{\omega-1}} dj\right]^{\frac{\omega-1}{\omega}}. \tag{C.35}$$

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

$$\min_{\{m_t(j,i)\}_{i=1}^{F_t(j)}} \sum_{i=1}^{F_t(j)} p_t(j,i)m_t(j,i) \qquad \text{s.t. } Q_t(j) = F_t(j)^{\tau+\frac{\rho-1}{\rho}} \left[\sum_{i=1}^{F_t(j)} m_t(j,i)^{\rho}\right]^{\frac{1}{\rho}}, \tag{C.36}$$

which has the first-order condition

$$p_t(j,i) - p_t(j)\frac{1}{\rho}F_t(j)^{\tau+\frac{\rho-1}{\rho}} \left[\sum_{i=1}^{F_t(j)} m_t(j,i)^{\rho}\right]^{\frac{1}{\rho}-1} \rho m_t(j,i)^{\rho-1} = 0. \tag{C.37}$$

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

$$m_t(j,i) = \left(\frac{p_t(j,i)}{p_t(j)}\right)^{\frac{1}{\rho-1}} \frac{Q_t(j)}{\left(F_t(j)^{\tau+\frac{\rho-1}{\rho}}\right)^{\frac{\rho}{\rho-1}}} = \left(\frac{p_t(j,i)}{p_t(j)}\right)^{\frac{1}{\rho-1}} \left(\frac{p_t(j)}{P_t}\right)^{\frac{1}{\omega-1}} \frac{Y_t}{\left(F_t(j)^{\tau+\frac{\rho-1}{\rho}}\right)^{\frac{\rho}{\rho-1}}}. \tag{C.38}$$

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

$$p_t(j) = \frac{1}{F_t(j)^{\tau+\frac{\rho-1}{\rho}}} \left[\sum_{i=1}^{F_t(j)} p_t(j,i)^{\frac{\rho}{\rho-1}}\right]^{\frac{\rho-1}{\rho}}. \tag{C.39}$$

C.4.2 Intermediate goods firms

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

$$\frac{p_t(j,i)}{P_t} m_t(j,i) - w_t^l n_t^l(j,i) - w_t^b n_t^b(j,i) - r_t^k k_{t-1}(j,i), \tag{C.40}$$

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

$$m_t(j,i) = k_{t-1}(j,i)^{\mu}\left[n_t^b(j,i)^{\alpha} n_t^l(j,i)^{1-\alpha}\right]^{1-\mu} - \varphi, \tag{C.41}$$

$$m_t(j,i) = \left(\frac{p_t(j,i)}{p_t(j)}\right)^{\frac{1}{\rho-1}}\left(\frac{p_t(j)}{P_t}\right)^{\frac{1}{\omega-1}}\frac{Y_t}{\left(F_t(j)^{\tau+\frac{\rho-1}{\rho}}\right)^{\frac{\rho}{\rho-1}}}, \tag{C.42}$$

$$p_t(j) = \frac{1}{F_t(j)^{\tau+\frac{\rho-1}{\rho}}}\left[\sum_{i=1}^{F_t(j)} p_t(j,i)^{\frac{\rho}{\rho-1}}\right]^{\frac{\rho-1}{\rho}}. \tag{C.43}$$

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

$$r_t^k = \mu\frac{p_t(j,i)}{P_t}\frac{k_{t-1}(j,i)^{\mu}\left[n_t^b(j,i)^{\alpha} n_t^l(j,i)^{1-\alpha}\right]^{1-\mu}}{x_t(j,i)k_{t-1}(j,i)}, \tag{C.44}$$

$$w_t^b = (1-\mu)\alpha\frac{p_t(j,i)}{P_t}\frac{k_{t-1}(j,i)^{\mu}\left[n_t^b(j,i)^{\alpha} n_t^l(j,i)^{1-\alpha}\right]^{1-\mu}}{x_t(j,i)n_t^b(j,i)}, \tag{C.45}$$

$$w_t^l = (1-\mu)(1-\alpha)\frac{p_t(j,i)}{P_t}\frac{k_{t-1}(j,i)^{\mu}\left[n_t^b(j,i)^{\alpha} n_t^l(j,i)^{1-\alpha}\right]^{1-\mu}}{x_t(j,i)n_t^l(j,i)}. \tag{C.46}$$

