Nonlinear Dynamics in Menu Cost Economies? Evidence from U.S. Data*

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Abstract

We show that standard menu cost models cannot simultaneously reproduce the dispersion in the size of micro-price changes and the extent to which the fraction of price changes increases with inflation in the U.S. time-series. Though the Golosov and Lucas (2007) model generates fluctuations in the fraction of price changes, it predicts too little dispersion in the size of price changes and therefore little monetary non-neutrality. In contrast, versions of the model that reproduce the dispersion in the size of price changes and generate stronger monetary non-neutrality predict a nearly constant fraction of price changes.

Keywords: menu costs, inflation, fraction of price changes.

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1 Introduction

A robust feature of the data is that the fraction of price changes increases in periods of high inflation.[1] It is widely believed that menu cost models can reproduce this pattern because firms choose endogenously the timing of price changes and are more likely to respond to larger shocks. Whether this is indeed the case or not has important implications for the extent to which the slope of the Phillips curve varies in the time-series and therefore the tradeoff between inflation and output stabilization faced by monetary policy.[2]

Because in menu cost models the distribution of price changes also critically shapes the slope of the Phillips curve and the degree of monetary non-neutrality,[3] we evaluate the ability of standard menu cost models to reproduce the comovement between inflation and the fraction of price changes in the U.S. data, while simultaneously accounting for the large dispersion in the size of price changes. We consider three commonly-used specifications for the technology of changing prices: *(i)* a fixed menu cost, as in Golosov and Lucas (2007), *(ii)* a Calvo-plus specification in which with some probability firms can change prices for free, and otherwise need to pay a fixed cost, as in Nakamura and Steinsson (2010) and *(iii)* a Calvo-plus specification augmented with a uniform distribution of menu costs, as in Blanco et al. (2024a).

We use the Krusell and Smith (1998) approach to characterize the non-linear solution of each menu cost economy in the presence of aggregate monetary policy shocks and then back out the sequence of shocks that exactly reproduces the time path of inflation in the U.S. time series. We find that the Golosov and Lucas (2007) model predicts that the fraction of price changes increases considerably in periods of high inflation, as first pointed out by Nakamura et al. (2018), though less than in the data.[4] As is well understood, however, this model predicts too little dispersion in the size of price changes. In contrast, the other two variants of the model that better reproduce the dispersion in the size of price changes imply a nearly constant fraction of price changes.

We illustrate the implications of these predictions by studying how each of these economies

[1] See, for example, Gagnon (2009), Alvarez et al. (2018), Karadi and Reiff (2019), Blanco et al. (2024a) and Cavallo et al. (2024).

[2] See Blanco et al. (2024b) who show that a model that reproduces the increase in the fraction of price changes in periods of high inflation implies a sharp increase in the slope of the Phillips curve in those periods, considerably reducing the output costs of reducing inflation.

[3] See Caballero and Engel (2007), Midrigan (2011) and Alvarez et al. (2016).

[4] Golosov and Lucas (2007) and Alvarez et al. (2018) show that the Golosov and Lucas (2007) model reproduces the relationship between inflation and the fraction of price changes in a cross-section of countries.

responds to monetary policy shocks of different sizes. Because the fraction of price changes increases considerably in the Golosov and Lucas (2007) model in response to large monetary shocks, the model predicts important non-linearities: the real effects from larger shocks are proportionally smaller than those from smaller shocks. However, because the model predicts little dispersion in the size of price changes, the real effects of monetary policy shocks are small even in response to small shocks. For example, the cumulative impulse response to a 1% monetary shock is 0.14 of that predicted by the Calvo model with the same average fraction of price changes, and that to a 5% monetary shock is 0.08 of that in the Calvo model. Thus, in the Golosov and Lucas (2007) money is approximately neutral regardless of the size of the shock.

In contrast, the other two variants of the model that better reproduce the dispersion in the size of price changes predict larger real effects. However, because in these economies the fraction of price changes barely responds, even to large shocks, the output effects are nearly linear in the size of the shock. Specifically, the Nakamura and Steinsson (2010) economy predicts a cumulative impulse response of output to a 1% monetary shock equal to 0.36 of that in the Calvo model and to a 5% shock equal to 0.30 of that in the Calvo model.[5] Similarly, the Calvo-plus economy augmented with uniform menu costs predicts a cumulative impulse response of output to a 1% monetary shock equal to 0.58 of that in the Calvo model and to a 5% shock equal to 0.56 of that in the Calvo model.

Our paper therefore corroborates the findings of Blanco et al. (2024a), where we made a similar point using U.K. micro-price data. Relative to that paper, here we focus on a simpler menu cost economy without strategic complementarities in price-setting and calibrate the model to match U.S. micro-price statistics. In addition, here we study the comovement between aggregate inflation and the fraction of price changes, whereas in Blanco et al. (2024a) we studied the comovement between sectoral inflation and the fraction of price changes.

