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Takushi Kurozumi*
takushi.kurozumi@boj.or.jp

Willem Van Zandweghe**
willem.vanzandweghe@clev.frb.org

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2-1-1 Nihonbashi-Hongokucho, Chuo-ku, Tokyo 103-0021, Japan

* Monetary Affairs Department, Bank of Japan
** Research Department, Federal Reserve Bank of Cleveland

Macroeconomic Changes with Declining Trend Inflation: Complementarity with the Superstar Firm Hypothesis*

Takushi Kurozumi† Willem Van Zandweghe‡

Abstract

Recent studies indicate that, since 1980, the US economy has undergone increases in the average markup and the profit share of income and decreases in the labor share and the investment share of spending. We examine the role of monetary policy in these changes as inflation has concurrently trended down. In a simple staggered price model with a non-CES aggregator of individual differentiated goods, a decline of trend inflation as measured since 1980 can account for a substantial portion of the changes. Moreover, adding a rise of highly productive "superstar firms" to the model can better explain not only the macroeconomic changes but also the micro evidence on the distribution of firms' markups, including the flat median markup.

JEL Classification: E52, L16

Keywords: Average markup; Profit share; Labor share; Trend inflation; Non-CES aggregator; Superstar firm hypothesis

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†Monetary Affairs Department, Bank of Japan, 2-1-1 Nihonbashi-Hongokucho, Chuo-ku, Tokyo 103-8660, Japan (e-mail: takushi.kurozumi@boj.or.jp)

‡Research Department, Federal Reserve Bank of Cleveland, 1455 E 6th St, Cleveland, OH 44114, USA (e-mail: willem.vanzandweghe@clev.frb.org)

1 Introduction

Recent studies have documented profound changes to the US economy since 1980. The average price–cost markup has increased, as illustrated by the solid line in panel (a) of Figure 1.[1] Concurrently, the profit share of income has increased, while the labor share (of income) and the investment share of spending have decreased, as shown by the solid lines in panels (b), (c), and (d) of the figure, respectively.[2] In each of the four panels a trend is plotted by the dashed regression line, whose slope is significantly different from zero.

A growing literature has studied possible linkages between the US macroeconomic changes. Barkai (2020) analyzes the link between the increase in the average markup and the decrease in the labor share using aggregate data, and shows that the decrease has been more than offset by a rise in pure profits, thus suggesting that the decreasing labor share is due to a decline in competition.[3] Autor et al. (2020) use micro panel data to address the link, and find evidence consistent with a rise since the early 1980s of highly productive "superstar firms," which have increased concentration and markups and reduced the labor share in their industries.[4] In this context, De Loecker et al. (2020) investigate firm-level data and indicate that the increase in the average markup is driven mainly by the upper tail of the distribution of firms' markups. The link between higher markups and the lower investment share is investigated by Covarrubias et al. (2020), who document a decrease in business investment during the last two decades and attribute it to a rise in concentration and a decline in compe-

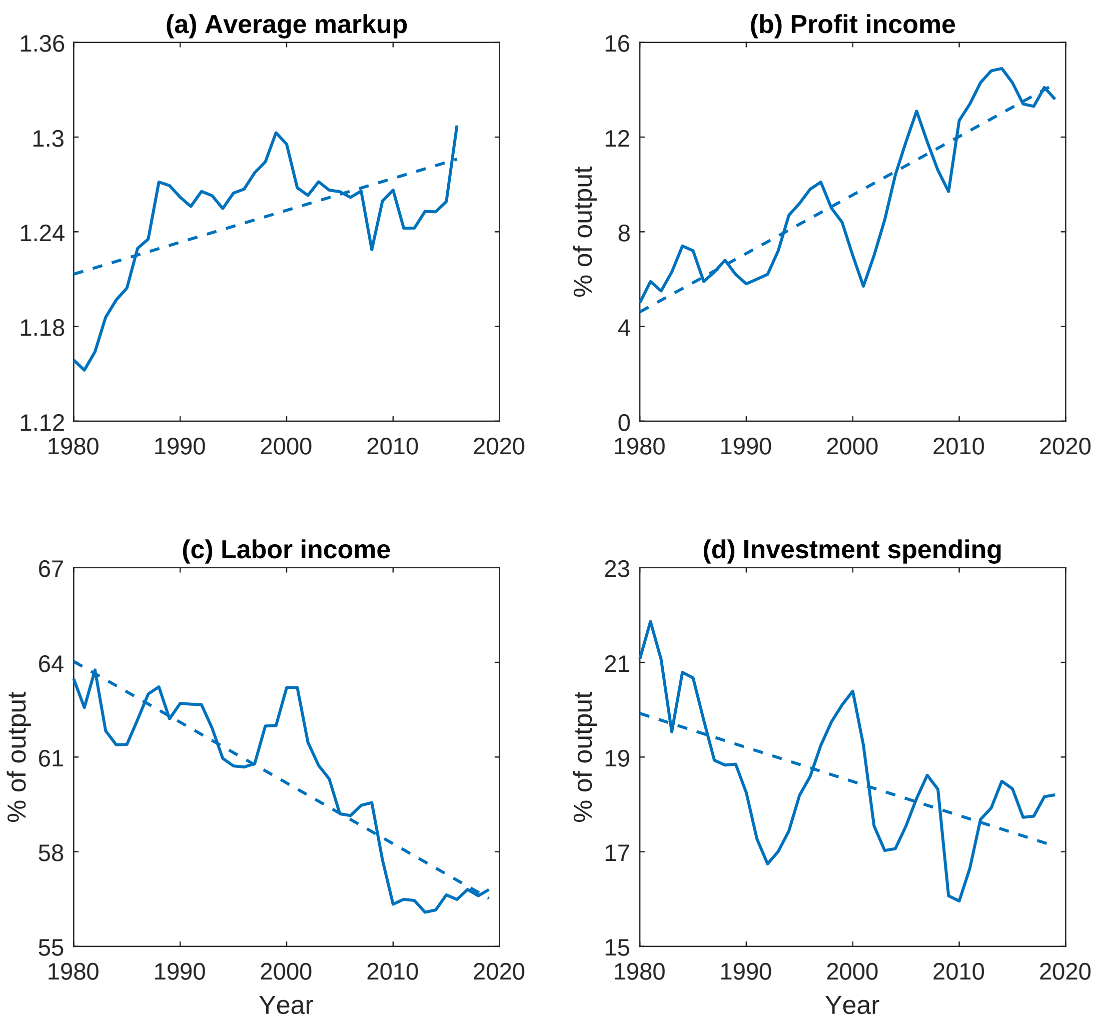
[1] Our measure of the average markup is the *sales-weighted harmonic* average markup, which coincides with the *cost-weighted arithmetic* average markup. De Loecker et al. (2020) employ the *sales-weighted arithmetic* average markup as displayed in their Figure I, and point out that such an average markup measure exhibits a larger increase, to 1.6 in 2016, because firms with higher markups tend to have higher sales weights relative to their cost weights. Edmond et al. (2021) indicate that the cost-weighted arithmetic average markup is the relevant statistic that summarizes the distortions to employment and investment decisions, and they show an average markup series similar to ours.

[2] Akcigit and Ates (2021) highlight 10 facts on declining business dynamism in the US since 1980, including the increases in the average markup and the profit share and the decrease in the labor share.

[3] Barkai (2020) calculates that the profit share of income rose by 13.5 percentage points during the period 1984–2014, a larger rise than that seen in panel (b) of Figure 1 (i.e., 7.5 percentage points during the same period). Karabarbounis and Neiman (2019) cast doubt on the rising profit share.

[4] See also Elsby et al. (2013), Karabarbounis and Neiman (2014), and Kehrig and Vincent (2021) for analyses of the decreasing labor share. Aum and Shin (2020) note that the decline in the labor share accelerated after 2000, and point to increased capital–labor substitution in the services sector as a possible factor behind the acceleration.

Figure 1: Evolution of key US macroeconomic variables.



Notes: In the figure the solid lines illustrate key US macroeconomic time series from 1980 to 2019 or the most recent available year and the dashed lines display their respective regression trends. Panel (a) shows the sales-weighted harmonic average markup, which coincides with the cost-weighted arithmetic average markup. Panel (b) exhibits corporate profits adjusted for inventory valuation and capital consumption as a share of value added of the nonfinancial corporate sector. Panel (c) plots the labor share in the nonfarm business sector. Panel (d) presents the share of business fixed investment in spending, where spending is measured as the sum of business fixed investment, personal consumption expenditures (PCE) for nondurable goods, and PCE for services.

Sources: De Loecker et al. (2020), US Bureau of Labor Statistics, US Bureau of Economic Analysis, Haver Analytics.

tition.[5] As for the forces driving some of the US macroeconomic changes, the literature has pointed to globalization and technological changes (Autor et al., 2020), weakened antitrust enforcement (Gutiérrez and Philippon, 2017), and patent concentration (Akcigit and Ates, 2021).

This paper examines the role of monetary policy as a driving factor behind the US macroeconomic changes. In tandem with the changes, Figure 2 shows that inflation has steadily trended down since 1980, a trend that is well-known but has hitherto not been linked to the changes illustrated in Figure 1. The dashed line in Figure 2 displays the estimated trend inflation series of Chan et al. (2018).[6] The series of the (trend) inflation rate and the average markup exhibit a negative association, suggesting the possibility of a structural relationship between the two macroeconomic variables.[7] Under lower trend inflation, firms' price–cost markups are less severely eroded in between infrequent price adjustments, which may raise the average markup. To see some empirical support for this, we conduct an industry-level regression analysis of the change in the average markup from 1980 to 2016 on the frequency of price adjustment, and find evidence that the change in the average markup under the concurrent decline in trend inflation was larger for industries that adjust prices less frequently.

To investigate the implications of declining trend inflation for the increasing average markup and the three other macroeconomic changes, we use a simple staggered price model.[8] A key feature of the model is that, each period, a fraction of individual goods' prices remains unchanged in line with micro evidence, while the other prices are set given demand curves arising from a not necessarily constant elasticity of substitution (CES) aggregator of individual differentiated goods of the sort proposed by Kimball (1995) and developed by

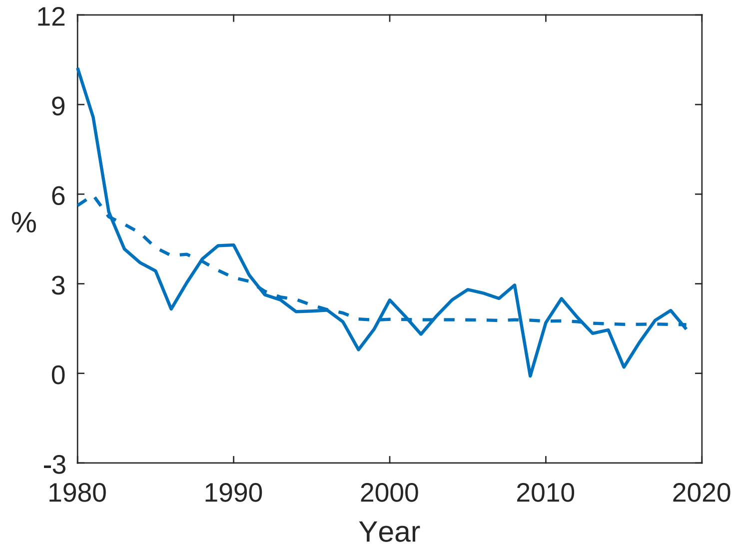
[5] Basu (2019) and Syverson (2019) provide excellent reviews of related studies.

[6] See, e.g., Ireland (2007), Cogley and Sbordone (2008), and Ascari and Sbordone (2014) for alternative estimates of trend inflation. The estimated series of Chan et al. (2018) exhibits a decline similar to those of the previous studies.

[7] The negative association between inflation and markups is consistent with the findings of empirical studies, such as Bénabou (1992) and Banerjee and Russell (2001, 2005).

[8] One may wonder how our paper treats the natural rate hypothesis. Despite the widely held view going back to Friedman (1968), the empirical evidence on whether monetary policy has long-run real effects is not as clear-cut. Various recent empirical research has supported the notion that monetary policy can have long-lasting real effects. For example, Moran and Queralto (2018) demonstrate increases in R&D and medium-run TFP following an expansionary monetary policy shock. Jordà et al. (2020) show that the effects of monetary policy on TFP, capital accumulation, and the production capacity of the economy are long lived.

Figure 2: Evolution of US inflation.



Notes: In the figure the solid and the dashed lines display the inflation rate of the personal consumption expenditures price index and its estimated trend of Chan et al. (2018) from 1980 to 2019, respectively.
Sources: Chan et al. (2018), US Bureau of Economic Analysis, Haver Analytics.

Dotsey and King (2005) and Levin et al. (2008), which includes the CES aggregator as a special case. The non-CES aggregator provides a parsimonious way of introducing variable price elasticity of demand and hence endogenous desired markups of firms. Specifically, for a lower (higher) relative price of a good, the price elasticity of demand for the good decreases (increases), which raises (reduces) the firm's desired markup and thus inhibits the firm from setting such a relative price.[9]

The calibrated model shows that a decline of trend inflation as measured since 1980 can account for a substantial portion of the US macroeconomic changes. The model attributes around 30% of the increases in the average markup and the profit share and the decrease in the labor share and about 20% of the decrease in the investment share to the decline in trend inflation. The endogenous desired markups arising from the non-CES aggregator make price-adjusting firms' markups less responsive to trend inflation, because a lower relative price induced by a decline in trend inflation raises the firms' desired markups. At the same time, the decline in trend inflation leads non-adjusting firms to experience a less severe erosion

[9]Because of this feature, the non-CES aggregator has been widely used as a source of real rigidity in the macroeconomic literature. See, e.g., Eichenbaum and Fisher (2007), Smets and Wouters (2007), Gopinath and Itskhoki (2011), and Kurozumi and Van Zandweghe (2016).

of their markups. This increases the average markup, which in turn raises the profit share and reduces the labor share and the investment share in the model. This result contrasts with that in the case of the CES aggregator in which the average markup is almost flat for the decline in trend inflation and thus keeps the other macroeconomic variables almost unchanged.