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

$$\varepsilon_{m_t(j,i)} = \left(\frac{m_t(j,i)}{p_t(j,i)}\frac{1}{\rho-1} + \left(\frac{1}{\omega-1} - \frac{1}{\rho-1}\right)\frac{m_t(j,i)}{p_t(j)}\frac{\rho-1}{\rho}\frac{p_t(j)}{\sum_{i=1}^{F_t} p_t(j,i)^{\frac{\rho}{\rho-1}}}\frac{\rho}{\rho-1}p_t(j,i)^{\frac{\rho}{\rho-1}-1}\right)\frac{p_t(j,i)}{m_t(j,i)}. \tag{C.47}$$

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

$$\varepsilon_{m_t(j,i)} = \frac{1}{\rho-1} + \left(\frac{1}{\omega-1} - \frac{1}{\rho-1}\right)\left(\frac{p_t(j,i)}{p_t(j)F_t(j)^{\tau}}\right)^{\frac{\rho}{\rho-1}}\frac{1}{F_t}. \tag{C.48}$$

Since the firm sells the good in a monopolistically 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:

$$x_t(j,i) = \frac{1}{1+\frac{1}{\varepsilon_{m_t(j,i)}}} = \frac{\varepsilon_{m_t(j,i)}}{1+\varepsilon_{m_t(j,i)}}. \tag{C.49}$$

Combining this expression with identical price setting returns the markup as a decreasing function of the number of firms:

$$x_t = \frac{(1-\omega)\,F_t - (\rho-\omega)}{\rho\,(1-\omega)\,F_t - (\rho-\omega)}. \tag{C.50}$$

The steady-state markup can be derived directly from (C.50):

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

C.4.3 Linearized symmetric firm equilibrium conditions

The log-linearized factor prices read as

$$\hat{r}_t^k = (1+\tau)\left(\left(\left(\mu - \frac{1}{1+\tau}\right)\hat{K}_{t-1} + (1-\mu)\left(\alpha\hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t, \tag{C.51}$$

$$\hat{w}_t^b = (1+\tau)\left(\left(\mu\hat{K}_{t-1} + (1-\mu)\left(\left(\alpha - \frac{1}{(1+\tau)(1-\mu)}\right)\hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t, \tag{C.52}$$

$$\hat{w}_t^l = (1+\tau)\left(\left(\mu\hat{K}_{t-1} + (1-\mu)\left(\alpha\hat{N}_t^b + \left(1-\alpha - \frac{1}{(1+\tau)(1-\mu)}\right)\hat{N}_t^l\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t. \tag{C.53}$$

while log-linearization of (4.6)—upon inserting from (5.20)—results in

$$\hat{Y}_t = (1+\tau)\left(\left(\mu\hat{K}_{t-1} + (1-\mu)\left(\alpha\hat{N}_t^b + (1-\alpha)\hat{N}_t^l\right)\right)\right) - \frac{x-(1+\tau)}{x-1}\hat{x}_t. \tag{C.54}$$

We can combine the production function (5.12) with (5.17) 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 (4.6) and (5.20), 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

$$\hat{F}_t = \frac{1}{1+\tau}\left(\hat{Y}_t + \frac{x}{x-1}\hat{x}_t\right). \tag{C.55}$$

We rewrite the markup (C.50) as

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

which is log-linearized as

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

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

$$\hat{F}_t = \frac{x}{x-1}\frac{\rho-1}{\rho x-1}\hat{x}_t. \tag{C.56}$$

The log-linearized expression for TFP is

$$T\hat{F}P_t = \tau\hat{F}_t - \hat{x}_t. \tag{C.57}$$

D A stylized version of the LOV 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.32), (B.39), (C.53), (C.54), (C.55), and (C.56), respectively:

$$\hat{w}_t = \psi\hat{n}_t + \hat{c}_t, \tag{D.1}$$

$$\hat{y}_t = (1-\theta)\hat{c}_t + \theta\hat{g}_t, \tag{D.2}$$

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

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

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

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

We now solve for the response of output to a government spending shock. First, combine (D.1) and (D.2) to eliminate $\hat{c}_t$. Then plug (D.3) in the resulting equation and rearrange to obtain