We conclude that an important challenge for the menu cost literature is to develop models that can simultaneously reproduce the micro-price data, as well as the extent to which the fraction of price changes comoves with inflation in the time-series. In Blanco et al. (2024a) we proposed one potential solution to this challenge using a menu cost economy in which the losses from misallocation from price dispersion within multi-product firms are low. Because that model reproduces the dispersion in the size of price changes in the data, it predicts

[5] Auclert et al. (2023) also find little evidence of non-linearity in the Nakamura and Steinsson (2010) but, unlike us, they do not confront the model's predictions on the comovement between inflation and the fraction of price changes in the time series.

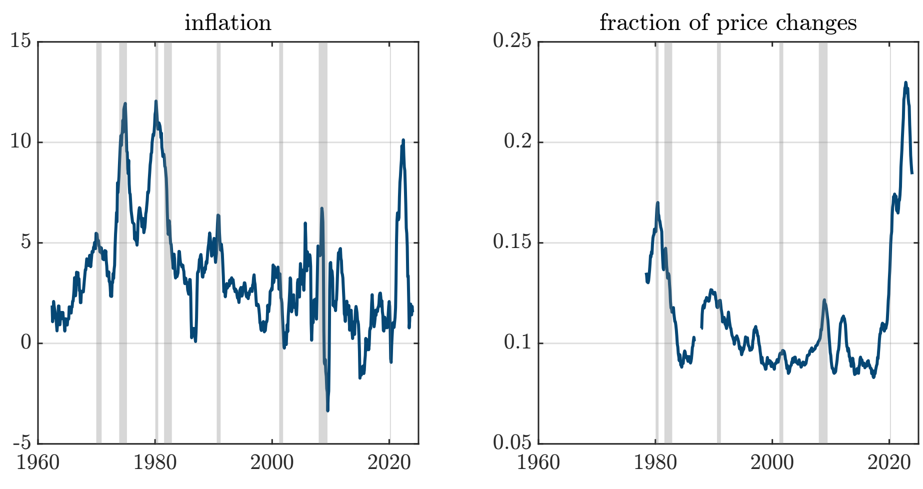
considerable real effects from monetary policy shocks. Moreover, because it predicts that the fraction of price changes increases considerably in times of high inflation, it also predicts highly non-linear output responses and therefore a time-varying slope of the Phillips curve.

2 Motivating Evidence

We briefly describe the evidence that motivates our paper. Figure 1 compares the time series of U.S. inflation with that of the fraction of price changes. We follow Nakamura et al. (2018) in measuring inflation using the growth of the Consumer Price Index excluding shelter. We use the data on the fraction of price changes computed by Nakamura et al. (2018) using price quotes collected by the Bureau of Labor Statistics. This series was recently updated by Montag and Villar (2023).[6] We report the year-on-year change in the price index and, for consistency, the average monthly fraction of price changes in the preceding year.

The figure shows that the fraction of price changes increases systematically with inflation, as pointed out by Nakamura et al. (2018) and Montag and Villar (2023).[7] For example, in the 1990s, when inflation was low, approximately 10% of prices changed in a given month. In contrast, in the 1980s, when inflation was high, approximately 17% of prices changed in a given month. More recently, the post-Covid increase in inflation was associated with an even higher fraction of price changes, approximately 23% per month.

Figure 1: Inflation and the Fraction of Price Changes



Notes: The shaded areas represent NBER recessions.

[6]We thank Hugh Montag and Daniel Villar for sharing the data with us.

[7]This is a robust feature of the data that has also been documented for other countries. See, for example, Gagnon (2009), Alvarez et al. (2018), Karadi and Reiff (2019), Blanco et al. (2024a) and Cavallo et al. (2024).

3 Model

We evaluate the ability of the menu cost model to reproduce the comovement between inflation and the fraction of price changes in the data. We consider an economy in which a continuum of monopolistically competitive firms are subject to idiosyncratic and aggregate shocks and face a menu cost of changing prices. We consider three specifications of the technology for changing prices, all widely used in the literature: a fixed menu cost as in Golosov and Lucas (2007), a Calvo-plus specification in which with some probability firms can change prices for free as in Nakamura and Steinsson (2010), and a uniform distribution of menu costs as in Blanco et al. (2024a).

3.1 Consumers

We follow Golosov and Lucas (2007) in assuming that preferences are logarithmic in consumption and linear in hours worked, so a representative consumer maximizes

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left(\log c_t - h_t\right),$$

subject to

$$P_t c_t + \frac{1}{1+i_t} B_{t+1} = W_t h_t + D_t + B_t,$$

where c_t is consumption, h_t is hours worked, P_t is the aggregate nominal price index, B_t is the amount of government bonds that pay the nominal interest rate i_t, and D_t denotes profits. The optimal labor supply choice implies that the nominal wage is equal to nominal spending, $W_t = P_t c_t$.

