Non-linearities, state-dependent prices and the transmission mechanism of monetary policy

Short title: Non-linearities and state-dependent prices

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Abstract

A sticky price theory of the transmission mechanism of monetary policy shocks based on state-dependent pricing yields two testable implications, that do not hold in time-dependent models. First, large monetary policy shocks should yield proportionally larger initial responses of the price level. Second, in a high trend inflation regime, the response of the price level to monetary policy shocks should be larger and real effects smaller. Our analysis provides evidence supporting these non-linear effects in the response of the price level in aggregate US data, indicating state-dependent pricing as an important feature of the transmission mechanism of monetary policy.

Keywords: Sticky prices, state-dependent pricing, monetary policy

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1 Introduction

The New Keynesian (NK) paradigm is one of the main frameworks for the analysis of business cycle fluctuations and the effects of monetary and fiscal policy, both among academic researchers and policy institutions. While there can be alternative mechanisms for the transmission of monetary policy shocks, price stickiness remains the major reason why monetary policy has effects on real variables in virtually all NK models. This pivotal assumption has been substantiated by numerous empirical studies that show that individual prices do indeed change infrequently (e.g. Bils and Klenow, 2004; Klenow and Kryvtsov, 2008; Nakamura and Steinsson, 2008; Nakamura, 2008; Eichenbaum *et al.*, 2011; Nakamura and Zerom, 2010). Moreover, this literature has demonstrated that microfounded state-dependent models are able to replicate the empirical distribution of price changes. Unlike time-dependent pricing models, such as Calvo (1983) and Rotemberg (1982), microfounded state-dependent models of nominal rigidities assume that the individual firm can change its price, subject to an adjustment cost. Hence, the frequency of price changes, the number of firms that decide to change their prices following a monetary policy shock and the size of these changes are endogenous (the so-called "selection effect", see Golosov and Lucas, 2007). In particular, these models predict that individual prices (i) are more flexible in response to large shocks (e.g., Karadi and Reiff, 2019); (ii) change more frequently when inflation is higher (e.g., Álvarez *et al.*, 2019).

Some papers find evidence of state-dependent pricing behaviour by investigating the response of prices at the micro-economic level to a particular event. Hobijn *et al.* (2006) document a dramatic increase in restaurant prices in the euro area after the introduction of the euro and explain it through the lens of a menu cost model. Karadi and Reiff (2019) document that micro-level price responses to three large value-added tax (VAT) changes, which occurred in Hungary in 2004 and 2006, were flexible and asymmetrical with respect to positive versus negative tax changes.[1] Bonadio *et al.* (2019) exploit the Swiss National Bank's (SNB) decision to discontinue the minimum exchange rate policy of one euro against 1.2 Swiss francs on January 15, 2015, to analyse the pass-through of the Swiss exchange rate shock into import product prices from the Euro area. The empirical literature on exchange rate pass-through is very large. It generally concerns measuring the extent of the exchange rate pass-through, and, while somewhat related, it is not concerned explicitly with state-dependent pricing models. Two notable exceptions are Álvarez *et al.* (2017) and Álvarez and Neumeyer (2020) that compare the observed price dynamics at the micro level after a large devaluations to the predictions of an adjustment cost model of price setting. Álvarez *et al.* (2017) use monthly data on Consumer Price Index (CPI) inflation and the nominal exchange rate from a large number of countries in a panel data analysis. Álvarez and Neumeyer (2020) analyse several episodes involving large changes in the nominal price of inputs in Argentina over 2012–2018 by using micro-level price data for the city of Buenos Aires. Álvarez *et al.* (2017) is the closest paper to ours because they use aggregate data and they link explicitly the empirical analysis to the prediction of state-dependent pricing model that the inflation response should depend on the size of the shock.

Yet, despite the large literature on sticky prices, to the best of our knowledge, there is surprisingly little direct evidence that *aggregate* prices behave as state-dependent models suggest and thus that (i) and (ii) affect the transmission mechanism of monetary policy in aggregate US data.[2] We aim to fill this gap. If staggered prices are of such paramount importance to the transmission mechanism of monetary policy, and if there is a significant fraction of state-dependent prices, then the two very general predictions

[1] Álvarez *et al.* (2006) and Gagnon *et al.* (2012) are earlier papers that find significant increases in the frequency of price changes in the months with VAT tax changes, consistent with state-dependent pricing models.

[2] As said, the closest reference is Álvarez *et al.* (2017) who however investigates only prediction (ii) above, use a panel of country and look at exchange rate pass-through interpreting exchange rate movements as cost shocks. Veirman (2009) investigates empirically whether the flattening of the Japanese Phillips Curve depends on lower trend inflation, as state dependent models would imply.