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

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

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

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

$$\hat{y}_t = \frac{\gamma x (1+\tau)\left[(\rho x - 1) - (\rho - 1)(1+\tau)\right]}{x\left[(\tau - \psi)(1-\gamma) - (1+\tau)\right]\left[(\rho - 1)(1+\tau) - (\rho x - 1)\right] + (\rho x - 1)(1-\gamma)\left[\tau - \psi - (1+\tau)\right]\left[x - (1+\tau)\right]}\hat{g}_t. \tag{D.9}$$

This expression allows us to analyze the importance of the different effects at play in the model. The first step consists of evaluating the role of endogenous entry in isolation. To this end, we set $\tau = 0$, so that (D.9) reduces to:

$$\hat{y}_t = \frac{\theta \rho x}{x\rho\left[1 + \psi(1-\theta)\right] - (\rho x - 1)(1-\theta)(1+\psi)}\hat{g}_t. \tag{D.10}$$

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

$$\rho x > 1 + \psi. \tag{D.11}$$

Since 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 substituting 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 (D.11) embodies the problems encountered in the existing literature when trying to generate consumption crowding-in through endogenous firm entry: Conditional on a realistic value of ρ, the condition can only be satisfied for unconventionally high values of the markup, x, consistent with the numerical results of Devereux et al. (1996); or for values of the Frisch elasticity, $\frac{1}{\psi}$, that are inconsistent with microeconometric studies, as discussed by Bilbiie (2011). A similar point was made by Lewis and Winkler (2017).

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

$$\tau > \frac{(\rho x - 1 - \psi)(1 - x)}{x(1-\rho)}. \tag{D.12}$$

Notice how this condition embeds (D.11): As long as this sufficient condition is satisfied, (D.12) always holds, as the overall expression on its right-hand 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, for given (realistic) values of the elasticity of labor supply and the markup.

It is important to stress that (D.12) is uninformative about the quantitative sensitivity of the house price to a fiscal shock, and how the variety and the competition effects combine, in this respect. To address this point, it is instructive to examine the log-linearized expression of TFP, whose reaction is key to limiting the negative wealth effect induced by a fiscal expansion:

$$T\hat{F}P_t = \tau \hat{F}_t - \hat{x}_t. \tag{D.13}$$

In the absence of a variety effect ($\tau = 0$), the model can only rely on a large variation in the markup to produce large shifts in TFP. This can be obtained through a high steady-state markup, x, which is key to index the comovement between the number of firms and the markup (i.e., the competition effect): When the steady-state markup is relatively high, the economy is characterized by poor competition and few firms—each with substantial market power—while fixed costs are high. Consequently, a marginal entrant has a rather large effect on the degree of competition, and thus on the response of the markup to a fiscal shock. However, as discussed above, from a quantitative viewpoint the steady-state markup that is necessary to attain sizable conditional variation in TFP might be counterfactually high. On the other hand, if the steady-state markup is relatively low—and the economy tends towards a perfect-competition benchmark—the marginal effect of an additional entrant is limited, so that complementing the competition effect with some taste for variety (i.e., $\tau > 0$) will be necessary to directly channel changes in the number of firms into the TFP response, all else equal. These considerations will be important for the quantitative implications of the LOV model.

E Additional numerical results

We now report some further computational findings from the quantitative analysis. We begin by reporting the parameter estimates of the standard RBC model considered in Figure 5 in the main text, i.e., a model without any of the endogenous productivity mechanisms we have considered. The parameter estimates are collected in Table E.1. Next, we report the results from the estimation of a set of models with some plausible parameter restrictions imposed. These are collected in Table E.2. Finally, we return to the model versions considered in Section 7.2, where we allow the equilibrium house price to be determined by credit-constrained borrowers, instead of by unconstrained lenders. Since the responses of most variables are quite similar to those obtained in the baseline model, we do not report all of them. Instead, we choose to focus on the responses of the house price (reported in Figure 7 in the main text) and of mortgage debt, which we report in Figure E.1. As seen from the figure, the increase in mortgage debt in each of these models falls short of that observed in the data by a significant amount, reflecting the fact that the borrowers are now unable to increase their stock of housing in response to shocks.