3.2 Monetary Policy

We assume that monetary policy targets nominal spending $M_t = P_t c_t$, which evolves according to

$$\log \frac{M_{t+1}}{M_t} = \mu_{t+1},$$

where $\mu_{t+1} \sim N\left(\mu, \sigma_m^2\right)$. Our assumptions on preferences imply that $W_t = M_t$.

3.3 Technology

There is a continuum of monopolistically competitive intermediate goods producers $i \in (0,1)$, each producing a differentiated variety using a linear technology

$$y_{it} = z_{it} l_{it},$$

where z_{it} is the quality of the variety produced by the firm and l_{it} is the amount of labor used in production.

A perfectly competitive final goods sector aggregates individual varieties using a CES aggregator with elasticity of substitution σ

$$y_t = \left(\int \left(\frac{y_{it}}{z_{it}} \right)^{\frac{\sigma-1}{\sigma}} \mathrm{d}i \right)^{\frac{\sigma}{\sigma-1}}. \tag{1}$$

The final good is used for consumption only, so $y_t = c_t$. In addition to affecting the firm's productivity, the quality z_{it} also affects demand. If prices were flexible, firms would respond to an increase in z_{it} by reducing prices one-for-one, leaving quality-adjusted prices and revenues unchanged. These shocks generate an idiosyncratic motive for firms to change their prices and are widely used in the menu cost literature due to their tractability. We assume that the z_{it} evolves according to

$$\log z_{it+1} = \log z_{it} + \varepsilon_{it+1},$$

where $\varepsilon_{it+1} \sim N\left(0, \sigma_z^2\right)$.

Letting P_{it} denote an individual firm's price, the demand function for the firm's output is given by

$$y_{it} = z_{it} \left(\frac{z_{it} P_{it}}{P_t} \right)^{-\sigma} y_t,$$

where

$$P_t = \left(\int \left(z_{it} P_{it} \right)^{1-\sigma} \mathrm{d}i \right)^{\frac{1}{1-\sigma}}$$

is the aggregate price index.

3.4 Price Adjustment Technology

As Caballero and Engel (2007), Midrigan (2011) and Alvarez et al. (2016) point out, the aggregate implications of menu cost models are critically shaped by the distribution of price changes. We therefore consider three different specifications for the price adjustment technology that increasingly allow the model to reproduce the distribution of price changes in the data. First, we follow Golosov and Lucas (2007) and assume that resetting prices requires paying a fixed menu cost $\bar{\xi}$. Second, we follow Nakamura and Steinsson (2010) and assume that with probability $1 - \lambda$ firms can change prices for free and with probability λ they have to pay a fixed menu cost $\bar{\xi}$. Third, we follow Blanco et al. (2024a) and assume that, in addition to the possibility of free price changes, the menu cost is idiosyncratic and drawn each period for a uniform distribution $U\left[0,\bar{\xi}\right]$. In all these cases, we assume that the price adjustment cost is denominated in units of labor. From now on, we refer to these economies as GL, NS and Uniform, respectively.

3.5 Firm Objective

The firm maximizes the expected present value of profits

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \frac{1}{P_t c_t} \left((1+\tau) P_{it} y_{it} - W_t \frac{y_{it}}{z_{it}} - \xi_{it} W_t \mathbb{I}_{it} \right),$$

where $\tau = 1/(\sigma - 1)$ is a subsidy that eliminates the markup distortion that would arise even with flexible prices, ξ_{it} is the possibly random cost of changing prices and $\mathbb{I}_{it}$ is an indicator equal to one if the firm adjusts its price and zero otherwise.

To characterize the firm's optimal choices, we first express its objective in terms of its *price gap* which we define as

$$x_{it} \equiv \frac{z_{it} P_{it}}{M_t}.$$

Similarly, we define the *aggregate price gap* as the CES weighted average of firm price gaps

$$X_t \equiv \left(\int x_{it}^{1-\sigma} \mathrm{d}i \right)^{\frac{1}{1-\sigma}} = \frac{P_t}{M_t}.$$

With this notation in place, we can write the firm's objective as a function of only its own and aggregate price gaps

$$\mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left(X_t^{\sigma-1} \left((1+\tau) x_{it}^{1-\sigma} - x_{it}^{-\sigma} \right) - \xi_{it} \mathbb{I}_{it} \right). \tag{2}$$