To reconcile our analysis with the recent literature that has advocated the superstar firm hypothesis as a leading explanation of the increasing average markup, the model is extended by introducing such firms, which are distinguished from ordinary firms by their higher productivity. The non-CES aggregator plays a dual role in the extended model. One role is acting as a source of endogenous variation in desired markups as in the baseline model. The other is serving as a source of markup heterogeneity between firms with different productivity levels, since the aggregator implies that more productive firms face less price-elastic demand and thus choose higher markups when they can adjust prices.[10] In the calibrated model, adding a rise of superstar firms increases the average markup further, by raising the upper tail of the distribution of firms' markups in line with the micro evidence (e.g., De Loecker et al., 2020), and therefore it can better explain the US macroeconomic changes. Moreover, the rise of superstar firms and the decline in trend inflation have offsetting effects on the median markup, thus keeping it almost unchanged. The flat median markup is consistent with the micro evidence reported by Autor et al. (2020) and De Loecker et al. (2020) for the period since the early 1980s. In this way, the decline in trend inflation complements the rise of superstar firms in accounting for the empirical changes in the macroeconomic variables and the distribution of firms' markups.

The remainder of the paper proceeds as follows. Section 2 reviews existing and new evidence on the relationship between inflation and markups. Section 3 presents a simple staggered price model with trend inflation and a Kimball-type aggregator. Section 4 investigates the implications of declining trend inflation for the average markup, the profit share of income, the labor share, and the investment share of spending in the model. Section 5 extends the analysis by adding a rise of superstar firms to the model. Section 6 concludes.

[10] See, e.g., Autor et al. (2020) and Edmond et al. (2021) for the use of non-CES aggregators as a source of markup heterogeneity between firms.

2 Empirical Evidence

This section summarizes existing macroeconomic evidence of a negative association between inflation and markups, and presents new industry evidence supporting the view that lower inflation has contributed to the increase in the average markup.

Macroeconomic time series evidence indicates that higher inflation is associated with lower markups. Since 1980, the first differences of the average markup series displayed in panel (a) of Figure 1 and the trend inflation series plotted in Figure 2 have a statistically significant negative correlation of -0.33, which survives for the longer period from 1962 (-0.39). The negative correlation is consistent with the findings of empirical studies. Bénabou (1992) shows that anticipated and unanticipated inflation have small but significantly negative effects on markups in the US retail sector during the period from 1947 to 1985. Banerjee and Russell (2001, 2005) present evidence of a negative long-run (cointegrating) relationship between inflation and the average markup in the US and other industrialized economies during more recent periods.

As for how trend inflation affects markups, the subsequent analysis emphasizes firms' staggered price-setting as a key mechanism, based on the micro evidence documenting infrequent adjustment of consumer prices (e.g., Klenow and Malin, 2010).[11] In staggered price models, higher inflation erodes firms' markups faster in between price adjustments. Let p^* denote the relative price that a firm chooses for its product when it can adjust the price, mc its real marginal cost, and π the trend inflation rate. Then the markup of a firm that has not adjusted its price for j periods is given by $p^*/(mc\,\pi^j)$, which becomes smaller under higher trend inflation. Therefore, stickier prices—a higher age j of the nominal price—imply a larger effect of trend inflation on markups; if prices are flexible (i.e., $j = 0$), trend inflation has no effect on markups. The erosion of non-adjusting firms' markups by higher trend inflation in turn tends to reduce the average markup.

To see some empirical support for the transmission mechanism from trend inflation to the average markup, we examine whether changes in the average markup since 1980 are negatively associated with the frequency of price adjustment (FPA) across detailed US

[11] Alternatively, Bénabou (1988, 1992) propose that higher trend inflation may increase households' product market search, which intensifies competition and lowers markups.

industries. The changes in the average markup are constructed from those of De Loecker et al. (2020) between 1980 and 2016 at the four-digit NAICS industry level.[12] The FPA is based on the one calculated by Pasten et al. (2020) from the micro data underlying the producer price index for the period from 2005 to 2011 at the industry-level of the detailed input–output tables; we use industry output shares to aggregate to the four-digit industry level. With the resulting sample of industry data, we regress the change in the average markup on the FPA and a constant.[13]

Table 1 presents the regression results. The ordinary least squares (OLS) estimator for the FPA is negative and statistically significant at the 5% level, consistent with the prediction of staggered price models that less frequent price adjustment leads to a larger change in the average markup from 1980 to 2016 under the concurrent decline in trend inflation.[14] In column (2) of the table, the regression includes a change in the growth rate of industry real value added. This variable can control for other possible drivers of the change in the average markup, such as changes in regulation or openness to global competition that could affect industry growth.[15] The value added data is aggregated into broader industries than the markup data, so an observation for value added may correspond to multiple markup observations. After including this additional control, the estimator for the FPA remains negative and statistically significant. Columns (3) and (4) show that the regression results are robust to excluding the 1% smallest and largest observations for the change in the average markup. The evidence in the table encourages a quantitative analysis on the effects of trend inflation on the average markup and related macroeconomic variables.

[12]We adapted the replication code of De Loecker et al. (2020) to calculate the sales-weighted harmonic average markup, which coincides with the cost-weighted arithmetic average markup.

[13]The change in the industry inflation rate is omitted from the regression, because the producer price index is available for a large number of detailed industries only from 2004.

[14]Prior evidence on the role of nominal price rigidity pertains to the level, rather than the change, in markups or other measures of market power. Carlton (1986) analyzes firm-level data and finds that industry concentration has a strong positive correlation with nominal price rigidity in the industry. Meier and Reinelt (2021) examine firm-level data and find that firms with stickier prices tend to charge higher markups.

[15]We include the change in the average annual growth rate of real value added from the period 1980–1982 to the period 1995–1997. The cut-off in 1997 is determined by a change in industry definitions in the Bureau of Economic Analysis' industry GDP accounts.

Table 1: Regression of change in the average markup on the frequency of price adjustment.

	(1)	(2)	(3)	(4)
FPA	-0.543	-0.566	-0.545	-0.576
	(0.234)	(0.248)	(0.227)	(0.243)
ΔGDP		-0.401		-0.517
		(0.469)		(0.445)
Observations	107	107	105	105
R^2	0.034	0.039	0.044	0.054

Notes: Each column displays regression results for the change in the average markup from 1980 to 2016 in four-digit NAICS industries. In column (1), the regressors are the frequency of price adjustment (FPA) and a constant. In column (2), the regressors also include the change in the average annual growth rate of real value added from the period 1980–1982 to the period 1995–1997 (ΔGDP). Columns (3) and (4) show the results of excluding the 1% smallest and largest observations on the change in the average markup. White (heteroskedasticity-consistent) standard errors are shown in parentheses.

3 Model

For the quantitative analysis, we use a staggered price model with trend inflation. In particular, the model introduces a Kimball-type aggregator of individual differentiated goods, which includes the CES aggregator as a special case. In the model economy there are households, composite-good producers, firms, and a monetary authority. This section describes each economic agent's behavior in turn.

3.1 Households

There is a representative household that consumes a composite good C_t, makes a capital investment I_t, and supplies labor l_t so as to maximize the utility function

$$E_0 \sum_{t=0}^{\infty} \beta^t \left(\log \left(C_t \right) - \frac{l_t^{1+1/\chi}}{1+1/\chi} \right)$$

subject to the budget constraint

$$P_t \, C_t + P_t \, I_t = P_t \, W_t \, l_t + P_t \, r_{k,t} \, K_{t-1} + P_t \, J_t$$

and the capital accumulation equation

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

where E_t denotes the expectation operator conditional on information available in period t, $\beta \in (0,1)$ is the subjective discount factor, $\chi > 0$ is the elasticity of labor supply, $\delta \in (0,1)$ is the depreciation rate of capital, P_t is the price of the composite good, W_t is the real wage rate, $r_{k,t}$ is the real rental rate of capital K_{t-1}, and J_t is the real value of firm profits received.

Combining the first-order conditions for utility maximization with respect to consumption, labor supply, and capital investment yields

$$W_t = l_t^{1/\chi} C_t, \tag{2}$$

$$1 = E_t\left[\frac{\beta\, C_t}{C_{t+1}}\left(r_{k,t+1} + 1 - \delta\right)\right]. \tag{3}$$

3.2 Composite-good producers

There is a representative composite-good producer that combines the outputs of a continuum of firms $f \in [0,1]$, each of which produces an individual differentiated good $Y_t(f)$ and is subject to staggered price-setting as detailed later. As in Kimball (1995), the composite good Y_t is produced by aggregating individual differentiated goods $\{Y_t(f)\}$ with

$$\int_0^1 F\left(\frac{Y_t(f)}{Y_t}\right) df = 1. \tag{4}$$

Following Dotsey and King (2005) and Levin et al. (2008), the function $F(\cdot)$ is assumed to be of the form

$$F\left(\frac{Y_t(f)}{Y_t}\right) = \frac{\gamma}{(1+\epsilon)(\gamma-1)}\left((1+\epsilon)\frac{Y_t(f)}{Y_t} - \epsilon\right)^{\frac{\gamma-1}{\gamma}} + 1 - \frac{\gamma}{(1+\epsilon)(\gamma-1)},$$

where $\gamma \equiv \theta(1+\epsilon)$. The parameter ϵ governs the curvature $(-\epsilon\theta)$ of the demand curve for each individual good. In the special case of $\epsilon = 0$, the aggregator (4) is reduced to the CES one $Y_t = [\int_0^1 (Y_t(f))^{(\theta-1)/\theta} df]^{\theta/(\theta-1)}$, where $\theta > 1$ represents the elasticity of substitution between individual goods. The case of $\epsilon < 0$ is of particular interest in this paper because it gives rise to endogenous variation in firms' desired markups as explained later.

The composite-good producer maximizes profit $P_t Y_t - \int_0^1 P_t(f)\, Y_t(f)\, df$ subject to the aggregator (4), given individual goods' prices $\{P_t(f)\}$. Combining the first-order conditions for profit maximization and the aggregator (4) leads to

$$\frac{Y_t(f)}{Y_t} = \frac{1}{1+\epsilon}\left[\left(\frac{P_t(f)}{P_t\, d_t}\right)^{-\gamma} + \epsilon\right], \tag{5}$$

$$d_t = \left[\int_0^1 \left(\frac{P_t(f)}{P_t}\right)^{1-\gamma} df\right]^{\frac{1}{1-\gamma}}, \tag{6}$$

$$1 = \frac{1}{1+\epsilon}\, d_t + \frac{\epsilon}{1+\epsilon}\, e_t\,, \tag{7}$$

where d_t is the Lagrange multiplier on the aggregator (4) and

$$e_t \equiv \int_0^1 \frac{P_t(f)}{P_t}\, df \tag{8}$$

is the average relative price.

Eq. (5) is the demand curve for each individual good $Y_t(f)$ and features a variable price elasticity of demand for the good given by $\eta_t(f) = \theta\left[1 + \epsilon - \epsilon\left(Y_t(f)/Y_t\right)^{-1}\right]$. When $\epsilon < 0$, the elasticity $\eta_t(f)$ varies inversely with relative demand $Y_t(f)/Y_t$. That is, relative demand for each individual good becomes more (less) price-elastic for an increase (a decrease) in the relative price of the good. Consequently, the firm's desired markup for the good, $\eta_t(f)/\left(\eta_t(f) - 1\right)$, decreases (increases) for a higher (lower) relative price, which gives rise to real rigidity in relative prices.[16] In the special case of $\epsilon = 0$, where the aggregator (4) becomes the CES one as noted above, the demand curve is reduced to $Y_t(f)/Y_t = \left(P_t(f)/P_t\right)^{-\theta}$, so that the elasticity of demand and the desired markup become the constants $\eta_t(f) = \theta$ and $\theta/\left(\theta - 1\right)$, respectively.

The Lagrange multiplier d_t represents the real marginal cost of producing the composite good, and coincides with the aggregate of all individual goods' relative prices that corresponds to the quantity aggregator (4), as shown in eq. (6). In the special case of $\epsilon = 0$, eqs. (6) and (7) can be reduced to $P_t = \left[\int_0^1 \left(P_t(f)\right)^{1-\theta} df\right]^{1/(1-\theta)}$ and $d_t = 1$, respectively. The last equation shows that the real marginal cost d_t is constant in the case of the CES

[16] Models incorporating real rigidity are sometimes criticized for inducing unrealistically small absolute price changes. See Klenow and Willis (2016) for the criticism applied to Kimball-type aggregators and Nakamura and Steinsson (2010) for an in-depth discussion. In our model, although the Kimball-type aggregator dampens firms' price changes, positive trend inflation amplifies them by inducing greater variation over time in firms' relative prices.

aggregator.[17]

The output of the composite good is equal to the household's consumption and investment:

$$Y_t = C_t + I_t. \tag{9}$$

3.3 Firms

Each firm f produces an individual differentiated good $Y_t(f)$ using the Cobb-Douglas production technology

$$Y_t(f) = A_t K_t(f)^{\alpha} l_t(f)^{1-\alpha}, \tag{10}$$

where $\alpha \in (0, 1)$ is the capital elasticity of output, A_t represents the level of economy-wide technology and grows at a constant rate $A_t/A_{t-1} = g^{1-\alpha}$, and $K_t(f)$ and $l_t(f)$ are firm f's inputs of capital and labor.

Firm f minimizes cost $TC_t(f) = P_t W_t l_t(f) + P_t r_{k,t} K_t(f)$ subject to the production technology (10), given the wage rate and the capital rental rate. In the presence of economy-wide, perfectly competitive factor markets, combining the first-order conditions for cost minimization shows that all firms choose an identical capital–labor ratio[18]

$$\frac{K_t(f)}{l_t(f)} = \frac{\alpha}{1-\alpha} \frac{W_t}{r_{k,t}} = \frac{K_{t-1}}{l_t} \tag{11}$$

and incur the same real marginal cost of producing their individual goods

$$mc_t(f) = \frac{1}{A_t} \left(\frac{W_t}{1-\alpha} \right)^{1-\alpha} \left(\frac{r_{k,t}}{\alpha} \right)^{\alpha} = mc_t. \tag{12}$$

Taking into account the demand curve (5) and the marginal cost (12), firms set their product prices on a staggered basis as in Calvo (1983). In each period, a fraction $\xi \in (0, 1)$ of firms keeps prices unchanged, while the remaining fraction $1 - \xi$ sets the price $P_t(f)$ so as

[17] Moreover, if all firms share the same production technology (as assumed later) and all individual goods' prices are flexible, the prices are all identical and thus eqs. (6) and (7) imply that $d_t = 1$ even in the case of the non-CES aggregator.