(i) and (ii) should also emerge in aggregate data. First, *large absolute value* monetary policy shocks should lead more firms to adjust their prices, and hence yield a proportionally larger response of inflation, whereas the real effects should be subdued. Second, the frequency of price changes should be an increasing function of underlying levels of inflation, that is, prices should be more flexible in a *high trend inflation* regime than otherwise. Hence, the higher trend inflation, the larger the response of inflation and the smaller should be the real effects of monetary policy shocks.

These theoretical results are quite intuitive and general since they derive from a broad variety of state-dependent sticky price models in the literature. In particular, Álvarez *et al.* (2017) show that the impulse response after a monetary shock is size-independent in time-dependent models, whilst it is not in state-dependent models. In this sense, our contribution is really a minimal, first-pass test for state-dependent price theories: if state-dependency is pivotal, aggregate data should exhibit these two features. We take these two theoretical predictions to US data between 1969 and 2007, applying smooth local projections (Jordà, 2005; Barnichon and Brownlees, 2019) and the smooth transition function methodology of Granger and Teräsvirta (1993) together with US monetary policy shocks identified with the narrative method of Romer and Romer (2004). The empirical methodology of this paper is most closely related to the work of Tenreyro and Thwaites (2016), focusing on the differential effects of monetary policy in recessions and expansions. As such, our paper contributes also to the literature about the state dependent effects of monetary shocks.

Our analysis provides new and statistically significant evidence in favour of state-dependent pricing models in aggregate US data. First, large absolute value shocks have disproportionately larger effects on prices on impact, but are less persistent and have weaker real effects, matching the first theoretical prediction. Second, the impulse response functions (IRFs) in the high and low trend inflation regimes are significantly different for prices and inflation and also in line with the second theoretical prediction of higher price flexibility in the high trend inflation regime. However, for the second prediction, we do not find statistically significant evidence of muted real effects. Importantly, the non-linearity of the impulse response functions is not due to a different feedback of monetary policy to inflation in response to small vs. large shocks or in periods of high vs. low trend inflation. In the Appendix we conduct a comprehensive sensitivity analysis to establish the robustness of these results, regarding different empirical specifications, sub-samples, controls, and measures of monetary policy surprises.

Our results are the first ones (to the best of our knowledge) that point towards a significant presence of state-dependent pricing in the US economy from an aggregate perspective. Our macro evidence is therefore a useful complement to existing micro evidence.

2 Theory and Testable Implications

Contrary to early standard time-dependent pricing models (e.g., Fischer, 1977; Taylor, 1979, 1980; Calvo, 1983), microfounded state-dependent pricing models feature endogenous price adjustments triggered by changes in the economic environment. The firms that decide to change their prices by paying the adjustment costs after a monetary shock are the ones further off from their optimal price ("selection effect"). Hence, since these are not random firms and the sizes of the price changes are relatively large, early studies (Caplin and Spulber, 1987; Golosov and Lucas, 2007) find that the aggregate price level can mimic a flexible price environment. Nonetheless, later studies investigate the robustness of this result to various extensions and show that state-dependent models yield a large degree of aggregate price stickiness and are important for the transmission mechanism of monetary policy.[3]

[3]These extensions are informational costs (Gorodnichenko, 2008; Bonomo *et al.*, 2013; Álvarez *et al.*, 2011), multi-product firms (Álvarez and Lippi, 2014; Midrigan, 2011; Bhattarai and Schoenle, 2014), multiple sectors and intermediate inputs (Nakamura and Steinsson, 2010), or errors in the timing and the precision of the adjustment (Costain *et al.*, 2019). The same result holds in a variety of state-

First Testable Implication: The impulse response functions of inflation and output to a monetary policy shock should depend on the size of the shock. If there is a non-negligible fraction of state-dependent price-setters, the impulse response should be a non-linear function of the size of the shock. The larger the shock, the larger the number of firms that decide to pay the adjustment costs and change their prices immediately, such that the reaction of the aggregate price level at short horizon is increasing in the size of the monetary policy shock (see Álvarez and Lippi, 2014). Moreover, the effect on the price level should be less persistent, because the larger the shock, the higher the number of firms that adjust immediately and hence the lower the number of firms that eventually would change their price as the shock tapers off. The real effects of monetary policy shocks, instead, are hump-shaped with respect to the size of the shock, because of two counteracting effects. First, larger shocks, *ceteris paribus*, give rise to stronger real effects, just like in a time-dependent model. Second, larger shocks also increase the number of adjusting firms, strengthening the reaction of the aggregate price level and thus reducing the real effects. For a small shock, the first effect prevails, so that both the impact and the cumulative effect on output is increasing in the size of the shock. For large shocks the opposite occurs. It follows that sufficiently big shocks should have lower real effects than smaller shocks.[4]