To characterize the price adjustment decision, let

$$s_{it} = \frac{z_{it} P_{it-1}}{M_t}$$

denote the firm's price gap in the absence of a price change, which is also the firm's idiosyncratic state. Each period the firm chooses whether to adjust its price or not. If the firm does not adjust, its price gap is $x_{it} = s_{it}$. If the firm adjusts its price, it resets the price gap to $x_{it} = x_t^*$, which is common to all firms that adjust. The firm recognizes that its state evolves over time according to

$$s_{it+1} = x_{it} \exp\left(\varepsilon_{it+1} - \mu_{t+1} \right).$$

Because this is a model with aggregate uncertainty in which heterogeneous firms follow non-linear decision rules, the entire distribution of price gaps, an infinite dimensional object, is necessary to characterize the equilibrium. To see why this is the case, let v_t^a denote the value of adjusting the price in period t and $v_t^n(s)$ the value of not adjusting the price for a firm with idiosyncratic state s. The firm adjusts its price with probability

$$h_t(s) = \begin{cases} \mathbb{I}\left(v_t^a - \bar{\xi} > v_t^n(s)\right), & \text{in GL} \\ 1 - \lambda + \lambda \mathbb{I}\left(v_t^a - \bar{\xi} > v_t^n(s)\right), & \text{in NS} \\ 1 - \lambda + \lambda \min\left\{\frac{v_t^a - v_t^n(s)}{\bar{\xi}}, 1\right\}, & \text{in Uniform.} \end{cases}$$

Letting $F_t(s)$ denote the distribution of firms, the equilibrium aggregate price gap satisfies

$$X_t = \left(\int \left[h_t\left(s\right)\left(x_t^*\right)^{1-\sigma} + \left(1 - h_t\left(s\right)\right) s^{1-\sigma}\right] \mathrm{d}F_t\left(s\right)\right)^{\frac{1}{1-\sigma}}.$$

Thus, computing the aggregate price gap needed to solve the firm's problem in equation (2) requires information about the entire distribution of firm price gaps.

We circumvent the curse of dimensionality by using the Krusell and Smith (1998) approach to characterize how X_t evolves over time as a function of a single moment of the distribution of $F_t\left(s\right)$, namely

$$S_t = \left(\int s^{1-\sigma} \mathrm{d}F_t\left(s\right)\right)^{\frac{1}{1-\sigma}}.$$

Specifically, we simulate a long history of monetary policy shocks, calculate the time series path for S_t and X_t implied by the model and use projection methods to update the perceived aggregate price gap function

$$X_t = \mathcal{X}\left(S_t\right),$$

where $\mathcal{X}\left(\cdot\right)$ is a linear combination of Chebyshev polynomials that allow us to capture potential non-linearities.

3.6 Parameterization

A period is a month and we set the discount factor β to an annualized value of 0.96 and the elasticity of substitution σ equal to 3. Following Nakamura and Steinsson (2010) and Auclert et al. (2022), in the NS and Uniform specifications of the adjustment cost we assume that the value of λ is such that 75% of price changes are free.

Table 1 reports the result of the parameterization. We calibrate the mean and volatility of the money growth rate μ and σ_m, the volatility of idiosyncratic quality shocks σ_z and the menu cost parameter $\bar{\xi}$ to target the average fraction of price changes, the median size of a price change, and the mean and standard deviation of inflation. These statistics are computed for the period 1979-2014 for which Nakamura et al. (2018) report a median size of regular price changes of 0.075. Using the data in Figure 1, we compute an average frequency of price changes of 0.105, an average annualized inflation rate of 3.4% and a standard deviation of annualized inflation of 2.6%. As Panel A of the table shows, all economies match these targets perfectly.

Panel B of the table reports the parameter values required to match the moments. Since, as pointed out by Bils and Klenow (2004) and Golosov and Lucas (2007), the average size of a price change is large relative to aggregate inflation, the model implies that idiosyncratic shocks are approximately 3 – 4 times more volatile than aggregate shocks. The menu cost parameter is relatively small in the GL economy and increases as we introduce randomness in the price adjustment costs.

Table 1: Parameterization

A. Moments

	Data	GL	NS	Uniform
I. Targeted				
fraction Δp	0.105	0.105	0.105	0.105
median $\lvert\Delta p\rvert$	0.075	0.075	0.075	0.075
mean inflation	0.034	0.034	0.034	0.034
std dev. inflation	0.026	0.026	0.026	0.026
II. Not targeted				
kurtosis Δp		1.466	2.402	3.346
10^{th} pctile $\lvert\Delta p\rvert$		0.065	0.012	0.013
25^{th} pctile $\lvert\Delta p\rvert$		0.068	0.032	0.034
75^{th} pctile $\lvert\Delta p\rvert$		0.084	0.174	0.136
90^{th} pctile $\lvert\Delta p\rvert$		0.095	0.196	0.202

B. Calibrated Parameter Values

		GL	NS	Uniform
μ	mean money growth rate	0.034	0.034	0.034
σ_m	s.d. monetary shocks	0.008	0.009	0.010
σ_z	s.d. idios. shocks	0.024	0.037	0.037
$\bar{\xi}$	menu cost	0.015	0.246	2.818

Note: The money growth rate is annualized. The menu cost parameter $\bar{\xi}$ is expressed relative to total revenue in a given period.