[18] The last equality in eq. (11) can be obtained by using the capital and labor market clearing conditions $K_{t-1} = \int_0^1 K_t(f) df$ and $l_t = \int_0^1 l_t(f) df$.

to maximize relevant profit

$$E_t \sum_{j=0}^{\infty} \xi^j \, Q_{t,t+j} \left(P_t(f) - P_{t+j} \, mc_{t+j}\right) \frac{Y_{t+j}}{1+\epsilon} \left[\left(\frac{P_t(f)}{P_{t+j} \, d_{t+j}} \right)^{-\gamma} + \epsilon \right],$$

where $Q_{t,t+j}$ is the (nominal) stochastic discount factor between period t and period $t+j$. Using the equilibrium condition $Q_{t,t+j} = \beta^j (C_t/C_{t+j})/(P_t/P_{t+j})$, the first-order condition for profit maximization can be written as

$$E_t \sum_{j=0}^{\infty} (\beta\xi)^j \frac{Y_{t+j}}{C_{t+j}} \left[\left(\frac{p_t^*}{d_{t+j}} \right)^{-\gamma} \prod_{\tau=1}^{j} \pi_{t+\tau}^{\gamma} \left(p_t^* \prod_{\tau=1}^{j} \pi_{t+\tau}^{-1} - \frac{\gamma}{\gamma-1} mc_{t+j} \right) - \frac{\epsilon}{\gamma-1} p_t^* \prod_{\tau=1}^{j} \pi_{t+\tau}^{-1} \right] = 0, \tag{13}$$

where $p_t^* \equiv P_t^*/P_t$, P_t^* is the price set by firms that can change prices in period t, and $\pi_t \equiv P_t/P_{t-1}$ is the gross inflation rate of the composite good's price. Moreover, under staggered price-setting, eqs. (6) and (8) can be reduced to, respectively,

$$d_t^{1-\gamma} = \xi \, \pi_t^{\gamma-1} \, d_{t-1}^{1-\gamma} + (1-\xi)(p_t^*)^{1-\gamma}, \tag{14}$$

$$e_t = \xi \, \pi_t^{-1} \, e_{t-1} + (1-\xi) \, p_t^*. \tag{15}$$

3.4 Monetary authority and equilibrium conditions

The monetary authority is assumed to choose the trend inflation rate π, which represents its inflation target in the model. The trend inflation rate influences real outcomes in steady state through its effects on two distortions: the average markup and the relative price distortion. We discuss each of the distortions in turn.[19]

In the model, the relative price distortion can be calculated as

$$\Delta_t \equiv \frac{s_t + \epsilon}{1+\epsilon}, \tag{16}$$

[19]While most studies attribute measured declines in trend inflation after the Great Inflation to improved monetary policy, some have pointed to other factors (e.g., globalization) as possible drivers of the declines (e.g., Chen et al., 2004; Borio and Filardo, 2007). In this context, Rogoff (2004) argues that globalization and deregulation have contributed to the declines in trend inflation, by reducing the wedge between the socially optimal level and the market level of output, thus weakening the incentive for central banks to produce unanticipated inflation. Such factors are captured succinctly in our simple model by the parameter θ that governs the elasticity of substitution between goods. If a discretionary monetary policymaker sets the trend inflation rate optimally, given the values of model parameters, then a larger value of θ would lead the policymaker to choose a lower trend inflation rate. However, the effect on the average markup would be ambiguous, as the larger value of θ reduces the average markup but the lower trend inflation rate raises it.

where

$$s_t \equiv \int_0^1 \left(\frac{P_t(f)}{P_t\, d_t}\right)^{-\gamma} df, \tag{17}$$

since we have

$$Y_t\, \Delta_t = A_t K_{t-1}^{\alpha} l_t^{1-\alpha} \tag{18}$$

by combining the demand curve (5), the production function (10), and the capital and labor market clearing conditions $K_{t-1} = \int_0^1 K_t(f)df$ and $l_t = \int_0^1 l_t(f)df$ under firms' identical capital–labor ratio (11). The relative price distortion Δ_t then measures the inefficiency of aggregate production arising from demand dispersion under staggered price-setting as combining eqs. (5), (16), and (17) leads to

$$\Delta_t = \int_0^1 \frac{Y_t(f)}{Y_t} df.$$

If all prices are flexible, all firms charge the same price because they share the same production technology (10). Consequently, eqs. (16) and (17) demonstrate no relative price distortion, i.e., $\Delta_t = s_t = 1$, and the aggregate production equation (18) implies no inefficiency in producing aggregate output using aggregate capital and labor. Staggered price-setting gives rise to demand dispersion and thus introduces an inefficiency in aggregate production, which is exacerbated under higher trend inflation. Eq. (17) can be reduced, under staggered price-setting, to

$$d_t^{-\gamma} s_t = \xi \pi_t^{\gamma} d_{t-1}^{-\gamma} s_{t-1} + (1-\xi)(p_t^*)^{-\gamma}. \tag{19}$$

We follow Edmond et al. (2021) to consider the cost-weighted average price–cost markup

$$\mu_t = \int_0^1 \frac{TC_t(f)}{\int_0^1 TC_t(f)\, df} \frac{P_t(f)}{P_t\, mc_t} df = \frac{1}{mc_t\, \Delta_t}, \tag{20}$$

where each firm's cost weight is given by

$$\frac{TC_t(f)}{\int_0^1 TC_t(f)\, df} = \frac{P_t\, mc_t\, Y_t(f)}{\int_0^1 P_t\, mc_t\, Y_t(f)\, df} = \frac{Y_t(f)}{Y_t\, \Delta_t}.$$

Therefore, the average markup coincides with the reciprocal of the real marginal cost mc_t and the relative price distortion Δ_t. If all prices are flexible, firms can attain their desired markups. Under staggered price-setting, however, firms choose a price that meets the profit-maximizing condition (13) when they can adjust prices. Thus, a firm's markup varies depending on how long its price has remained unchanged, and higher trend inflation exacerbates the erosion of firms' relative prices and hence their markups in between price changes.

The equilibrium conditions of the model consist of eqs. (1)–(3), (7), (9), (11)–(16), (18), and (19). These conditions are rewritten in terms of detrended variables: $y_t \equiv Y_t/\Upsilon_t$, $c_t \equiv C_t/\Upsilon_t$, $i_t \equiv I_t/\Upsilon_t$, $k_t \equiv K_t/\Upsilon_t$, $w_t \equiv W_t/\Upsilon_t$, and $j_t \equiv J_t/\Upsilon_t$, where $\Upsilon_t = A_t^{1/(1-\alpha)}$. This implies that the growth rate of Υ_t (i.e., $\Upsilon_t/\Upsilon_{t-1} = g$) represents the rate of balanced growth.

3.5 Steady state

For the steady state to be well defined, the following condition is assumed to be satisfied:

$$\xi \max(\pi^{\gamma}, \pi^{\gamma-1}, \pi^{-1}) < 1. \tag{21}$$

This condition is rewritten as $\xi \max(\pi^{\theta}, \pi^{\theta-1}) < 1$ in the special case of the CES aggregator, i.e., $\epsilon = 0$.[20]

Using the equilibrium conditions, we can obtain the equations for the real marginal cost mc and the relative price distortion Δ in the steady state with trend inflation π

$$mc = \frac{\gamma - 1 - \tilde{\epsilon}}{\gamma} \frac{1 - \beta\xi\pi^{\gamma}}{1 - \beta\xi\pi^{\gamma-1}} \left[\frac{1}{1+\epsilon} \left(\frac{1-\xi}{1-\xi\pi^{\gamma-1}} \right)^{\frac{1}{1-\gamma}} + \frac{\epsilon}{1+\epsilon} \frac{1-\xi}{1-\xi\pi^{-1}} \right]^{-1}, \tag{22}$$

$$\Delta = \frac{1}{1+\epsilon} \frac{1-\xi}{1-\xi\pi^{\gamma}} \left(\frac{1-\xi}{1-\xi\pi^{\gamma-1}} \right)^{\frac{\gamma}{1-\gamma}} + \frac{\epsilon}{1+\epsilon}, \tag{23}$$

where $\tilde{\epsilon} \equiv \epsilon[(1-\xi\pi^{\gamma-1})/(1-\xi)]^{\gamma/(1-\gamma)}(1-\beta\xi\pi^{\gamma-1})/(1-\beta\xi\pi^{-1})$. The average markup equation (20) gives its steady-state value

$$\mu = \frac{1}{mc\,\Delta}. \tag{24}$$

In the steady state, the profit share of income and the labor share are given by

$$\frac{j}{y} = 1 - \frac{1}{\mu}, \tag{25}$$

$$\frac{wl}{y} = (1-\alpha)\frac{1}{\mu}, \tag{26}$$

[20] The condition is always met in the special case of zero trend inflation, i.e., $\pi = 1$.

and the investment share of spending i/y is the product of the investment–capital ratio $i/k = 1-(1-\delta)/g$ and the capital–output ratio $k/y = \alpha g/(r_k\,\mu)$, where $r_k = g/\beta-(1-\delta)$, so that

$$\frac{i}{y} = \alpha\beta\frac{g-(1-\delta)}{g-\beta(1-\delta)}\frac{1}{\mu}. \tag{27}$$

Note that the cost-weighted average markup is the only and the common driver of the profit share (25), the labor share (26), and the investment share (27).[21] That is, trend inflation π influences the three shares through its effect on the steady-state average markup μ.

4 Effects of declining trend inflation

Using the steady-state equations (22)–(27), this section evaluates the quantitative effects of a decline in trend inflation as measured since 1980 on the steady-state values of the average markup, the profit share of income, the labor share, and the investment share of spending.

4.1 Calibration of model parameters

As seen in the preceding section, the steady-state equations (22)–(27) are highly nonlinear functions of trend inflation π. We therefore use a calibration of model parameters to illustrate how the steady-state values vary with π.

Table 2 summarizes the calibration of parameters in the quarterly model.[22] We set the subjective discount factor at $\beta = 0.99$, the depreciation rate of capital at $\delta = 0.025$, and the capital elasticity of output at $\alpha = 0.3$, which all are common values in the macroeconomic literature. The rate of balanced growth is chosen at $g = 1.005$, that is, 2% annually. The probability of no price change—the so-called Calvo parameter—is set at $\xi = 0.75$, which implies that prices change every four quarters on average in line with the micro evidence (e.g., Klenow and Malin, 2010). The parameter governing the elasticity of substitution between individual goods is chosen at $\theta = 4.1$ to target an average markup of 1.31 at the annualized trend inflation rate of 1.6%, their values for 2016 displayed in Figures 1 and 2.

[21]As indicated by Basu (2019) and Syverson (2019), the average markup can be written as $\mu = [1/(1-j/y)]\,(ac/mc)$, where the term in brackets involves the profit share of income and the term ac/mc is the inverse of the elasticity of costs with respect to quantity, with ac denoting the average cost. Comparing this equation with eq. (25) shows that the cost elasticity is equal to one in the model.

[22]The value for the elasticity of labor supply χ is unspecified because it has no effect on the steady-state values of the average markup and the three other macroeconomic variables.

Nakamura and Steinsson (2010) observe that such a value of θ matches estimates from the literature on industrial organization and international trade. As for the parameter governing the curvature of demand curves, we select a value of $\epsilon = -8$, which implies, given our calibration of θ, a curvature of $-\epsilon\theta = 32.8$.[23] This degree of curvature is close to the high value of 33 considered by Eichenbaum and Fisher (2007), intermediate between the values of 16.7 and 65.9 implied by the two estimates of Guerrieri et al. (2010), and below the estimate by Hirose et al. (2021) of 49.1 for the post-1982 period.[24] To meet the assumption (21) under the model parameter values presented above, the trend inflation rate needs to be greater than -3.9% annually.

Table 2: Calibration of parameters in the quarterly model.

Parameter	Description	Value
β	Subjective discount factor	0.99
δ	Depreciation rate of capital	0.025
α	Capital elasticity of output	0.3
g	Rate of balanced growth	1.005
ξ	Probability of no price change	0.75
θ	Parameter governing the elasticity of substitution between goods	4.1
ϵ	Parameter governing the curvature of demand curves	-8

4.2 Quantitative effects

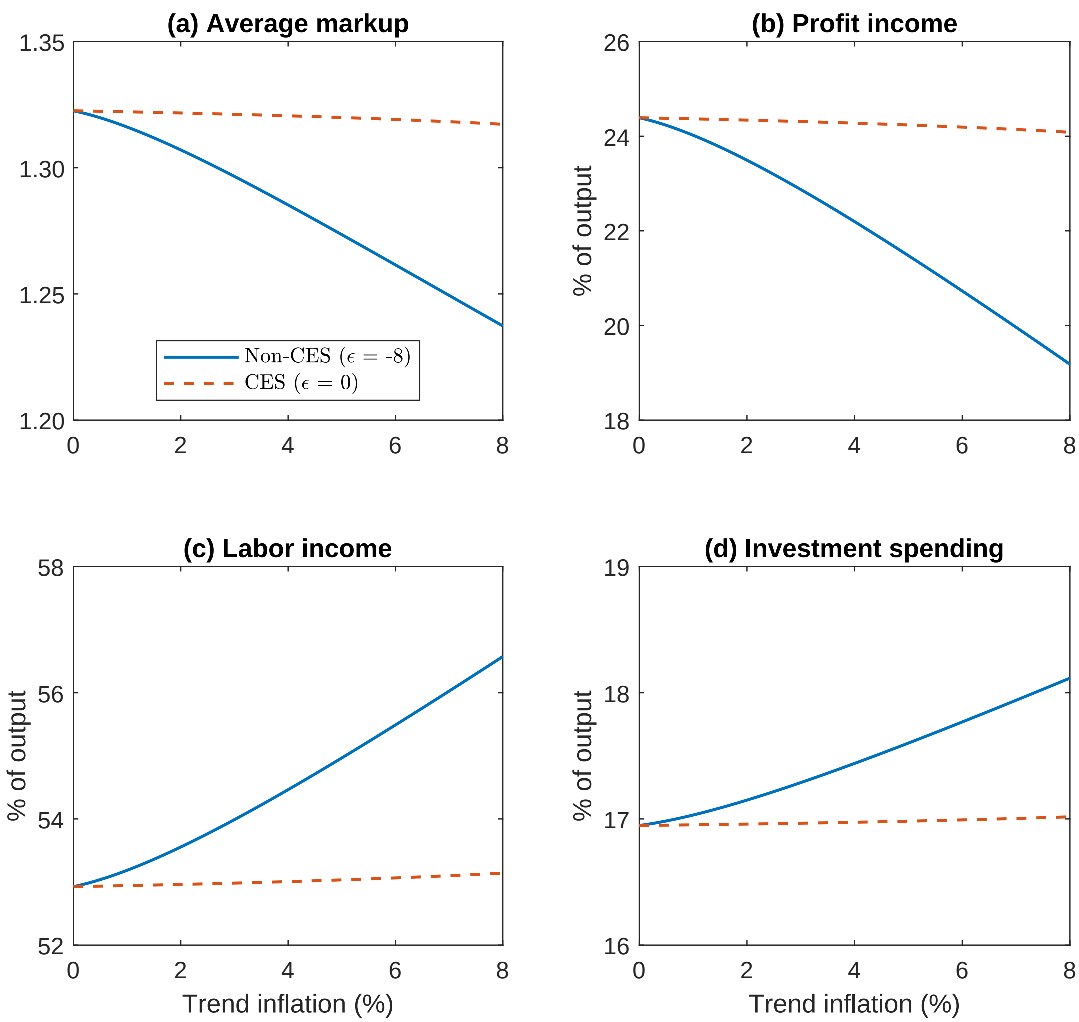
Trend inflation has declined steadily since the Volcker disinflation in the early 1980s, as displayed in Figure 2. Figure 3 illustrates the steady-state effects of lower trend inflation in the model. As shown by the solid lines, the non-CES aggregator leads lower trend inflation to increase the average markup (panel a) and the profit share (panel b) and to decrease the labor share (panel c) and the investment share (panel d). The increasing average markup is consistent with the evidence of De Loecker et al. (2020), Edmond et al. (2021), and Hall (2018) that the average markup has risen since the early 1980s, and the increasing profit