Table E.1: Estimated parameter values: RBC model

Parameter	Description	Estimate
σ_c	Curvature in utility of C	4.998 [4.775−5.000]
σ_h	Curvature in utility of H	0.912 [0.315−1.633]
h^l	Habit formation, lenders	0.747 [0.570−0.900]
h^b	Habit formation, borrowers	0.743 [0.306−0.900]
ψ	Inverse Frisch elasticity	0.250 [0.250−0.264]
ϕ	Capital adjustment costs	24.770 [10.046−24.997]
γ	Inertia of mortgage debt	0.950 [0.935−0.950]
γ_τ	Tax response to gov't debt	0.816 [0.197−0.894]
ρ_τ	Inertia of tax rate	0.337 [0.101−0.900]
γ_G	Persistence of G shock	0.958 [0.922−0.968]
σ_g	Std. dev. of G shock	0.096 [0.082−0.107]

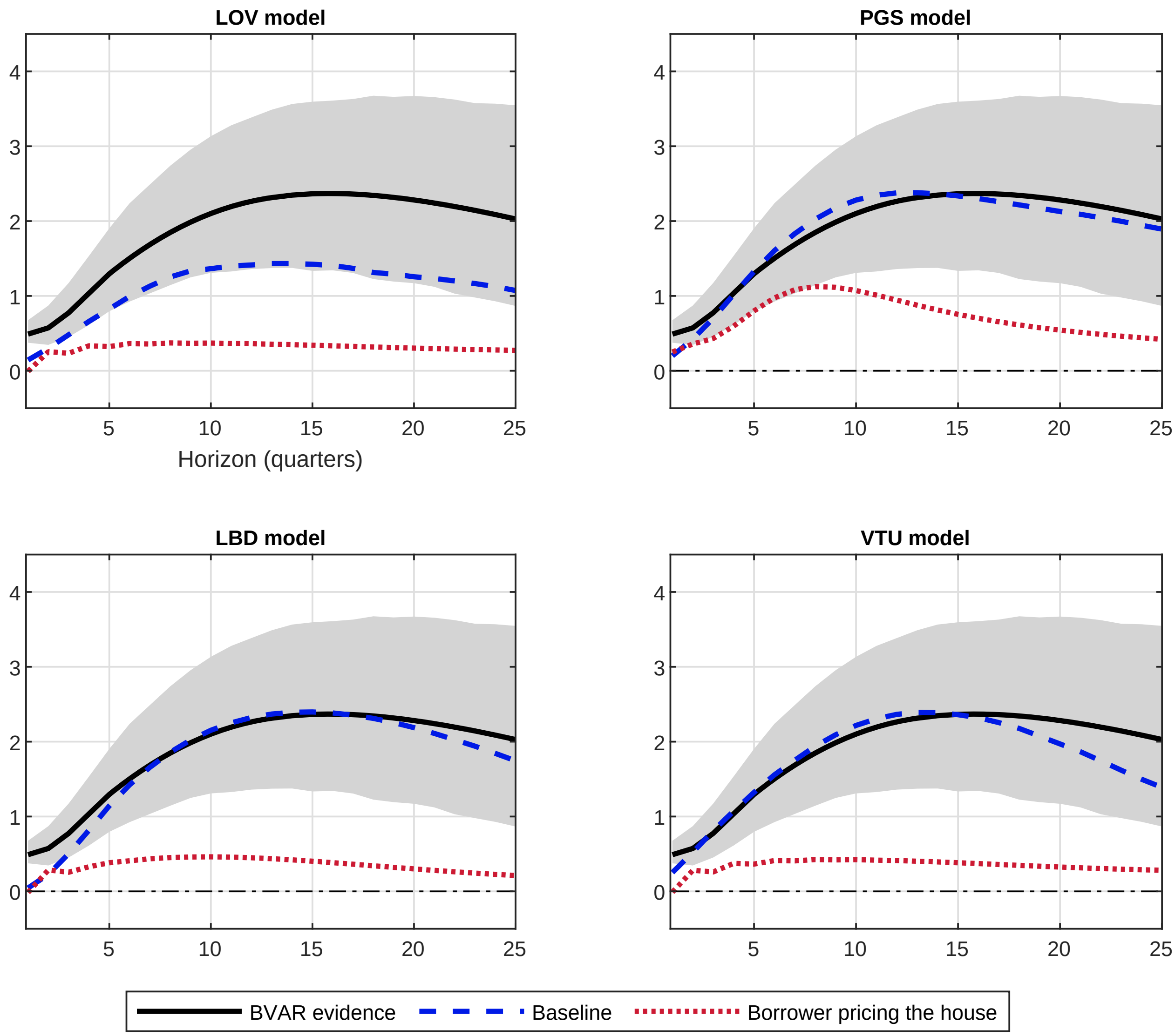
Notes: The table reports the median parameter estimates from the model, with the 68 percent credible sets (i.e., the 16th and 84th percentiles) reported in brackets.