3.7 The Distribution of Price Changes

We next investigate the models' implications for broader moments of the distribution of price changes. Figure 2 plots the distribution of price changes implied by each of the three economies. As is well understood, the GL model produces a bi-modal distribution of price changes with neither very small nor very large price changes. This is also reflected in the untargeted statistics reported in Panel A of Table 1 which shows that in this economy the distribution of the size (absolute value) of price changes is very compressed, ranging from a 10^{th} percentile of 6.5% to a 90^{th} percentile of 9.5%. An alternative way of summarizing the dispersion in the size of price changes is the kurtosis of price changes, which is equal to 1.5.

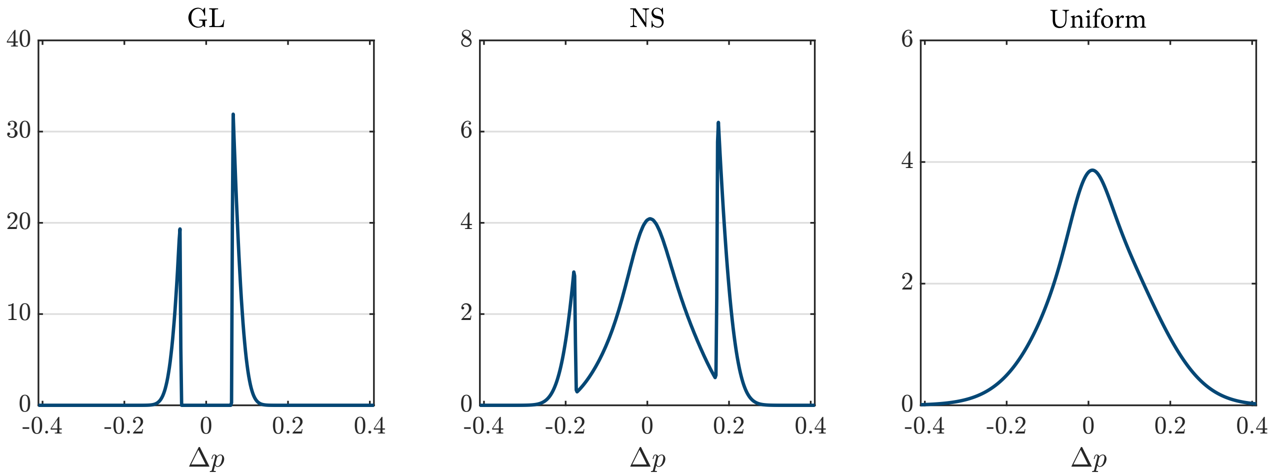
Both the NS and the Uniform economies generate more dispersion in the size of price changes and a higher kurtosis, but the distribution in the NS economy has jumps, reflecting that the adjustment hazard is a step function, as in the GL economy. In contrast, the smoother adjustment hazard induced by the uniform menu cost distribution in the Uniform economy produces a smooth distribution of price changes. Though we do not have direct access to the BLS micro-price data, the evidence on the distribution of price changes in existing studies is most consistent with the predictions of the Uniform economy. For example, using the same BLS data Klenow and Kryvtsov (2008) report a unimodal distribution of regular price changes, with 25% of price changes below 2.5%. This pattern has also been documented for other countries (see Alvarez et al., 2016 for France, Blanco et al., 2024a for the UK and Gautier et al., forthcoming for 11 European countries), and for specific sectors (see Midrigan, 2011).[8]

4 Nonlinear Dynamics?

Because the fraction of price changes varies over time in menu cost economies, these models predict potentially nonlinear responses to aggregate shocks. We next investigate the extent of these nonlinearities.

[8]A recent paper by Alvarez et al. (2021) uses data from a grocery store and documents that, after controlling for measurement error, the distribution of price changes is bimodal, featuring a considerable dip near zero. This is true for all the countries in their sample, except for the U.S., for which the distribution is unimodal. Such a bimodal distribution can be rationalized with a model with random menu costs but no free price changes. We study this model in the Appendix and show that it has similar aggregate implications as the NS and Uniform economies.

Figure 2: Distribution of Price Changes



4.1 The Fraction of Price Changes

We start by gauging the ability of each of the three economies above to reproduce the fluctuations in the fraction of price changes observed in the U.S. time series. To that end, we use the nonlinear solution of each model to back out the history of monetary shocks μ_t that allows each model to exactly reproduce the time series of inflation in Figure 1.[9]