[23]We also considered selecting values of θ and ϵ so as to minimize the distance between the empirical cost-weighted average markup and its model counterpart at different trend inflation rates. Such values are computed as $\theta = 4.6$ and $\epsilon = -25.6$. Because these values imply a high curvature of $-\theta\epsilon = 116$, we adopted the more conservative values of $\theta = 4.1$ and $\epsilon = -8$.

[24]Micro evidence of Dossche et al. (2010) and Beck and Lein (2020) also points to curvature in demand curves, although its magnitude is substantially smaller and pertains to household spending on relatively narrow categories of retail goods in European countries.

share is in line with the evidence of Barkai (2020), who documents a rise in the profit share from 1984. Likewise, the declining labor share is pointed out by Elsby et al. (2013), Karabarbounis and Neiman (2014), Autor et al. (2020), and Kehrig and Vincent (2021), while the decreasing investment share is indicated by Covarrubias et al. (2020). In contrast, the dashed lines represent the case of the CES aggregator in which lower trend inflation has minor effects on all the macroeconomic variables.

Figure 3: Steady-state values of the macroeconomic variables as functions of trend inflation.



Notes: The figure illustrates the effects of the annualized trend inflation rate $\bar{\pi}\,(\equiv 400 \log \pi)$ on the average markup, the profit share, the labor share, and the investment share in the steady states of the models with the non-CES aggregator (solid lines) and with the CES aggregator (dashed lines). The values of model parameters used here are reported in Table 2.

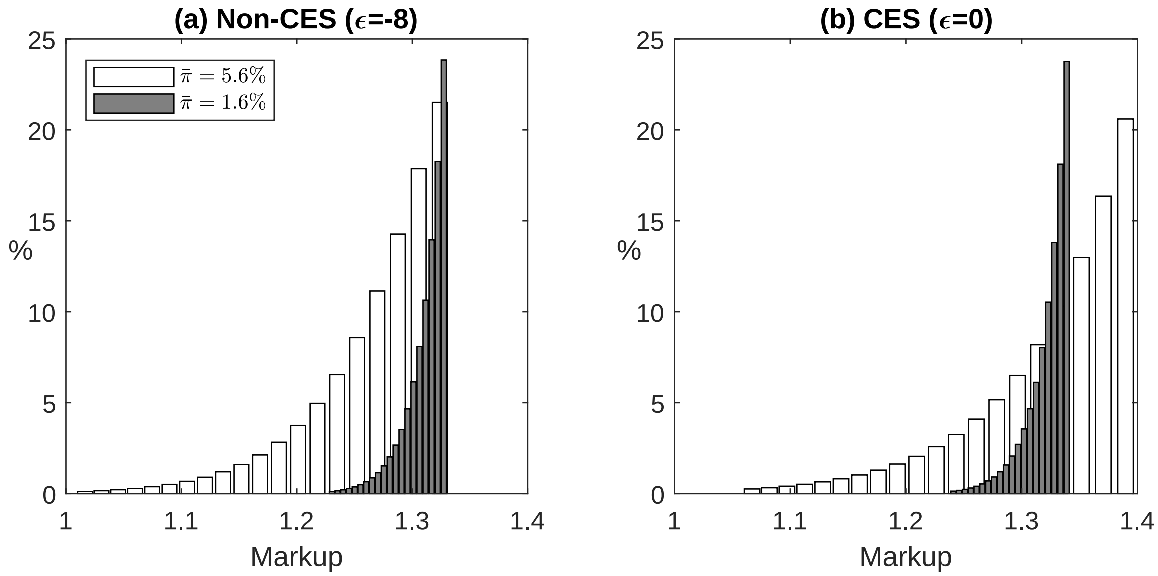
To understand why the non-CES aggregator leads lower trend inflation to raise the average markup, we can look at how the aggregator affects the distribution of markups across firms. In steady state, the average markup consists of the markups of $1-\xi$ firms that set prices in the current period, those of $(1-\xi)\,\xi$ firms that set prices in the previous period, and so forth. The age of a firm's price also determines its cost weight, as can be seen from each individual good's demand curve (5). Thus, the distribution of markups is represented by the density function

$$f(j) = (1-\xi)\,\xi^j \frac{1}{\Delta}\left[\frac{1}{1+\epsilon}\left(\frac{1-\xi}{1-\xi\pi^{\gamma-1}}\right)^{\frac{\gamma}{1-\gamma}}\pi^{\gamma j} + \frac{\epsilon}{1+\epsilon}\right],$$

where j is the age of a firm's price. The average markup is then given by $\mu = \sum_{j=0}^{\infty} f(j)\,[p^*/(mc\,\pi^j)]$, where $p^*/mc = [\gamma/(\gamma-1-\tilde{\epsilon})][(1-\beta\xi\pi^{\gamma-1})/(1-\beta\xi\pi^{\gamma})]$ is the markup chosen by price-adjusting firms.

Figure 4 displays the steady-state markup distribution obtained with the non-CES aggregator (panel a) and with the CES aggregator (panel b) for two values of annualized trend inflation, 5.6% in 1980 (white bars) and 1.6% in 2019 (gray bars), which are plotted in Figure 2. The lower trend inflation reduces the lower tail of the markup distribution, regardless of the CES or non-CES aggregator, because firms that keep their nominal prices unchanged experience a less severe erosion of their relative prices and hence their markups. Then, in the case of the CES aggregator, the lower trend inflation also induces price-adjusting firms to choose a smaller price increase because firms are forward-looking, as pointed out by King and Wolman (1996). The resulting lower markups of price-adjusting firms offset most of the contribution of the less severely eroding markups of non-adjusting firms, and as a consequence, the average markup is almost flat for the lower trend inflation, as shown by the dashed line in panel (a) of Figure 3. In contrast, the non-CES aggregator gives rise to variable elasticity of demand, which leads price-adjusting firms' desired markup to increase for a lower relative price of the firms induced by the lower trend inflation. Consequently, price-adjusting firms' markups are less responsive to the lower trend inflation. Therefore, the lower trend inflation raises the average markup through the thinner lower tail of non-adjusting firms' markups. In short, the non-CES aggregator leads the lower trend inflation to reduce the skewness to the left of the steady-state markup distribution, thereby increasing the average markup.

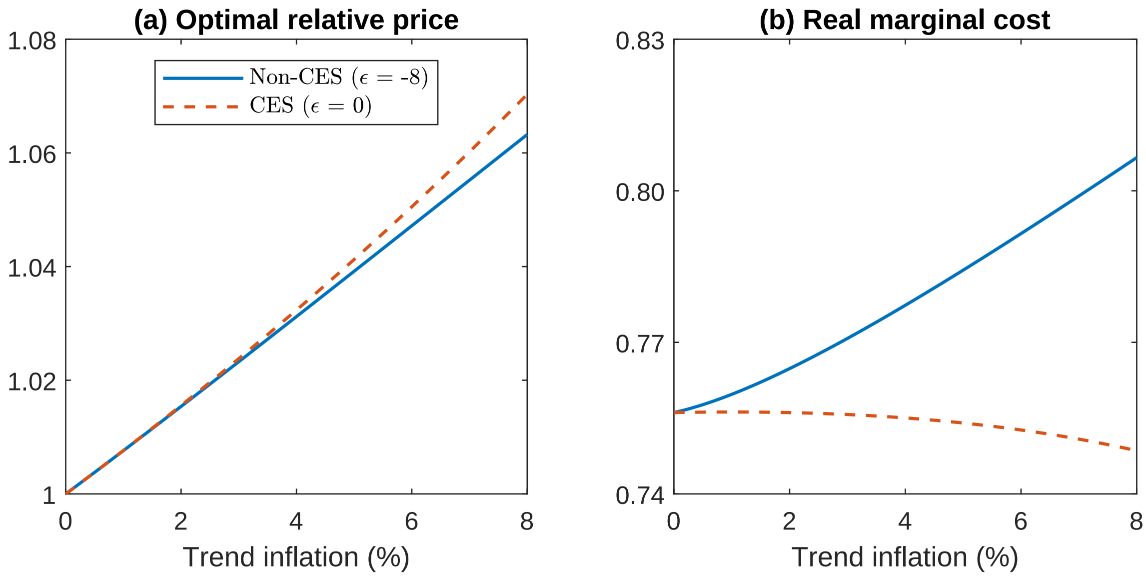
Figure 4: Steady-state distribution of firms' markups with cost weights.



Notes: The figure illustrates the distribution of firms' markups with cost weights in the steady states with the annualized trend inflation rate $\bar{\pi}$ ($\equiv 400 \log \pi$) of 5.6% (white bars) and 1.6% (gray bars) in the models with the non-CES aggregator (left panel) and with the CES aggregator (right panel). The plotted distributions display the 20 most recent price vintages, which account for more than 99% of the probability mass. The values of model parameters used here are reported in Table 2.

In the case of the non-CES aggregator, price-adjusting firms' less responsive markup to the lower trend inflation implies that the price set by the firms changes roughly proportionally with their marginal cost. Figure 5 plots the steady-state effects of lower trend inflation on price-adjusting firms' optimal relative price p^* (panel a) and real marginal cost mc (panel b). Lower trend inflation reduces both variables in the case of the non-CES aggregator, whereas it decreases only the price p^* in the case of the CES aggregator. In the latter case, price-adjusting firms choose a lower relative price for lower trend inflation, not because they face a lower real marginal cost today, but because they anticipate a slower future erosion of their markups, as noted above. In contrast, the endogenous desired markups arising from the non-CES aggregator weaken the responsiveness of price-adjusting firms' optimal relative price to trend inflation, but this firm-level effect is partly offset by a macroeconomic effect: in steady state, lower trend inflation raises the average markup, which reduces output and hence the real marginal cost, thus pulling down the optimal relative price. On balance, the optimal relative price is only modestly less sensitive to trend inflation than in the case of

Figure 5: Steady-state values of price-adjusting firms' optimal relative price and real marginal cost as functions of trend inflation.



Notes: The figure illustrates the effects of the annualized trend inflation rate $\bar{\pi}\,(\equiv\, 400\log\pi)$ on price-adjusting firms' optimal relative price and real marginal cost in the steady states of the models with the non-CES aggregator (solid lines) and with the CES aggregator (dashed lines). The values of model parameters used here are reported in Table 2.

the CES aggregator. This is why evidence parsing the relevance of our model's mechanism must consider markups as in Section 2, rather than reset prices.

To quantify how much of the US macroeconomic changes can be attributed to the decline of trend inflation measured since 1980, panel (a) of Table 3 presents the percentage point changes in the annualized trend inflation rate, the average markup, the profit share, the labor share, and the investment share from 1980 to 2019 or the most recent available year as displayed in Figures 1 and 2, while panel (b) of the table reports the model-predicted changes in the steady-state values of the macroeconomic variables induced by a decline in trend inflation of equal size. The annualized trend inflation rate declined by 4 percentage points from 5.6% in 1980 to 1.6% in 2019. The average markup concurrently increased by 14.9 percentage points, whereas the model predicts that the decline in trend inflation increases the average markup by 4.5 percentage points. The profit share rose by 8.6 percentage points, while the model predicts a rise of 2.7 percentage points. The labor share and the investment share decreased by 6.7 percentage points and 2.9 percentage points, respectively, and their counterparts in the model prediction are decreases of 1.9 percentage points and 0.6 percentage point, respectively. In short, the model attributes around 30% of the increases in the average markup and the profit share and the decrease in the labor share and about 20% of the decrease in the investment share to the decline in trend inflation.[25]

Table 3: Macroeconomic changes from 1980 to 2019.

	Trend Inflation (% pa)	Average markup	Profit share (%)	Labor share (%)	Investment share (%)
(a) *US economy*					
1980	5.6	1.16	5.00	63.48	21.07
2019	1.6	1.31[a]	13.60	56.80	18.20
Change (percentage points)	−4.0	14.86	8.60	−6.68	−2.87
(b) *Model*					
Steady-state value	5.6	1.27	21.03	55.28	17.70
Steady-state value	1.6	1.31	23.72	53.40	17.10
Change (percentage points)	−4.0	4.46	2.69	−1.88	−0.60
(c) *Model with superstar firms*					
Steady-state value with $z = 1.0$	5.6	1.27	21.03	55.28	17.70
Steady-state value with $z = 1.6$	1.6	1.35	25.82	51.92	16.63
Change (percentage points)	−4.0	8.18	4.79	−3.36	−1.07

Notes: The data for the US economy shown in panel (a) of the table are described in the notes to Figures 1 and 2. The value marked by 'a' pertains to 2016, the most recent available observation. The values of model parameters used for panel (b) of the table are reported in Table 2. They are also used for panel (c) of the table, along with the values of ordinary firms' share of $n = 0.86$ and superstar firms' relative productivity z indicated in the first column of the table.