Second Testable Implication: The impulse response functions of inflation and output to a monetary policy shock should depend on the average level of inflation. As outlined in Dotsey *et al.* (1999) or Costain and Nakov (2011), average inflation affects the frequency of price adjustments in state-dependent models, because it erodes a firm's relative price so that firms adjust prices more frequently. Indeed, the

dependent models that are carefully calibrated to match the main features of retail price microdata (Álvarez *et al.*, 2016; Eichenbaum *et al.*, 2011; Costain and Nakov, 2011, 2019), and, more recently, of both the size and frequency of price *and wage* changes (Costain *et al.*, 2019).

[4] Álvarez and Lippi (2014) show that the monetary shock that maximises the cumulated effect on output (i.e., the area under the impulse response function) is about one-half of the standard deviation of price changes. Costain *et al.* (2019) show that a similar effect occurs in a model with state-dependent prices and wages.

empirical analysis in Álvarez *et al.* (2019) provides solid evidence of how the frequency of price changes varies with inflation.[5] This evidence implies different impulse responses to a monetary policy shock in high trend inflation regimes compared to low trend inflation regimes. In particular, we should observe a quicker and less persistent reaction of prices in high inflation regimes. Further, Álvarez *et al.* (2016) show that in a large class of sticky-price models, the total cumulative output effect of a small unexpected monetary shock is inversely related to the average number of price changes per year.[6] This theoretical prediction provides the second testable implication.

3 Empirical Methodology

In this Section we describe the data and the empirical methodology used to estimate smooth impulse responses and conduct inference. We test the two predictions by analysing the presence of non-linearities in the impulse response functions to a monetary policy shock for large and small shocks, and during high and low trend inflation. Our empirical methodology follows a growing body of literature employing local projections (Jordà, 2005) to account for the response to non-linear terms and state-dependency in empirical impulse responses (e.g. Auerbach and Gorodnichenko, 2012a,b; Caggiano *et al.*, 2014, 2015; Tenreyro and Thwaites, 2016; Ramey and Zubairy, 2018; Furceri *et al.*, 2016).

Data The monthly sample for our response variables runs from 1969m1 up to 2007m12, excluding the most recent financial crisis, where monetary policy has been very different and the zero-lower bound on nominal interest rates has been binding.

[5] Figure 5 and 6 therein show that the frequency of price changes do not react much for levels of annual inflation up to 5%, then it starts accelerating and finally it increases linearly for values of annual inflation above 14% with an elasticity of about two-thirds. This is in line with Sheshinski and Weiss (1977)'s adjustment cost model with no idiosyncratic shocks.

[6] The same holds in a model in which nominal rigidities in both wages and prices are state-dependent (see Figure 11 in Costain *et al.*, 2019), and in a model that allows for temporary price changes, because firms can set a price plan, rather than a fixed price as in the standard adjustment cost model (see Figure 7 in Álvarez and Lippi, 2020).

We analyse three response variables: output, inflation and the nominal interest rate. The series for output is the industrial production index, for inflation we use personal consumption expenditure (PCE) inflation, and for the nominal interest rate we use the effective federal funds rate, all from the Federal Reserve Bank of St. Louis Database (FRED).[7]

The main shock variable used in this analysis is based on the narrative analysis of Romer and Romer (2004). They identify monetary policy surprises by using a narrative approach to infer the intended federal funds rate at every Federal Open Market Committee (FOMC) meeting from 1969 onwards. By regressing changes of this intended rate on Greenbook forecasts they derive a measure of monetary policy surprises that is arguably exogenous to the Fed's information set about the future state of the economy. We utilise this methodology and the extended shock sample until December 2003 by Tenreyro and Thwaites (2016).

There are a number of other ways of identifying monetary policy surprises, for example High Frequency Identification (HFI) (Kuttner, 2001; Gürkaynak *et al.*, 2005; Barakchian and Crowe, 2013; Gertler and Karadi, 2015) or identification restrictions in a Vector Autoregression (VAR) (Christiano *et al.*, 1999; Bernanke and Mihov, 1998; Kim and Roubini, 2000; Uhlig, 2005; Bernanke *et al.*, 2005). We conduct robustness tests with respect to the latter and find largely similar conclusions.[8]

Smooth Local Projections We estimate local projection coefficients using the recently developed methodology by Barnichon and Brownlees (2019) to improve accuracy

[7] The variables are all in natural log levels and then multiplied by 100 except for the federal funds rate which remains in percentage points. This standard transformation enables the interpretation of the strength of the coefficients as approximate percentage points.