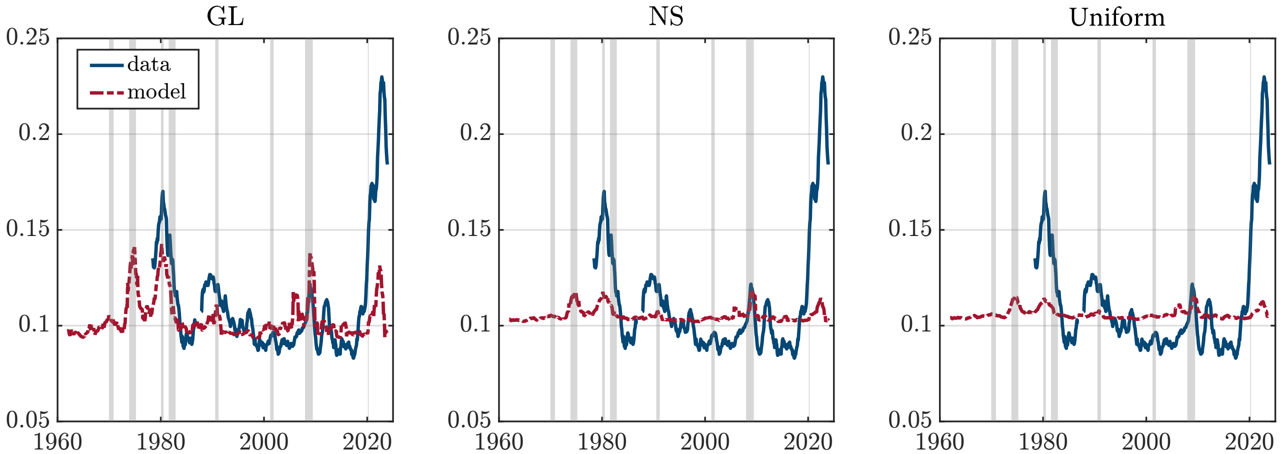
Figure 3 contrasts the series for the fraction of price changes in the model and in the data. As the left panel shows, the GL model predicts significant movement in the fraction of price changes, but not as much as in the data. For example, in 1980 the fraction of price changes in the data increased to 0.17 and in the model it increased to 0.14.[10] Similarly, during the post-COVID inflation episode, the fraction of price changes increased to 0.23 in the data and to only 0.13 in the model. Overall, a regression of the fraction of price changes on the absolute value of inflation for the 1979-2014 sample period in Nakamura et al. (2018) implies a slope coefficient of 0.63% in the data and 0.32% in the model, so the GL model generates half of the comovement between inflation and the fraction of price changes in the data.

The middle and right panels of Figure 3 show that the NS and Uniform economies imply much smaller fluctuations in the fraction of price changes: the fraction of price changes never exceeds 12% in either economy. The slope coefficient from regressing the fraction of price changes on the absolute value of inflation is 0.1% in the NS economy and 0.07% in the

[9]For simplicity, we assumed that the monetary policy shocks are the only source of aggregate uncertainty. Our results would be identical if we also allowed for aggregate productivity shocks, provided these also follow a random walk, since what matters is the process for nominal marginal costs.

[10]These numbers are similar to those obtained by Nakamura et al. (2018).

Figure 3: Fraction of Price Changes



Notes: The shaded areas represent NBER recessions.

Uniform economy, much smaller than the 0.63% in the data.

In Blanco et al. (2024a) we show that the response of the fraction of price changes to aggregate shocks depends on the size of the shock relative to the dispersion in price changes. Models that account for the dispersion of price changes predict that idiosyncratic, rather than aggregate shocks determine most of the adjustment decisions, rendering the fraction of price changes nearly time-invariant when we confront the model with aggregate shocks of similar magnitude to those needed to account for time-series fluctuations in the U.S. inflation rate. For larger shocks, even these versions of the menu cost model generate fluctuations in the fraction of price changes. However, such large shocks would generate a lot more variation in inflation than observed in the data.

Thus, we conclude that none of the menu cost models studied in this paper can simultaneously reproduce the micro-price data and the comovement between inflation and the fraction of price changes. On the one hand, versions of the model that generate substantial fluctuations in the fraction of price changes, such as GL, predict too little dispersion in the size of price changes. On the other hand, versions of the model that reproduce the large dispersion in the size of price changes, such as NS and Uniform, predict a nearly constant fraction of price changes.

As we illustrate below, the dispersion in the size of price changes critically determines the real effects of monetary shocks, and fluctuations in the fraction of price changes are crucial in determining how nonlinear these real effects are. Therefore, an important challenge for menu cost models is to simultaneously account for both of these features of the data. In Blanco et al. (2024a) we provide one potential solution to this challenge using a multi-product menu cost model with a low degree of misallocation inside the firm.