Before proceeding to a robustness analysis, it is worth noting that the decline in trend inflation substantially reduces the lost profit stemming from staggered price-setting in the model (with the non-CES aggregator). Using the values of model parameters reported in Table 2, steady-state average profit forgone as a result of staggered price-setting—i.e., the percentage change in steady-state average profit from the case of flexible prices to that of staggered price-setting—in 1980 is 10%. This level is comparable to the lost profit due to sticky prices calculated for the US economy by Zbaracki et al. (2004) (about 20%) and that by Gorodnichenko and Weber (2016) (about 25%). The steady-state average lost profit

[25] We obtained similar numbers under the calibration of $\theta = 4.6$ and $\epsilon = -25.6$ presented in footnote 23: 33%, 37%, 34%, and 25% for the increases in the average markup and the profit share and the decreases in the labor share and the investment share, respectively. Therefore, the results are robust for a wide range of values for the curvature of demand curves.

decreases to 1.9% in 2019, as the decline in trend inflation from 5.6% to 1.6% raises steady-state average profit closer to its level under flexible prices. Furthermore, the decline in trend inflation raises total factor productivity by mitigating the relative price distortion, but the gain in productivity is small (0.1%).[26]

4.3 Robustness analysis

The aforementioned quantitative effects of the decline in trend inflation on the macroeconomic variables may depend on some modeling assumptions. In this subsection we thus inspect the robustness of the effects by altering key assumptions. The alternative assumptions include (i) increasing nominal price rigidity with the decline in trend inflation, (ii) assuming alternative specifications of nominal price rigidity, (iii) incorporating some key features of medium-scale models that are absent from our model, and (iv) introducing time variation in trend inflation. Details of the last three sets of robustness exercises are provided in Appendix A.

4.3.1 Increase in nominal price rigidity

One may wonder if the Calvo parameter ξ changes with trend inflation. Previous studies exhibit mixed empirical results. While Galí and Gertler (1999) estimate the New Keynesian Phillips curve (NKPC) during two periods before and after 1980 and report an increase in the Calvo parameter, Eichenbaum and Fisher (2007) introduce a Kimball-type non-CES aggregator in the NKPC and show no significant change in the estimate of the Calvo parameter between the pre-1979 and the post-1982 periods. Hirose et al. (2020) also demonstrate no change of the Calvo parameter in a dynamic stochastic general equilibrium (DSGE) model that is estimated during the pre-1979 and the post-1982 periods while allowing for indeterminacy of equilibrium. In contrast, Smets and Wouters (2007) estimate a DSGE model for the two periods within only the determinacy region of its parameter space and report an increase in the Calvo parameter. Similarly, Hirose et al. (2021) find a rise of the Calvo parameter from the pre-1979 to the post-1982 period in a DSGE model augmented with Kimball-type

[26] In the case of the CES aggregator, the steady-state average lost profit due to staggered price-setting is much smaller, that is, 1.4% in 1980 and 0.2% in 2019, while the gain in total factor productivity is somewhat larger but remains small (0.6%).

non-CES aggregators that is estimated in both determinacy and indeterminacy regions. As for estimates based on micro data, Nakamura et al. (2018) point to a decrease in the frequency of price change after the Great Inflation, whereas Alvarez et al. (2019) stress that the frequency of price adjustment does not vary with inflation at low inflation rates (less than 10%), although at high inflation rates the elasticity of the price adjustment frequency with respect to inflation is close to two-thirds.

In our calibrated model, if the decline in trend inflation is accompanied by a rise in the Calvo parameter from $\xi = 0.67$—which implies that prices change every three quarters on average—to $\xi = 0.75$, then panel (a) of Table 4 shows that the contribution of the decline in trend inflation to the changes in the steady-state values of all the macroeconomic variables becomes smaller but remains more than half as large as that in the baseline model presented in panel (b) of Table 3.[27]

4.3.2 Specification of nominal price rigidity

We have emphasized staggered price-setting as a key mechanism for the effects of the decline in trend inflation on the macroeconomic variables. To evaluate to what extent the specification of nominal price rigidity matters for the effects, we consider two alternatives, a fixed duration staggered contract of Taylor (1980) and a quadratic price adjustment cost of Rotemberg (1982), although the latter is at odds with the micro evidence on price adjustment because it implies that firms continually change their product prices.

Panel (b) of Table 4 indicates that replacing the Calvo-style random duration staggered contracts with Taylor-style fixed duration ones leads to a somewhat smaller contribution of the decline in trend inflation to the changes in the steady-state values of all the macroeconomic variables, using a calibration that delivers the same average age of contracts between those with fixed and random duration, as suggested by Dixon and Kara (2006). In contrast, employing a quadratic price adjustment cost reduces the effects of the decline in trend inflation on the macroeconomic variables substantially as seen in panel (c) of the table, since it induces no price dispersion so that the distribution of markups is degenerate.

[27] The assumed rise in the Calvo parameter is roughly consistent with the empirical relationship between trend inflation and the frequency of price adjustment estimated by L'Huillier and Schoenle (2021), which, averaged across the four specifications reported in their Table 1, implies the Calvo parameter of $\xi = 0.63$ and $\xi = 0.77$ at the annualized trend inflation rate of 5.6% and 1.6%, respectively.

Table 4: Macroeconomic changes from 1980 to 2019 in robustness exercises.

	Trend Inflation (% pa)	Average markup	Profit share (%)	Labor share (%)	Investment share (%)
(a) *Increasing nominal price rigidity*					
Steady-state value with $\xi = 0.67$	5.6	1.29	22.34	54.36	17.41
Steady-state value with $\xi = 0.75$	1.6	1.31	23.72	53.40	17.10
Change (percentage points)	−4.0	2.33	1.38	−0.96	−0.31
(b) *Fixed duration contract:* $J = 7$					
Steady-state value	5.6	1.28	21.96	54.63	17.49
Steady-state value	1.6	1.32	24.07	53.15	17.01
Change (percentage points)	−4.0	3.56	2.11	−1.48	−0.47
(c) *Price adjustment cost:* $\zeta = 25$					
Steady-state value	5.6	1.32	24.06	52.99	16.97
Steady-state value	1.6	1.32	24.35	52.94	16.95
Change (percentage points)	−4.0	0.11	0.29	−0.04	−0.01
(d) *Roundabout production:* $\phi = 0.5$					
Steady-state value	5.6	1.27	34.75	45.67	14.63
Steady-state value	1.6	1.31	38.34	43.16	13.82
Change (percentage points)	−4.0	4.46	3.59	−2.51	−0.80
(e) *Fixed production cost:* $\omega/y = 0.133$					
Steady-state value	5.6	1.27	10.54	62.62	20.05
Steady-state value	1.6	1.31	13.57	60.50	19.37
Change (percentage points)	−4.0	4.41	3.03	−2.12	−0.68
(f) *Time-varying trend inflation*					
Mean value for 1980	5.6	1.27	21.08	55.32	17.29
Mean value for 2019	1.6	1.32	24.04	53.19	17.27
Change (percentage points)	−4.0	5.05	2.96	−2.12	−0.02

Note: The values of model parameters used for the table are reported in its first column as well as in Table 2.

4.3.3 Medium-scale model features

Ascari et al. (2018) point out that the business cycle and welfare implications of positive trend inflation are amplified by the presence of staggered wage-setting and a roundabout production structure in a medium-scale DSGE model otherwise similar to our model. Thus, we assess whether these features alter the effects of the decline in trend inflation on the macroeconomic variables.

First, we augment our model with staggered wage-setting following Erceg et al. (2000) along with a possibly non-CES labor aggregator that takes the same functional form as the goods aggregator $F(\cdot)$ in eq. (4). We find that these features do not alter the results obtained with the baseline model, since the steady-state values of real marginal cost and relative price distortion—which are given by eqs. (22) and (23) and determine the four macroeconomic variables in eqs. (24)–(27)—do not depend on the specification of the labor disutility function (for example, the labor supply elasticity χ).

Next, we embed a roundabout production structure as in Basu (1995) in the model. As shown in panel (d) of Table 4, such a structure does not alter the steady-state average markup either, as the steady-state values of real marginal cost and relative price distortion do not depend on the specification of the production function (for example, the capital elasticity of output α). Yet the roundabout production structure increases the profit share and decreases the labor share and the investment share, moving the values of these three variables away from their respective data counterparts, although the model can better explain the changes in the data on the three variables from 1980 to 2019.[28]

4.3.4 Time variation in trend inflation

The effects of the decline in trend inflation reported in panel (b) of Table 3 are obtained under the assumption that agents anticipate no change in trend inflation when forming expectations.[29] In this exercise we let agents account for time variation in trend inflation by assuming that they form rational expectations while the monetary authority's inflation target π_t^* evolves according to an autoregressive process

$$\log \pi_t^* = (1 - \rho_\pi) \log \pi + \rho_\pi \log \pi_{t-1}^* + \varepsilon_t, \tag{28}$$

where ρ_π is a persistence parameter and ε_t is a shock to the target. For this exercise, we will refer to π_t^* as the time-varying trend inflation rate (or inflation target) and to π as the steady-state inflation rate. We also assume that the representative household purchases

[28]We also incorporated a fixed cost of production in the model and calibrated it by targeting the profit share observed for the US economy. The fixed cost aligns the labor share and the investment share in the model more closely to their data counterparts and increases changes in the three shares induced by the decline in trend inflation, as seen in panel (e) of Table 4.

[29]This setting can alternatively be interpreted as a driftless random walk process of trend inflation under the assumption of an anticipated utility model of Kreps (1998), as in Cogley and Sbordone (2008). Candia et al. (2021) study survey expectations of US firms and find that, while firms are not always well-informed about the Federal Reserve's inflation target, their long-run inflation expectations reflect their perceived inflation target of the central bank.

one-period bonds B_t that earn the gross interest rate r_t, which leads to the consumption Euler equation

$$1 = E_t\left(\frac{\beta\, C_t}{C_{t+1}} \frac{r_t}{\pi_{t+1}}\right). \tag{29}$$

Moreover, the monetary authority sets the interest rate r_t following a Taylor (1993)-type rule

$$\log r_t = \log r + \phi_\pi \left(\log \pi_t - \log \pi_t^*\right). \tag{30}$$

To evaluate the macroeconomic effects of the time-varying trend inflation, we use a log-linear approximation to the model augmented with the consumption Euler equation (29) and the monetary policy rule (30). The calibration of Table 2 is extended by adopting common values for the policy response of $\phi_\pi = 1.5$ and the labor supply elasticity of $\chi = 1$. To calibrate the persistence parameter of the inflation target shock and the steady-state inflation rate, the autoregressive process (28) is estimated on the trend inflation series of Chan et al. (2018) from 1960Q2 to 2019Q4. The OLS estimators yield the values of the persistence parameter of $\rho_\pi = 0.996$ and the (annualized) steady-state inflation rate of $\bar{\pi}$ ($\equiv 400 \log \pi$) $= 2.89\%$. We then conduct a counterfactual exercise wherein the OLS residuals are taken to be realizations of the inflation target shock innovation ε_t in the calibrated model, so that the model replicates the trend inflation series and simulates the associated changes in the macroeconomic variables. As presented in panel (f) of Table 4, the exercise indicates that the decline of trend inflation measured from 1980 to 2019 explains a somewhat larger portion of the changes in the average markup, the profit share, and the labor share than that in the baseline model reported in panel (b) of Table 3, though it does not explain the decline in the investment share.

5 Complementarity with superstar firm hypothesis

The analysis in the preceding section indicates that the decline of trend inflation as measured since 1980 may have contributed substantially to the concurrent rise in the average markup but is not the only driving factor. In the recent literature, the superstar firm hypothesis first proposed by Autor et al. (2020)—that a rise in the average markup since the early 1980s stems from the increased importance of highly productive superstar firms with large markups—is a leading explanation for the rising average markup.[30] While the decline in trend inflation leads to a thinner lower tail of the steady-state distribution of firms' markups, the rise of superstar firms gives rise to a thicker upper tail. This section thus examines the joint effect of the decline in trend inflation and the rise of superstar firms on the average markup and the three other macroeconomic variables. To this end, our model is extended by introducing highly productive superstar firms.[31] Specifically, the extended model assumes that a fraction $1 - n \in [0, 1)$ of firms (i.e., superstar firms) is more productive than the other n firms (i.e., ordinary firms). The production function (10) and the real marginal cost (12) are then replaced with

$$Y_t(f) = z(f) A_t K_t(f)^{\alpha} l_t(f)^{1-\alpha}, \tag{31}$$

$$mc_t(f) = \frac{1}{A_t z(f)} \left(\frac{W_t}{1-\alpha} \right)^{1-\alpha} \left(\frac{r_{k,t}}{\alpha} \right)^{\alpha}, \tag{32}$$

where $z(f) = 1$ if $f \in [0, n]$ and $z(f) = z > 1$ otherwise. Consequently, the steady-state cost-weighted average markup can be calculated as

$$\mu = \frac{1}{mc_1 \, \Delta_1 + mc_z \, \Delta_z}, \tag{33}$$

where $mc_1 = mc(f)$, $f \in [0, n]$; $mc_z = mc(f)$, $f \in (n, 1]$; and Δ_i, $i \in \{1, z\}$ denote the steady-state relative price distortions that affect the production of ordinary and superstar firms, respectively. Given a value of the steady-state average markup μ, the steady-state values of the profit share, the labor share, and the investment share continue to be determined by eqs. (25)–(27), respectively.

To investigate the joint effect of the decline in trend inflation and the rise of superstar firms, the calibration presented in Table 2 is supplemented with values for the two new parameters, n and z. The value of superstar firms' relative productivity z is set to target the empirical ratio of the 90th percentile to the median markup. In their Table 3, Edmond et al. (2018) report that the former exceeded the latter by 50% in 2012, and thus we choose $z = 1.6$

[30] Autor et al. (2020) raise the hypothesis to explain the rising average markup and the declining labor share. Meanwhile, Kehrig and Vincent (2021) present evidence that supports a demand-driven explanation of the declining labor share.