Table E.2: Estimated parameter values: Models with parameter restrictions

Parameter	Description	LOV model	PGS model	LBD model	VTU model
σ_c	Curvature in utility of C	0.545 [0.510−3.045]	1.508 [0.422−2.734]	4.683 [4.374−4.944]	2.083 [1.220−4.621]
σ_h	Curvature in utility of H	1.410 [0.454−1.657]	0.421 [0.137−4.197]	0.306 [0.276−0.976]	0.159 [0.100−0.778]
h^l	Habit formation, lenders	0.415 [0.271−0.450]	0.712 [0.461−0.844]	0.755 [0.580−0.839]	0.390 [0.000−0.668]
h^b	Habit formation, borrowers	0.367 [0.312−0.690]	0.771 [0.501−0.894]	0.372 [0.251−0.621]	0.683 [0.343−0.742]
ψ	Inverse Frisch elasticity	0.304 [0.279−0.378]	0.279 [0.251−0.625]	0.260 [0.251−0.270]	0.250 [0.250−0.307]
ϕ	Capital adjustment costs	10.013 [2.515−10.100]	10.681 [9.699−11.747]	11.541 [10.059−22.641]	0.002 [0.000−10.911]
γ	Inertia of mortgage debt	0.577 [0.465−0.884]	0.892 [0.767−0.948]	0.938 [0.921−0.950]	0.880 [0.673−0.950]
γ_τ	Tax response to gov't debt	0.650 [0.125−0.708]	0.588 [0.072−0.750]	0.819 [0.492−0.876]	0.899 [0.216−0.900]
ρ_τ	Inertia of tax rate	0.481 [0.441−0.512]	0.484 [0.309−0.766]	0.313 [0.246−0.842]	0.233 [0.100−0.730]
γ_G	Persistence of G shock	0.942 [0.921−0.966]	0.955 [0.905−0.980]	0.974 [0.932−0.984]	0.990 [0.929−0.990]
σ_g	Std. dev. of G shock	0.106 [0.082−0.120]	0.109 [0.085−0.124]	0.099 [0.082−0.118]	0.098 [0.082−0.124]
τ	Love for variety	1.867 [1.777−1.941]	*N/A*	*N/A*	*N/A*
x	Steady-state value of markup	1.130 [1.121−1.141]	*N/A*	*N/A*	*N/A*
μ_g	Productivity of public capital	*N/A*	0.396 [0.305−0.400]	*N/A*	*N/A*
δ_g	Depreciation of public capital	*N/A*	0.098 [0.093−0.100]	*N/A*	*N/A*
θ_n	Skill acquisition from working	*N/A*	*N/A*	0.109 [0.104−0.111]	*N/A*
ρ_x	Persistence of acquired skills	*N/A*	*N/A*	0.793 [0.769−0.796]	*N/A*
χ_2	Tech. utilization adj. cost	*N/A*	*N/A*	*N/A*	1.500 [1.500−1.589]

Notes: The table reports the median parameter estimates from each model, with the 68 percent credible sets (i.e., the 16th and 84th percentiles) reported in brackets.

Figure E.1: Effects of a government spending shock on mortgage debt: The role of credit frictions