4.2 Real Effects of Monetary Shocks

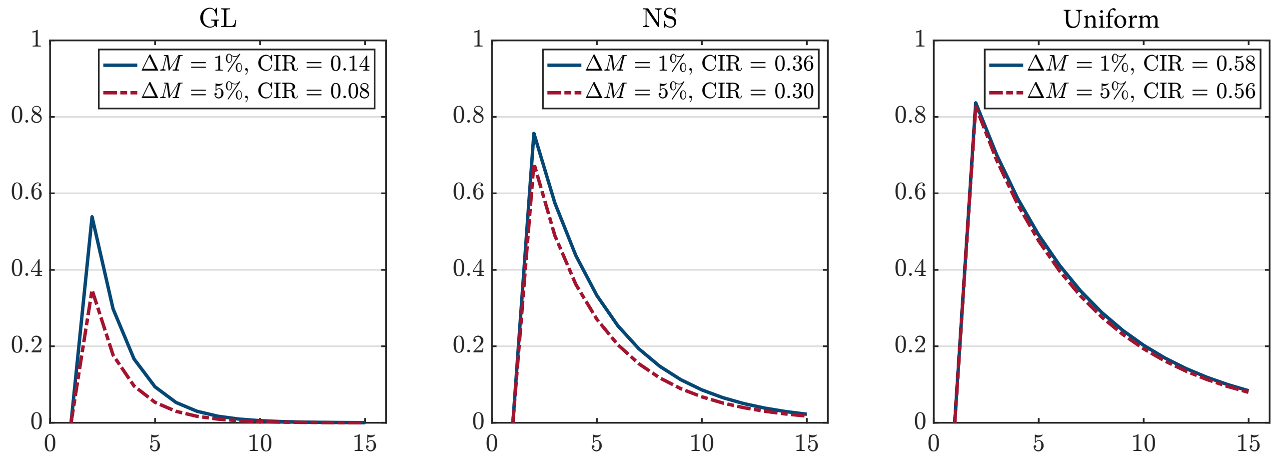
We next investigate the extent of nonlinearities that the three models imply by studying how each of these economies responds to monetary shocks. We consider two shocks: a relatively small increase in nominal spending of 1% and a larger one of 5% and trace out the implied impulse response of real output. We compute this impulse response using the nonlinear solution $X_t = \mathcal{X}\left(\frac{X_{t-1}}{\exp(\mu_t)}\right)$, initializing the economy at its stochastic steady state.

Figure 4 plots the output impulse responses which, for comparability, are scaled by the size of the monetary shock. The left panel of the figure reproduces the Golosov and Lucas (2007) finding that the output effects are small and short-lived. A summary statistic commonly used in the literature is the cumulative impulse response (CIR), which in response to a 1% shock is 0.14 as large as that predicted by a Calvo model in which the frequency of price changes is set equal to the average frequency in the menu cost model. Because the fraction of price changes responds to the shock, in the GL economy the output response to a large shock is even smaller and has a CIR that is 0.08 of that in the Calvo model. As argued in Midrigan (2011), these responses are small and short-lived precisely because the model fails to match the dispersion in the size of price changes in the data.

The middle and right panels of the figure show that the NS and Uniform economies that better reproduce the dispersion in the size of price changes predict larger real effects of monetary shocks. However, because in these economies the fraction of price changes is relatively irresponsive to aggregate shocks, the output effects are nearly linear. In the NS economy, the CIR after a 1% monetary shock is 0.36 of that in Calvo, not much larger than the 0.3 relative CIR after a 5% shock. In the Uniform economy, which predicts the greatest degree of monetary non-neutrality, the output responses to small and large shocks are nearly identical.

We therefore conclude that although the GL economy predicts significant nonlinearities, the real effects of monetary policy shocks are small because the model fails to match the distribution of price changes in the data. Versions of the menu cost model that are consistent with the micro-price data predict larger real effects and a modest degree of nonlinearity.

Figure 4: Output Impulse Response to Monetary Shocks



Notes: The output responses are scaled by the size of the monetary shock. The CIR is computed relative to that in a linearized Calvo model in which the frequency of price changes is the same as the average frequency of price changes in the menu cost model.

5 Conclusions

It is widely believed that menu cost models predict important nonlinearities in response to aggregate shocks because firms choose endogenously the timing of their price changes and are more likely to respond to larger shocks. Whether this is the case or not has important implications for the slope of the Phillips curve and the tradeoff faced by monetary policy. We evaluate the importance of nonlinearities by confronting several variants of the menu cost model used in the literature with the history of shocks necessary to exactly reproduce the time series of U.S. inflation.

The canonical Golosov and Lucas (2007) model indeed predicts that the fraction of price changes increases with inflation, albeit not as much as in the U.S. data. Therefore, the model features a substantial degree of nonlinearity. However, because the model fails to match the dispersion in the size of price changes, it predicts that money is nearly neutral. In contrast, variants of the menu cost model that can reproduce the large dispersion in the size of price changes predict larger real effects of monetary policy shocks, but imply a nearly constant fraction of price changes and therefore linear output dynamics.

We therefore conclude that an important challenge for the menu cost literature is to develop models that can simultaneously reproduce the micro-price data, as well as the extent to which the fraction of price changes comoves with inflation.

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Appendix

For Online Publication

A An Economy Without Free Price Changes

In this section we consider a menu cost economy in which there are no free price changes and menu costs are drawn each period from a uniform distribution $\xi_{it} \sim U\left[0, \bar{\xi}\right]$. We calibrate the parameters of this model using the same strategy as in the main text. Table A.1 reports the results of the parameterization. Once again the model matches the targeted moments perfectly.