[31] Appendix B presents the details of the model with superstar firms and proves the proposition that, for the two types of firms with product prices of the same age, a superstar firm has a higher markup in steady state than an ordinary firm if and only if the aggregator is the non-CES one, i.e., $\epsilon < 0$.

to target a ratio of the 90th percentile to the median markup of 1.5 at the trend inflation rate of 1.6% annually. The share of ordinary firms n is then set at its largest possible value that enables the ratio of the 90th percentile to the median markup, which is $n = 0.86$.

Panel (c) of Table 3 reports the joint effect of the decline in trend inflation and the rise of superstar firms on the average markup, the profit share, the labor share, and the investment share. The extended model predicts that these two factors increase the average markup by 8.2 percentage points and the profit share by 4.8 percentage points, while they decrease the labor share by 3.4 percentage points and the investment share by 1.1 percentage points. In short, the extended model attributes more than half of the increases in the average markup and the profit share and the decrease in the labor share, and more than one third of the decrease in the investment share to the two factors. Therefore, these factors together can account for the macroeconomic changes better than only the decline of trend inflation does.

To understand the joint effect of the two factors in more detail, Table 5 shows to what extent each of them affects the average, the median, and the 90th percentile markups. The decline in trend inflation (going from the second to the third row of the table) raises all three markups: the average and the median markups from 1.27 to 1.31 and the 90th percentile markup from 1.31 to 1.32. Then, adding the rise of superstar firms (going from the third to the fourth row) increases the average markup to 1.35 and the 90th percentile markup to 1.92 while decreasing the median markup to 1.28. Therefore, it raises the average markup further by increasing the 90th percentile; that is, the upper tail of the markup distribution rises in line with the micro evidence provided by, for example, De Loecker et al. (2020) (their Figure III).

Table 5: Effects of a decline in trend inflation and a rise of superstar firms.

Parameter values	Average	Median	p90
$\overline{\pi} = 5.6\%$ and $z = 1.0$	1.27	1.27	1.31
$\overline{\pi} = 1.6\%$ and $z = 1.0$	1.31	1.31	1.32
$\overline{\pi} = 1.6\%$ and $z = 1.6$	1.35	1.28	1.92

Notes: This table reports the effects of a decline in trend inflation and a rise of superstar firms on the average, the median, and the 90th percentile markups. The share of ordinary firms is set at $n = 0.86$, $\overline{\pi}\,(\equiv 400 \log \pi)$ and z denote the annualized trend inflation rate and superstar firms' relative productivity, and the values of other model parameters used here are reported in Table 2.

Moreover, the decline in trend inflation and the rise of superstar firms have offsetting effects on the median markup, thus keeping it roughly unchanged. The flat median markup is consistent with the micro evidence reported by Autor et al. (2020) (their Figure 10), who show that the median markup in manufacturing has been essentially flat over time, and De Loecker et al. (2020) (their Figure III), who indicate that the median markup among publicly traded firms is likewise invariant. Therefore, the two factors together can explain the empirical observation of the rising average markup and the flat median markup better than either does in isolation. In this way, the decline in trend inflation complements the rise of superstar firms in accounting for the empirical changes in the distribution of firms' markups (and the other macroeconomic variables).

Why does the rise in superstar firms' relative productivity z reduce the median markup? Facing the different marginal costs (32), ordinary and superstar firms that can adjust prices in period t set different prices $P^*_{1,t}$ and $P^*_{z,t}$, respectively. The steady-state distribution of firms' markups is represented by the density function

$$n f_1(j) + (1-n) f_z(j), \tag{34}$$

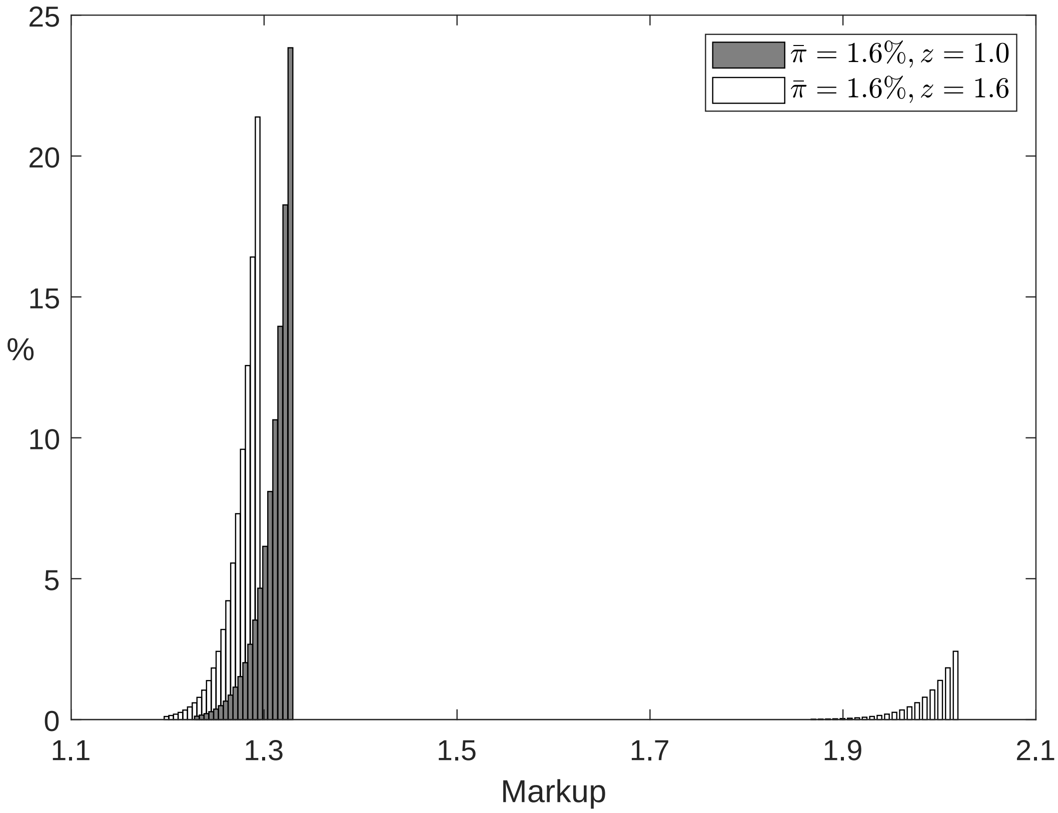
where

$$f_i(j) = (1-\xi)\,\xi^j \frac{mc_i}{mc_1\,\Delta_1 + mc_z\,\Delta_z}\left[\frac{1}{1+\epsilon}\left(\frac{d}{p_i^*}\right)^{\gamma}\pi^{\gamma j} + \frac{\epsilon}{1+\epsilon}\right],$$

for $i \in \{1, z\}$, is the cost-weighted proportion of ordinary firms ($i = 1$) or superstar firms ($i = z$) whose price was last updated j periods ago. The average markup can then be written as $\mu = \sum_{j=0}^{\infty}\{n\, f_1(j)[p_1^*/(mc_1\pi^j)] + (1-n)\, f_z(j)[p_z^*/(mc_z\pi^j)]\}$, where p_1^* and p_z^* are respectively the steady-state relative prices of ordinary and superstar firms that can change prices. The distribution shows that there are two markups associated with each price vintage j, $\{p_1^*/(mc_1\pi^j),\ p_z^*/(mc_z\pi^j)\}$. Figure 6 plots the steady-state markup distribution at the annualized trend inflation rate of $\bar{\pi} = 1.6\%$ for the cases of $z = 1$ (gray bars) and $z = 1.6$ (white bars). Comparing the two cases demonstrates that the rise of superstar firms (i.e., $z = 1.6$) leads the markup distribution to become more diffuse, by not only giving the distribution an upper tail but also increasing the fraction of firms with low markups. On balance these two effects then lead to a lower median markup. Under the calibration presented in Table 2, it can be verified that the presence of superstar firms raises ordinary firms' steady-state marginal cost, that is, $mc_1 > mc$. Because eq. (12) for mc_t and eq. (32)

for $mc_t(f)$, $f \in [0, n]$ are identical, and because the steady-state capital rental rate r_k is the same for any value of z, it follows that the larger value of $z = 1.6$ induces a larger steady-state real wage rate w. Therefore, the rise of superstar firms drives up the marginal cost, thereby lowering ordinary firms' markups and hence the median markup.

Figure 6: Steady-state distribution of markups across two types of firms.



Note: The figure illustrates the distribution of ordinary and superstar firms' markups with cost weights in the steady states with the annualized trend inflation rate of $\bar{\pi} \equiv 400 \log \pi = 1.6\%$ and superstar firms' relative productivity of $z = 1.6$ (white bars) and $z = 1.0$ (gray bars) in the model (with the non-CES aggregator). The plotted distributions display the 20 most recent price vintages, which account for more than 99% of the probability mass. The share of ordinary firms is set at $n = 0.86$ and the values of other model parameters used here are reported in Table 2.

6 Concluding remarks

Since 1980, the US economy has undergone increases in the average markup and the profit share of income and decreases in the labor share and the investment share of spending. In tandem with these macroeconomic changes, inflation has trended down steadily. Thus, this paper has examined the role of monetary policy in the macroeconomic changes using a simple staggered price model with a non-CES aggregator of individual differentiated goods, which introduces endogenous variation in firms' desired markups and leads the average markup to rise for lower trend inflation. The calibrated model has shown that a decline of trend inflation as measured since 1980 can account for a substantial portion of the changes.[32] Moreover, adding a rise of highly productive superstar firms to the model can better account for not only the macroeconomic changes but also the micro evidence on the distribution of firms' markups.

[32]This result implies that zero trend inflation—which the previous literature reviewed by, for example, Schmitt-Grohé and Uribe (2010) considers to be the optimal inflation rate when nominal price rigidity is the main source of monetary non-neutrality—leads to a high average markup, that is, a large distortion in the model. In this context, Kurozumi and Van Zandweghe (2020) propose variable elasticity of demand as a new rationale for the positive inflation targets of central banks.

Appendix

A Details of robustness analysis

This appendix describes three sets of robustness exercises. The first set addresses the specification of nominal price rigidity. The second raises some key features of medium-scale models that are absent from our model. The third pertains to time variation in trend inflation.

A.1 Specification of nominal price rigidity

Two alternative specifications of nominal price rigidity are a fixed duration staggered contract of Taylor (1980) and a price adjustment cost of Rotemberg (1982).

As in Taylor (1980), every firm sets its product price for J periods, and each period a fraction $1/J$ of firms adjust their prices. Hence the firms choose the price $P_t(f)$ so as to maximize the relevant profit

$$E_t \sum_{j=0}^{J-1} Q_{t,t+j} \left(P_t(f) - P_{t+j}\, mc_{t+j}\right) \frac{Y_{t+j}}{1+\epsilon} \left[\left(\frac{P_t(f)}{P_{t+j}\, d_{t+j}}\right)^{-\gamma} + \epsilon\right].$$

The first-order condition for profit maximization (13) is then replaced by

$$E_t \sum_{j=0}^{J-1} \beta^j \frac{Y_{t+j}}{C_{t+j}} \left[\left(\frac{p_t^*}{d_{t+j}}\right)^{-\gamma} \prod_{k=1}^{j} \pi_{t+k}^{\gamma} \left(p_t^* \prod_{k=1}^{j} \pi_{t+k}^{-1} - \frac{\gamma}{\gamma-1} mc_{t+j}\right) - \frac{\epsilon}{\gamma-1} p_t^* \prod_{k=1}^{j} \pi_{t+k}^{-1}\right] = 0.$$

In addition, the equations for the real marginal cost of producing the composite good d_t in eq. (14), the average relative price e_t in eq. (15), and the dynamic component of the relative price distortion s_t in eq. (19) are replaced with

$$d_t^{1-\gamma} = \frac{1}{J} \sum_{j=0}^{J-1} \left(p_{t-j}^* \prod_{\tau=0}^{j-1} \pi_{t-\tau}^{-1}\right)^{1-\gamma}, \quad e_t = \frac{1}{J} \sum_{j=0}^{J-1} p_{t-j}^* \prod_{\tau=0}^{j-1} \pi_{t-\tau}^{-1}, \quad d_t^{-\gamma} s_t = \frac{1}{J} \sum_{j=0}^{J-1} \left(p_{t-j}^* \prod_{\tau=0}^{j-1} \pi_{t-\tau}^{-1}\right)^{-\gamma}.$$

Consequently, the equations for the steady-state values of real marginal cost and relative price distortion in eqs. (22) and (23) become

$$mc = \frac{\gamma - 1 - \tilde{\epsilon}}{\gamma} \frac{1-\beta\pi^{\gamma}}{1-\beta\pi^{\gamma-1}} \frac{1-(\beta\pi^{\gamma-1})^J}{1-(\beta\pi^{\gamma})^J} \left[\frac{1}{1+\epsilon} \left(\frac{1}{J} \frac{1-\pi^{(\gamma-1)J}}{1-\pi^{\gamma-1}}\right)^{\frac{1}{1-\gamma}} + \frac{\epsilon}{1+\epsilon} \frac{1}{J} \frac{1-\pi^{-J}}{1-\pi^{-1}}\right]^{-1},$$

$$\Delta = \frac{1}{1+\epsilon}\frac{1}{J}\frac{1-\pi^{\gamma J}}{1-\pi^{\gamma}}\left(\frac{1}{J}\frac{1-\pi^{(\gamma-1)J}}{1-\pi^{\gamma-1}}\right)^{\frac{\gamma}{1-\gamma}} + \frac{\epsilon}{1+\epsilon},$$

where $\tilde{\epsilon} \equiv \epsilon\,[J(1-\pi^{\gamma-1})/(1-\pi^{(\gamma-1)J})]^{\gamma/(1-\gamma)}[(1-\beta\pi^{\gamma-1})/(1-\beta\pi^{-1})]\{[1-(\beta\pi^{-1})^J]/[1-(\beta\pi^{\gamma-1})^J]\}$.