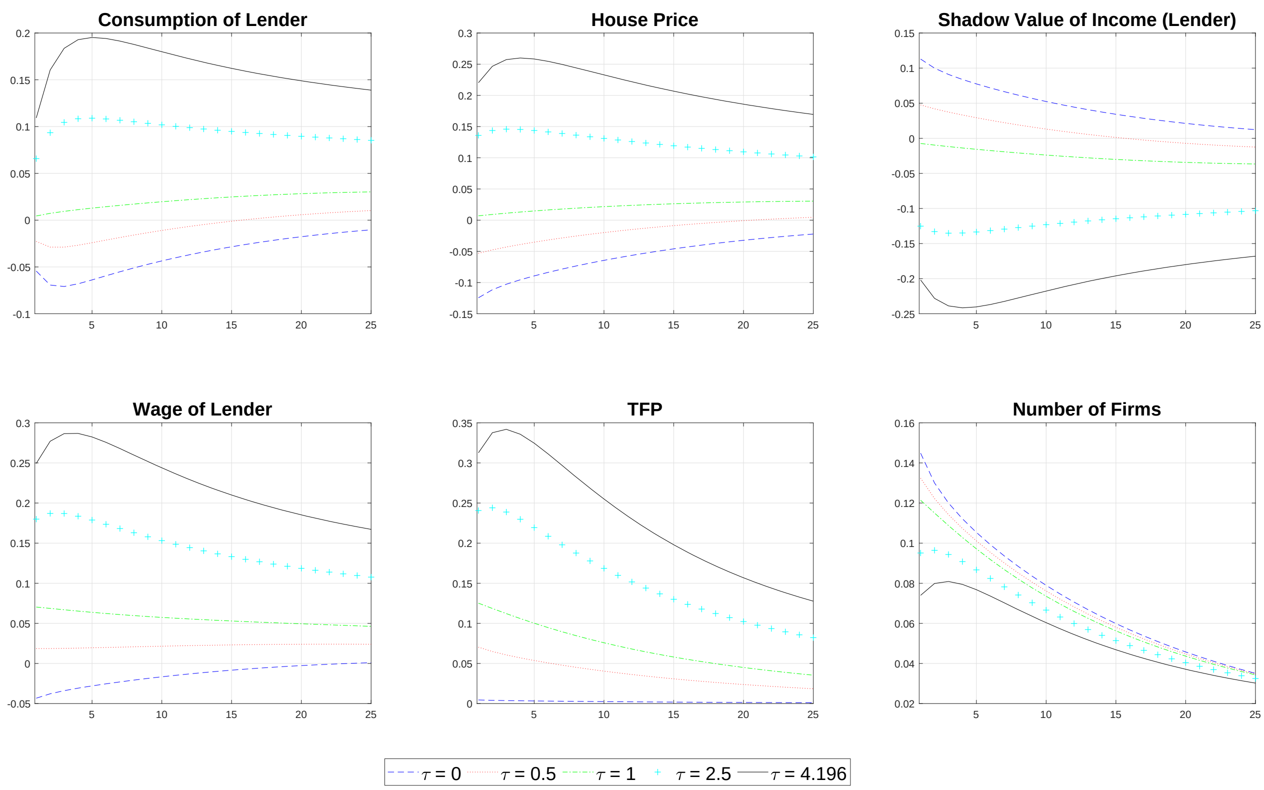


Notes: The figure shows the mortgage debt response to a government spending shock under different assumptions regarding which agent is determining the house price in equilibrium. Solid black line: BVAR model. Grey areas: 68 percent credible sets from BVAR model. Dashed blue line: Baseline model with lender pricing the house. Dotted red line: Alternative model with constrained borrower pricing the house.

F On the role of the competition and variety effects in the LOV model

Figure F.1 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 and 2).[33] As the figure illustrates, the model is capable of producing a negative response of the shadow value of lenders' income, and thus an increase in the house price, provided that the variety effect is strong enough, that is, that τ is high enough. Given the other model parameters, values of $\tau > 1$ are sufficient for this to happen. In these cases, the increase in TFP drives up the wage rate, allowing lenders to increase their consumption. Note that the response of the number of firms is more modest when τ is high. This reflects that the model appears to favour the variety effect over the competition effect in order to generate a sizeable increase in TFP.