Table A.1: Parameterization of Economy Without Free Price Changes

A. Moments

	Data	Model
I. Targeted		
fraction Δp	0.105	0.105
median $\lvert\Delta p\rvert$	0.075	0.075
mean inflation	0.034	0.034
std dev. inflation	0.026	0.026
II. Not targeted		
kurtosis Δp		2.001
10^{th} pctile $\lvert\Delta p\rvert$		0.036
25^{th} pctile $\lvert\Delta p\rvert$		0.052
75^{th} pctile $\lvert\Delta p\rvert$		0.102
90^{th} pctile $\lvert\Delta p\rvert$		0.126

B. Calibrated Parameter Values

		Model
μ	mean money growth rate	0.034
σ_m	s.d. monetary shocks	0.008
σ_z	s.d. idios. shocks	0.026
$\bar{\xi}$	menu cost	0.128

Note: The money growth rate is annualized. The menu cost parameter $\bar{\xi}$ is expressed relative to total revenue in a given period.

Figure A.1 plots the distribution of price changes implied by this model, which is bimodal and features a dip near zero, consistent with the empirical evidence in Alvarez et al. (2021). The model therefore predicts fewer small price changes than the economies with free price changes. As Panel A of Table A.1 shows, the 10^{th} percentile of the distribution of the size of price changes is 3.6%. The model also produces fewer large price changes: the 90^{th} percentile is 12.6%.

Figure A.1: Distribution of Price Changes in Model without Free Price Changes

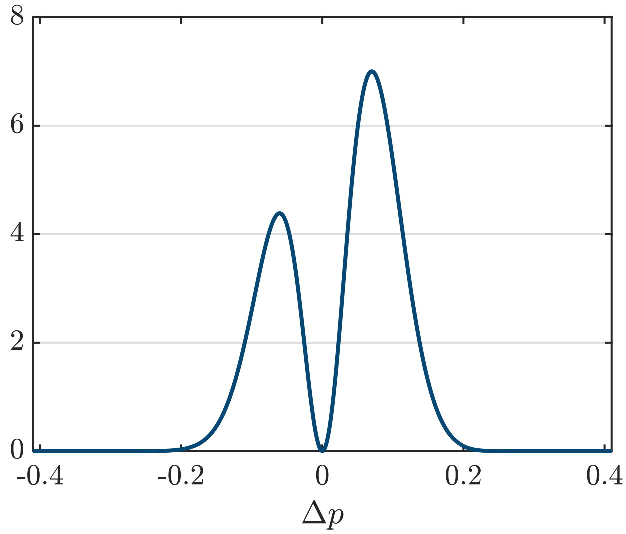
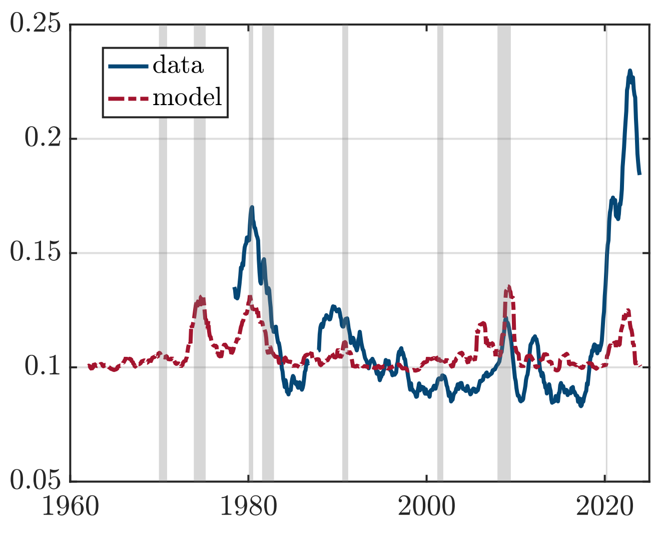


Figure A.2 shows that the fraction of price changes predicted by the model, when confronted with the history of shocks necessary to exactly reproduce the U.S. inflation time series, increases in times of high inflation, but not as much as in the data. The slope coefficient from regressing the fraction of price changes on the absolute value of inflation is equal to 0.2%, one-third of the 0.6% in the data.

Figure A.2: Fraction of Price Changes in Model without Free Price Changes



Notes: The shaded areas represent NBER recessions.

Because the fraction of price changes does not fluctuate a lot in response to aggregate shocks, the model predicts little nonlinearity in response to small and large monetary shocks, as shown in Figure A.3. Specifically, the output CIR after a 1% shock is 0.28 of that in the linearized Calvo model, similar to that after a 5% shock, which is 0.23 of that in the Calvo model.

Figure A.3: Output Impulse Response in Model without Free Price Changes