We turn next to a quadratic price adjustment cost. As in Rotemberg (1982), such a cost implies that all firms set the price $P_t(f)$ in every period so as to maximize the profit

$$E_0\sum_{t=0}^{\infty} Q_{0,t}\left\{(P_t(f)-P_t\,mc_t)\,\frac{Y_t}{1+\epsilon}\left[\left(\frac{P_t(f)}{P_t\,d_t}\right)^{-\gamma}+\epsilon\right]-\frac{\zeta}{2}P_tY_t\left(\frac{P_t(f)}{P_{t-1}(f)}-1\right)^2\right\},$$

where the parameter $\zeta > 0$ governs the size of the price adjustment cost. The first-order condition for profit maximization (13) is replaced by

$$\begin{aligned}0 = {} & \left(\frac{p_t^*}{d_t}\right)^{\gamma}\left(p_t^*-\frac{\gamma}{\gamma-1}mc_t\right)-\frac{\epsilon}{\gamma-1}p_t^* \\ & +\frac{\zeta(1+\epsilon)}{\gamma-1}\left\{\frac{p_t^*}{p_{t-1}^*}\,\pi_t\left(\frac{p_t^*}{p_{t-1}^*}\,\pi_t-1\right)-\beta E_t\left[\frac{C_t}{C_{t+1}}\frac{Y_{t+1}}{Y_t}\frac{p_{t+1}^*}{p_t^*}\,\pi_{t+1}\left(\frac{p_{t+1}^*}{p_t^*}\,\pi_{t+1}-1\right)\right]\right\},\end{aligned}$$

and the symmetry across firms implies that eqs. (14), (15), and (19) are replaced with $d_t = e_t = s_t = p_t^* = 1$. Consequently, the equations for the steady-state values of real marginal cost and relative price distortion become

$$mc = \frac{\gamma-1-\epsilon+\zeta(1+\epsilon)(1-\beta)\pi(\pi-1)}{\gamma}, \quad \Delta = 1.$$

The equations for the steady-state values of the average markup, the labor share, and the investment share remain unchanged—eqs. (24), (26), and (27)—for both specifications of nominal price rigidity. The equation for the steady-state profit share is also unchanged from eq. (25) for the fixed duration staggered contract, whereas the price adjustment cost implies that such a cost is subtracted from the profit share, so that $j/y = 1-1/\mu-(\zeta/2)(\pi-1)^2$.

As for the calibration of parameters pertaining to nominal price rigidity, Dixon and Kara (2006) suggest comparing models with Taylor-style fixed duration staggered contracts and with Calvo-style random duration ones that have the same average age of contracts, that is, $J = (1+\xi)/(1-\xi)$, so that $J = 7$ periods correspond to the probability of $\xi = 0.75$. Calibrating the price adjustment cost parameter ζ is less straightforward. We choose it such that in the steady state with the annualized trend inflation rate of 1.6%, the price adjustment cost is equal to the relative price distortion in the baseline model, which is $\Delta = 1.0002$. This yields a value of $\zeta = 25$.

A.2 Medium-scale model features

In this subsection we augment the baseline model with either staggered wage-setting, a roundabout production structure, or a fixed cost of production.

To incorporate staggered wage-setting as in Erceg et al. (2000), we assume that the representative household has a continuum of members $h \in [0, 1]$, each of whom supplies a differentiated labor service $l_t(h)$, and derives disutility from the labor services. Consequently, the household's preferences are represented as

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[\log\left(C_t\right) - \int_0^1 \frac{l_t(h)^{1+1/\chi}}{1+1/\chi} \, dh \right].$$

In each period, a fraction $\xi_w \in (0, 1)$ of nominal wages remains unchanged, while the other wages are set so as to maximize the relevant utility function

$$E_t \sum_{j=0}^{\infty} (\beta \xi_w)^j \left[-\frac{l_{t+j|t}(h)^{1+1/\chi}}{1+1/\chi} + \Lambda_{t+j} \frac{P_t \, W_t(h)}{P_{t+j}} \, l_{t+j|t}(h) \right]$$

subject to the labor demand curve

$$l_{t+j|t}(h) = \frac{l_{t+j}}{1+\epsilon_w} \left[\left(\frac{P_t W_t(h)}{P_{t+j} W_{t+j} \, d_{w,t+j}} \right)^{-\gamma_w} + \epsilon_w \right],$$

which arises from a Kimball-type aggregator that takes the same functional form as the goods aggregator $F(\cdot)$ but with parameters θ_w and ϵ_w. Here Λ_t denotes the marginal utility of consumption, $\gamma_w \equiv \theta_w(1 + \epsilon_w)$, and

$$d_{w,t} = \left[\int_0^1 \left(\frac{P_t W_t(h)}{P_t W_t} \right)^{1-\gamma_w} dh \right]^{\frac{1}{1-\gamma_w}}$$

is the real marginal cost of aggregating differentiated labor services and meets

$$1 = \frac{1}{1+\epsilon_w} \, d_{w,t} + \frac{\epsilon_w}{1+\epsilon_w} \, e_{w,t} \,,$$

where

$$e_{w,t} \equiv \int_0^1 \frac{P_t W_t(h)}{P_t W_t} \, dh.$$

The first-order condition for utility maximization with respect to the wage is given by

$$0 = E_t \sum_{j=0}^{\infty} (\beta \xi_w)^j \frac{\Lambda_{t+j}}{\Lambda_t} l_{t+j} \left[\left(\frac{W_t^*/W_t}{d_{w,t+j}} \right)^{-\gamma_w} \prod_{\tau=1}^{j} \pi_{w,t+\tau}^{\gamma_w} \left(\frac{W_t^*}{W_t} \prod_{\tau=1}^{j} \pi_{t+\tau}^{-1} \right. \right.$$
$$\left. \left. - \frac{\gamma_w}{\gamma_w - 1} \left\{ \frac{l_{t+j}}{1+\epsilon_w} \left[\left(\frac{W_t^*/W_t}{d_{w,t+j}} \right)^{-\gamma_w} \prod_{\tau=1}^{j} \pi_{w,t+\tau}^{\gamma_w} + \epsilon_w \right] \right\}^{1/\chi} \frac{1}{\Lambda_{t+j} W_{t+j}} \prod_{\tau=1}^{j} \frac{W_{t+\tau}}{W_{t+\tau-1}} \right) - \frac{\epsilon_w}{\gamma_w - 1} \frac{W_t^*}{W_t} \prod_{\tau=1}^{j} \pi_{t+\tau}^{-1} \right],$$

where $\pi_{w,t} = \pi_t W_t / W_{t-1}$ is the wage inflation rate. The last five equations, including the definition of wage inflation, jointly replace our model's labor supply condition (2), while the other equilibrium conditions remain unchanged. Then, the equations for the steady-state values of the four macroeconomic variables are the same as eqs. (24)–(27), because they do not depend on the specification of the labor disutility function.

We turn next to a roundabout production structure. To embed it as in Basu (1995), we extend the production function (10) to

$$X_t(f) = A_t \left[K_t(f)^{\alpha} l_t(f)^{1-\alpha} \right]^{1-\phi} O_t(f)^{\phi},$$

where $X_t(f)$ is the output of firm f, $O_t(f)$ is its intermediate input, $\phi \in (0,1)$ is the intermediate-input elasticity of output, and the technology level grows at a constant rate $A_t/A_{t-1} = g^{(1-\alpha)(1-\phi)}$. Then, the equations for the steady-state values of real marginal cost, relative price distortion, and hence the average markup are the same as eqs. (22)–(24) because they do not depend on the specification of the production function, whereas those of the profit share, the labor share, and the investment share are now given by

$$\frac{j}{y} = 1 - \frac{1-\phi}{\mu - \phi}, \quad \frac{wl}{y} = (1-\alpha) \frac{1-\phi}{\mu - \phi}, \quad \frac{i}{y} = \alpha \beta \frac{g - (1-\delta)}{g - \beta(1-\delta)} \frac{1-\phi}{\mu - \phi},$$

which replace eqs. (25)–(27). As for the calibration of the intermediate-input elasticity ϕ, its value is set at $\phi = 0.5$, following Basu (1995).

Last, the presence of a fixed cost in production extends the production function (10) to

$$Y_t(f) = A_t K_t(f)^{\alpha} l_t(f)^{1-\alpha} - \omega \Upsilon_t, \tag{35}$$

if $A_t K_t(f)^{\alpha} l_t(f)^{1-\alpha} > \omega \Upsilon_t$; otherwise, $Y_t(f) = 0$, where $\omega \Upsilon_t > 0$ denotes the fixed cost. While this cost does not alter the equations for the steady-state values of real marginal cost and relative price distortion (22) and (23), it replaces those of the four macroeconomic variables (24)–(27) with

$$\mu = \frac{1+(\omega/y)e}{mc\,(\Delta+\omega/y)}, \quad \frac{j}{y} = 1 - \frac{1+(\omega/y)e}{\mu}, \quad \frac{wl}{y} = (1-\alpha)\frac{1+(\omega/y)e}{\mu},$$
$$\frac{i}{y} = \alpha\beta\left(\frac{g-(1-\delta)}{g-\beta(1-\delta)}\right)\frac{1+(\omega/y)e}{\mu},$$

where

$$e = \frac{1-\xi}{1-\xi\pi^{-1}}\left[\frac{1}{1+\epsilon}\left(\frac{1-\xi}{1-\xi\pi^{\gamma-1}}\right)^{\frac{1}{1-\gamma}} + \frac{\epsilon}{1+\epsilon}\frac{1-\xi}{1-\xi\pi^{-1}}\right]^{-1}$$

is the steady-state average relative price. The fixed cost is calibrated to target the observed profit share of the US economy in 2019 (i.e., 13.6%), which implies the value of $\omega/y = 0.133$.

A.3 Time variation in trend inflation

In this subsection we present the log-linearized equilibrium conditions of the model that is augmented with time-varying trend inflation, and describe the details of the counterfactual exercise.

The same equilibrium conditions as those for our model in Section 3 apply, in addition to the trend inflation process (28), the consumption Euler equation (29), and the monetary policy rule (30). To log-linearize the price-setting condition (13), we rewrite it as the four equations

$$p_t^* V_{1,t} + (p_t^*)^{1+\gamma} V_{2,t} = V_{3,t}, \quad V_{1,t} = \frac{y_t}{c_t} d_t^\gamma + \beta\xi E_t\left(\pi_{t+1}^{\gamma-1} V_{1,t+1}\right),$$
$$V_{2,t} = -\frac{\epsilon}{\gamma-1}\frac{y_t}{c_t} + \beta\xi E_t\left(\pi_{t+1}^{-1} V_{2,t+1}\right), \quad V_{3,t} = \frac{\gamma}{\gamma-1}\frac{y_t}{c_t} d_t^\gamma mc_t + \beta\xi E_t\left(\pi_{t+1}^{\gamma} V_{3,t+1}\right).$$

The complete set of log-linearized conditions is then given by

$$\left(1-\tilde{\epsilon}\frac{1+\gamma}{\gamma-1}\right)\hat{p}_t^* = -\hat{V}_{1,t} + \frac{\tilde{\epsilon}}{\gamma-1}\hat{V}_{2,t} + \left(1-\frac{\tilde{\epsilon}}{\gamma-1}\right)\hat{V}_{3,t},$$
$$\hat{V}_{1,t} = \left(1-\beta\xi\pi^{\gamma-1}\right)\left(\hat{y}_t - \hat{c}_t + \gamma\hat{d}_t\right) + \beta\xi\pi^{\gamma-1}(\gamma-1)E_t\hat{\pi}_{t+1} + \beta\xi\pi^{\gamma-1}E_t\hat{V}_{1,t+1},$$
$$\hat{V}_{2,t} = \left(1-\beta\xi\pi^{-1}\right)(\hat{y}_t-\hat{c}_t) - \beta\xi\pi^{-1}E_t\hat{\pi}_{t+1} + \beta\xi\pi^{-1}E_t\hat{V}_{2,t+1},$$
$$\hat{V}_{3,t} = (1-\beta\xi\pi^\gamma)\left(\hat{y}_t - \hat{c}_t + \gamma\hat{d}_t + \hat{mc}_t\right) + \beta\xi\pi^\gamma\gamma E_t\hat{\pi}_{t+1} + \beta\xi\pi^\gamma E_t\hat{V}_{3,t+1},$$
$$\hat{w}_t + \hat{l}_t = \hat{r}_{k,t} + \hat{k}_{t-1}, \quad \hat{mc}_t = (1-\alpha)\hat{w}_t + \alpha\hat{r}_{k,t}, \quad \hat{y}_t + \hat{\Delta}_t = (1-\alpha)\hat{l}_t + \alpha\hat{k}_{t-1},$$
$$\hat{\Delta}_t = \xi\pi^\gamma\hat{\Delta}_{t-1} + \frac{s}{s+\epsilon}\frac{\xi\pi^{\gamma-1}\gamma(\pi-1)}{1-\xi\pi^{\gamma-1}}\left(\hat{\pi}_t + \hat{d}_t - \hat{d}_{t-1}\right), \quad \hat{p}_t^* = \frac{\xi\pi^{\gamma-1}}{1-\xi\pi^{\gamma-1}}\left(\hat{\pi}_t + \frac{1}{\xi\pi^{\gamma-1}}\hat{d}_t - \hat{d}_{t-1}\right),$$

$$\hat{d}_t = \frac{\xi\pi^{-1}(1+\epsilon_1\pi^{\gamma})}{1+\epsilon_1}\hat{d}_{t-1} - \frac{\epsilon_1\xi\pi^{-1}(\pi^{\gamma}-1)}{(1+\epsilon_1)(1-\xi\pi^{-1})}\hat{\pi}_t, \quad \hat{w}_t = \frac{1}{\chi}\hat{l}_t + \hat{c}_t,$$
$$\hat{k}_t = \frac{1-\delta}{g}\hat{k}_{t-1} + \left(1-\frac{1-\delta}{g}\right)\hat{i}_t, \quad E_t\hat{c}_{t+1} - \hat{c}_t = \left[1 - \frac{\beta(1-\delta)}{g}\right]E_t\hat{r}_{k,t+1},$$
$$E_t\hat{c}_{t+1} - \hat{c}_t = \hat{r}_t - E_t\hat{\pi}_{t+1}, \quad \hat{y}_t = \frac{c}{y}\hat{c}_t + \frac{i}{y}\hat{i}_t, \quad \hat{r}_t = \phi_\pi\left(\hat{\pi}_t - \hat{\pi}^*_t\right), \quad \hat{\pi}^*_t = \rho_\pi\hat{\pi}^*_{t-1} + \varepsilon_t,$$
$$\hat{\mu}_t = -\hat{mc}_t - \hat{\Delta}_t, \quad \hat{js}_t = \frac{1}{\mu-1}\hat{\mu}_t, \quad \hat{ls}_t = -\hat{\mu}_t, \quad \hat{is}_t = \hat{i}_t - \hat{y}_t,$$

where hatted variables denote log-deviations from their steady-state values (e.g., $\hat{y}_t = \log y_t - \log y$) and the following definitions are employed: $js_t \equiv j_t/y_t$, $ls_t \equiv l_t/y_t$, $is_t \equiv i_t/y_t$, and $\epsilon_1 \equiv \epsilon[(1-\xi)/(1-\xi\pi^{\gamma-1})]^{\gamma/(\gamma-1)}$.