Figure F.1: Effects of a government spending shock for different values of τ



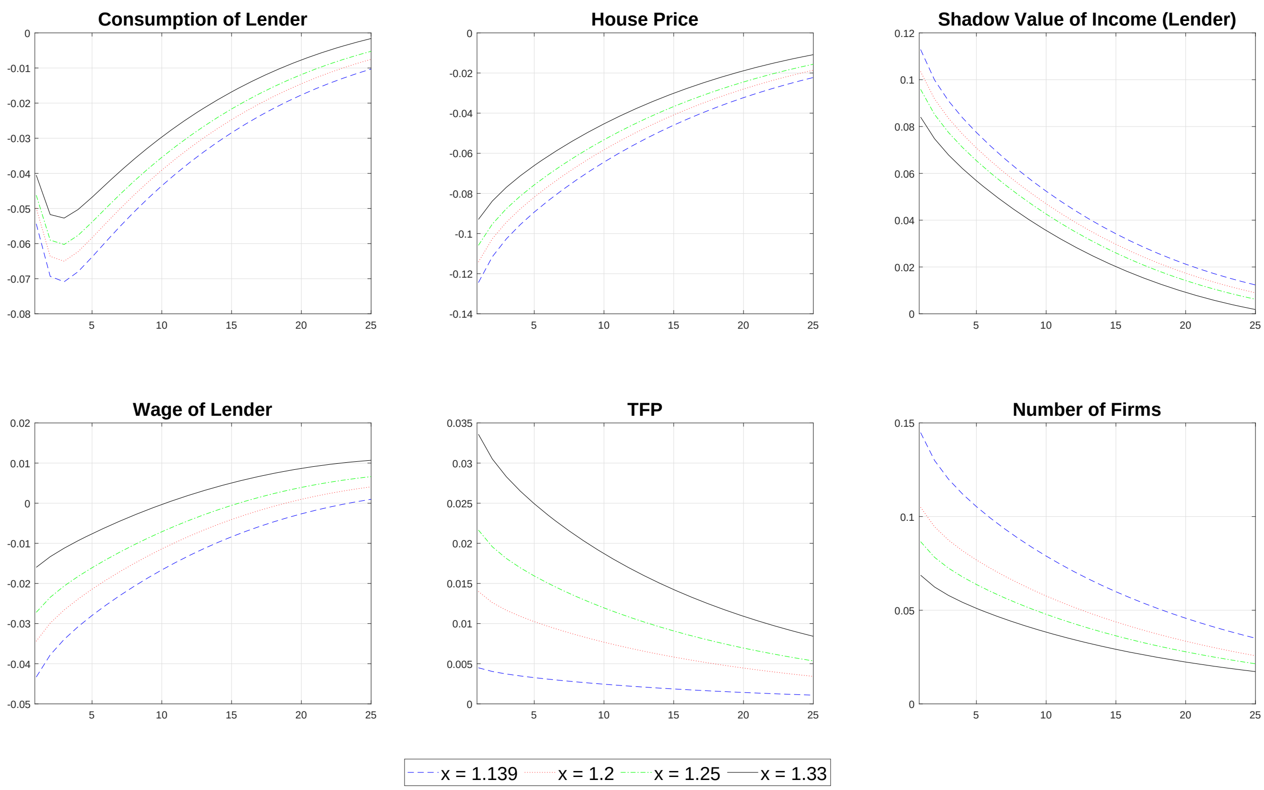
Notes: The figure shows the effects of a shock to government spending for various values of the love-of-variety parameter τ. Dashed blue line: $\tau = 0$. Dotted red line: $\tau = 0.5$. Dashed-dotted green line: $\tau = 1$. Crossed cyan line: $\tau = 2.5$. Solid black line: $\tau = 4.196$ (estimated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

To dig deeper into the endogenous drivers of the model, we find it useful to consider Figure

[33]Note that the impulse responses in Figures F.1 and F.2 do not correspond exactly to those in Figure 5. The reason is that the former display impulse responses based on the median parameter estimates reported in Table 2, while the latter displays the median impulse responses. As is well known, these two objects do not necessarily coincide. This, however, does not affect any of our qualitative conclusions.

F.2, 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 $TFP_t = 1/x_t$. However, even a relatively high value of x, which corresponds to a rather strong competition effect—recall that the marginal entrant has higher chances to produce a sizable impact on the markup in a relatively concentrated market—is not enough to produce a drop in the markup and, thus, a rise in TFP that is large enough.

Figure F.2: Government spending shock for different values of x without variety effect ($\tau = 0$)



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

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 Appendix D, alongside the indirect effect through the markup discussed above. The variety effect relies on a large increase in the number of firms to produce the maximal impact on TFP. This explains why the estimation of our baseline model returns a low value of the steady-state markup: The estimation procedure prefers an environment with strong competition and low entry costs, so that a fiscal shock leads to a large increase in the number of operating firms. This is true despite the fact that a low steady-state markup entails a rather weak competition effect, implying that TFP is only affected by a small decline in the markup.