To calibrate the model, the autoregressive process $x_t = \beta_0 + \beta_1 x_{t-1} + e_t$ is estimated by OLS for the trend inflation series of Chan et al. (2018) during the period from 1960Q2 to 2019Q4. The OLS estimators of β_0 and β_1 are computed as $b_0 = 0.000$ and $b_1 = 0.996$, and the Newey-West robust standard errors are $s_0 = 0.000$ and $s_1 = 0.011$, respectively. To conduct the counterfactual exercise, initial values of the four state variables $(\pi^*_{t-1}, \Delta_{t-1}, d_{t-1}, k_{t-1})$ must be pinned down. The initial value of the lagged inflation target is set equal to the trend inflation rate in 1960Q2 in the series of Chan et al. (2018), while those of the other three (endogenous) state variables are set at their steady-state values.[33] Given the initial values, the trend inflation series is replicated, and the associated counterfactual series for the average markup, the profit share, the labor share, and the investment share are generated, by feeding the OLS residuals $\{x_t - b_0 - b_1 x_{t-1}\}_{t=1960Q3}^{2019Q4}$ into the calibrated model.

B Model with superstar firms

This appendix describes the model with superstar firms, which extends the baseline model presented in Section 3 by considering two types of firms: ordinary firms $f \in [0, n]$, whose type is indexed by subscript $i = 1$, and highly productive superstar firms $f \in (n, 1]$, whose type is indexed by subscript $i = z$. The decision problems of the representative household and the representative composite-good producer remain unchanged in the extended model.

Firms' production function (10) is extended to the form (31), so that if $z = 1$ then the

[33]Varying the initial values of the three endogenous state variables by $\pm 5\%$ does not alter the qualitative comparison to the results obtained with the baseline model (as reported in panel (b) of Table 3).

extended model is reduced to the baseline model. All firms continue to choose an identical capital–labor ratio, so that

$$\frac{\alpha}{1-\alpha}\frac{W_t}{r_{k,t}} = \frac{K_{1,t-1}}{l_{1,t}} = \frac{K_{z,t-1}}{l_{z,t}},$$

where $K_{1,t-1} = \int_0^n K_t(f)\, df$, $l_{1,t} = \int_0^n l_t(f)\, df$, $K_{z,t-1} = \int_n^1 K_t(f)\, df$, and $l_{z,t} = \int_n^1 l_t(f)\, df$. In addition, the real marginal cost is given by the extended form (32); we denote $mc_{1,t} = mc_t(f)$ for $f \in [0, n]$ and $mc_{z,t} = mc_t(f)$ for $f \in (n, 1]$.

The first-order condition of the price-setting problem for each firm type $i \in \{1, z\}$ can be written as

$$E_t \sum_{j=0}^{\infty} (\beta\xi)^j \frac{Y_{t+j}}{C_{t+j}} \left[\left(\frac{p^*_{i,t}}{d_{t+j}} \right)^{-\gamma} \prod_{k=1}^{j} \pi^{\gamma}_{t+k} \left(p^*_{i,t} \prod_{k=1}^{j} \pi^{-1}_{t+k} - \frac{\gamma}{\gamma-1} mc_{i,t+j} \right) - \frac{\epsilon}{\gamma-1} p^*_{i,t} \prod_{k=1}^{j} \pi^{-1}_{t+k} \right] = 0,$$

where $p^*_{i,t} \equiv P^*_{i,t}/P_t$ and $P^*_{i,t}$ is the price set by firms that can change prices in period t. Moreover, the law of motion (14) for d_t is extended to

$$\begin{aligned}
d_t^{1-\gamma} &= d_{1,t}^{1-\gamma} + d_{z,t}^{1-\gamma}, \\
d_{1,t}^{1-\gamma} &= \xi\, \pi_t^{\gamma-1}\, d_{1,t-1}^{1-\gamma} + (1-\xi) n \left(p^*_{1,t}\right)^{1-\gamma}, \\
d_{z,t}^{1-\gamma} &= \xi\, \pi_t^{\gamma-1}\, d_{z,t-1}^{1-\gamma} + (1-\xi)(1-n) \left(p^*_{z,t}\right)^{1-\gamma},
\end{aligned}$$

where $d_{1,t} \equiv \left[\int_0^n (P_t(f)/P_t)^{1-\gamma}\, df\right]^{1/(1-\gamma)}$ and $d_{z,t} \equiv \left[\int_n^1 (P_t(f)/P_t)^{1-\gamma}\, df\right]^{1/(1-\gamma)}$. Similarly, the law of motion (15) for e_t is extended to

$$\begin{aligned}
e_t &= e_{1,t} + e_{z,t}, \\
e_{1,t} &= \xi\, \pi_t^{-1}\, e_{1,t-1} + (1-\xi) n\, p^*_{1,t}, \\
e_{z,t} &= \xi\, \pi_t^{-1}\, e_{z,t-1} + (1-\xi)(1-n)\, p^*_{z,t},
\end{aligned}$$

where $e_{1,t} \equiv \int_0^n (P_t(f)/P_t)\, df$ and $e_{z,t} \equiv \int_n^1 (P_t(f)/P_t)\, df$.

Aggregating the production function (31) for ordinary firms $f \in [0, n]$ and for superstar firms $f \in (n, 1]$ leads to, respectively,

$$Y_t\, \Delta_{1,t} = A_t\, l_{1,t}^{1-\alpha} K_{1,t-1}^{\alpha}, \tag{36}$$

$$Y_t\, \Delta_{z,t} = A_t\, l_{z,t}^{1-\alpha} K_{z,t-1}^{\alpha}, \tag{37}$$

where

$$\Delta_{1,t} = \frac{s_{1,t} + n\epsilon}{1+\epsilon}, \tag{38}$$

$$\Delta_{z,t} = \frac{s_{z,t} + (1-n)\epsilon}{1+\epsilon} \tag{39}$$

denote the relative price distortions that respectively affect the production of ordinary and superstar firms, $s_{1,t} \equiv \int_0^n \left[P_t(f)/(P_t\, d_t)\right]^{-\gamma} df$, and $s_{z,t} \equiv \int_n^1 \left[P_t(f)/(P_t\, d_t)\right]^{-\gamma} df$. Therefore, the aggregate production function (18) is replaced with ordinary and superstar firms' aggregate production functions (36) and (37), and the relative price distortion (16) is replaced with the two relative price distortions (38) and (39). The laws of motion for $s_{1,t}$ and $s_{z,t}$ are given by

$$d_t^{-\gamma} s_{1,t} = \xi \pi_t^{\gamma} d_{t-1}^{-\gamma} s_{1,t-1} + (1-\xi) n \left(p_{1,t}^*\right)^{-\gamma},$$
$$d_t^{-\gamma} s_{z,t} = \xi \pi_t^{\gamma} d_{t-1}^{-\gamma} s_{z,t-1} + (1-\xi)(1-n) \left(p_{z,t}^*\right)^{-\gamma},$$

which replace the law of motion (19) for s_t.

The cost-weighted average markup can then be calculated as

$$\mu_t = \int_0^1 \frac{TC_t(f)}{\int_0^1 TC_t(f)\, df} \frac{P_t(f)}{P_t\, mc_t(f)} df = \frac{1}{mc_{1,t}\, \Delta_{1,t} + mc_{z,t}\, \Delta_{z,t}}.$$

The equilibrium conditions in terms of detrended variables lead to the relevant steady-state equations

$$mc_i = \frac{\gamma-1}{\gamma} \frac{1-\beta\xi\pi^{\gamma}}{1-\beta\xi\pi^{\gamma-1}} \left[1 - \frac{\epsilon}{\gamma-1}\left(\frac{p_i^*}{d}\right)^{\gamma} \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{-1}}\right] p_i^*,\ i \in \{1, z\};$$
$$\Delta_1 = n\left[\frac{1}{1+\epsilon}\frac{1-\xi}{1-\xi\pi^{\gamma}}\left(\frac{d}{p_1^*}\right)^{\gamma} + \frac{\epsilon}{1+\epsilon}\right]; \quad \Delta_z = (1-n)\left[\frac{1}{1+\epsilon}\frac{1-\xi}{1-\xi\pi^{\gamma}}\left(\frac{d}{p_z^*}\right)^{\gamma} + \frac{\epsilon}{1+\epsilon}\right];$$
$$d^{1-\gamma} = d_1^{1-\gamma} + d_z^{1-\gamma}; \quad d_1 = \left[\frac{(1-\xi)n}{1-\xi\pi^{\gamma-1}}\right]^{\frac{1}{1-\gamma}} p_1^*; \quad d_z = \left[\frac{(1-\xi)(1-n)}{1-\xi\pi^{\gamma-1}}\right]^{\frac{1}{1-\gamma}} p_z^*;$$
$$e = e_1 + e_z; \quad e_1 = \frac{(1-\xi)n}{1-\xi\pi^{-1}} p_1^*; \quad e_z = \frac{(1-\xi)(1-n)}{1-\xi\pi^{-1}} p_z^*; \quad 1 = \frac{1}{1+\epsilon} d + \frac{\epsilon}{1+\epsilon} e\,.$$

Combining the steady-state equations yields the following three nonlinear equations for the three steady-state variables $\{d,\, p_1^*,\, p_z^*\}$:

$$1 = \frac{1}{1+\epsilon} d + \frac{\epsilon}{1+\epsilon} \frac{1-\xi}{1-\xi\pi^{-1}} \left[n p_1^* + (1-n) p_z^*\right],$$

$$d^{1-\gamma} = \frac{(1-\xi)n}{1-\xi\pi^{\gamma-1}} (p_1^*)^{1-\gamma} + \frac{(1-\xi)(1-n)}{1-\xi\pi^{\gamma-1}} (p_z^*)^{1-\gamma} \ ,$$

$$\left[\gamma - 1 - \epsilon \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{-1}} \left(\frac{p_1^*}{d}\right)^{\gamma}\right] p_1^* = \left[\gamma - 1 - \epsilon \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{-1}} \left(\frac{p_z^*}{d}\right)^{\gamma}\right] (z\, p_z^*) \, , \qquad (40)$$

which can be solved numerically using the values of model parameters presented in Table 2, $z = 1.6$, and $n = 0.86$. The solution allows us to calculate the steady-state real marginal costs mc_1, mc_z; the steady-state relative price distortions Δ_1, Δ_z; the steady-state average markup (33); the steady-state profit share (25); the steady-state labor share (26); the steady-state investment share (27); and the density function for the steady-state markup distribution (34).

In the model with superstar firms the non-CES aggregator plays a dual role: it serves as not only a source of endogenous variation of firms' desired markups but also a source of markup heterogeneity between firms with different productivity levels. The following proposition shows that, for the two types of firms with product prices of the same age, a superstar firm has a higher markup in steady state than an ordinary firm if and only if the aggregator is the non-CES one, i.e., $\epsilon < 0$.

Proposition 1 *Assume that the assumption (21) holds and that $z > 1$. Consider ordinary firms and superstar firms whose product prices have remained unchanged for j periods. Then, the steady-state markup of the superstar firms exceeds that of the ordinary firms if and only if $\epsilon < 0$.*

Proof. Let the steady-state markup of a firm of type $i \in \{1, z\}$ with a price of age j be denoted by $\mu_{i,j} = p_i^* / (mc_i\, \pi^j)$. The proposition then claims that $\mu_{1,j} < \mu_{z,j}$ iff $\epsilon < 0$.

First, assume $\epsilon < 0$ and suppose the contrary $\mu_{1,j} \geq \mu_{z,j}$. Then, $mc_1 = z\, mc_z$ implies that $p_1^* \geq z\, p_z^* > p_z^*$. Without loss of generality, we have $\gamma - 1 - \epsilon \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{-1}} \left(\frac{p_1^*}{d}\right)^{\gamma} = \gamma \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{\gamma}} \frac{mc_1}{p_1^*} \neq 0$ under the assumption (21). Then, from eq. (40), it follows that

$$\frac{\gamma - 1 - \epsilon \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{-1}} \left(\frac{p_z^*}{d}\right)^{\gamma}}{\gamma - 1 - \epsilon \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{-1}} \left(\frac{p_1^*}{d}\right)^{\gamma}} (z\, p_z^*) = p_1^* \geq z\, p_z^* \quad \Leftrightarrow \quad \frac{\gamma - 1 - \epsilon \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{-1}} \left(\frac{p_z^*}{d}\right)^{\gamma}}{\gamma - 1 - \epsilon \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{-1}} \left(\frac{p_1^*}{d}\right)^{\gamma}} \geq 1.$$

If $\gamma - 1 - \epsilon \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{-1}} \left(\frac{p_1^*}{d}\right)^{\gamma} = \gamma \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{\gamma}} \frac{mc_1}{p_1^*} > 0$, we have that $\gamma > 0$ and $(p_z^*)^{\gamma} \geq (p_1^*)^{\gamma}$ under the assumption (21). Then, it follows that $(p_z^*)^{\gamma} \geq (p_1^*)^{\gamma} > (p_z^*)^{\gamma}$, which is a contradiction. If $\gamma - 1 - \epsilon \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{-1}} \left(\frac{p_1^*}{d}\right)^{\gamma} = \gamma \frac{1-\beta\xi\pi^{\gamma-1}}{1-\beta\xi\pi^{\gamma}} \frac{mc_1}{p_1^*} < 0$, we have that $\gamma < 0$ and $(p_z^*)^{\gamma} \leq (p_1^*)^{\gamma}$ under the assumption (21). Then, it follows that $(p_z^*)^{-\gamma} \geq (p_1^*)^{-\gamma} > (p_z^*)^{-\gamma}$, which is a contradiction.

Next, assume that $\mu_{1,j} < \mu_{z,j}$ and suppose the contrary $\epsilon = 0$. Then, eq. (40) can be reduced, without loss of generality, to $p_1^* = z\, p_z^*$. From $mc_1 = z\, mc_z$, it follows that $\mu_{1,j} = \mu_{z,j} > \mu_{1,j}$, which is a contradiction. ■

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