[8] We do not, however, include HFI shocks for two main reasons. First, the available shock series are too short for our analysis. Even the backward extension up to 1979 by Gertler and Karadi (2015) omits a significant proportion of the Great Inflation period; a major source of variation in our smooth transition function. Second, Ramey (2016) argues that these shocks may not be robust to samples where anticipation effects are important. Moreover, (Ramey, 2016, p. 109) shows that the impulse response function to HFI shocks can look very different depending on whether one uses a VAR or local projections. For example, a contractionary Gertler and Karadi (2015) shock in a local projection *increases* output and leaves the price level largely unchanged whereas it produces the literature standard effect in their proxy SVAR.

and inference over the standard least squares approach.[9] With this technique, the impulse response coefficients, i.e. all shock-dependent coefficients, are modelled as linear combinations of B-spline basis functions. One can then estimate the coefficients of these linear combinations using generalised ridge estimation, with a penalty parameter that selects the degree of shrinkage. When the shrinkage parameter is close to zero, the estimation yields the standard least squares estimates. Conversely, if the parameter is high, the impulse response is converging to a smooth limit polynomial distributed lag model. We follow Barnichon and Brownlees (2019) and select the shrinkage parameter using k-fold cross validation (Racine, 1997).

We also follow Barnichon and Brownlees (2019) on conducting inference. In particular, in order to take into account potential autocorrelation and heteroskedasticity we estimate the variance of the coefficients using a modified Newey and West (1987) estimator, corrected for the penalty parameter. We use the resulting variance matrix to construct confidence intervals and t-statistics in the ordinary way. We provide further details on this and the estimation technique as a whole in the Appendix.

4 Results

This Section presents the main results regarding the two testable implications above.

4.1 Implication 1: Size-dependent effects of monetary policy shocks

Non-linear local projections. In order to test the size-dependence of impulse responses we consider the following non-linear local projection:

$$y_{t+h} = \alpha_h + \tau_h t + \beta_h e_t + \zeta_h(e_t \cdot |e_t|) + \sum_{k=1}^{K} \gamma_{h,k} w_{t,k} + v_{t+h}, \tag{1}$$

[9] We provide the results with standard least squares coefficients in the Appendix.

which is estimated for $h = 0, 1, \ldots, H$. We set $H = 48$ which corresponds to an impulse response horizon of four years. y_{t+h} denotes the variable of interest, in our case either the industrial production index, PCE inflation or the federal funds rate. e_t are the narrative Romer and Romer (2004) shocks. $w_{t,k}$ denotes the kth control variable and v_{t+h} the estimation error, possibly heteroskedastic and serially correlated. The set of control variables includes up to two months of lags of industrial production, PCE inflation and the federal funds rate. Moreover, we follow Ramey (2016) and include contemporaneous values of the industrial production index and PCE inflation. This is equivalent to assuming recursiveness between the three different variables of interest since inflation and industrial production can contemporaneously affect the federal funds rate but not vice versa.[10]

While the coefficient β_h captures the linear component, the coefficient ζ_h on the absolute value interaction term accounts for non-linearities in the impulse response function due to the size of the shock. The interaction term, $(e_t \cdot |e_t|)$, magnifies the size of the shock, but it keeps the same sign of the shock. Hence, contrary to a simple quadratic term, it isolates the pure effect of a change in the size of the shock. If ζ_h has the same sign of β_h then the non-linear interaction term *amplifies* the linear effects of the impulse response. On the contrary, if ζ_h has the opposite sign to β_h, it *counteracts* the linear impulse response, possibly even tilting the overall effect from one sign to another for large enough shocks. Whenever $\zeta_h = 0$ the impulse response function is linear with respect to the shock size.

Therefore, ζ_h is the main coefficient of interest to test our first theoretical prediction. Let y_t signify prices, and assume a large monetary contraction, i.e., a positive value of e_t. Large monetary policy shocks should induce a more price-flexible impulse response function of the price level and inflation. If firms exhibit state-dependent pricing we should

[10] As Ramey (2016) points out, relaxing this assumption would otherwise lead to a number of puzzles. A contractionary monetary policy shock would actually be expansionary for about a year and produce a very pronounced price puzzle.

see a negative ζ_h at small horizons as we expect β_h to be close to zero or negative. This would mean that firms decrease prices quicker and so prices decline disproportionately at small horizons. Furthermore, we would then expect to see a positive ζ_h at larger horizons, weakening the price response, as more firms have already changed prices earlier. Consequently a combination of a negative ζ_h at small horizons, as more firms change prices right away, and a positive ζ_h at larger horizons, as persistence is lower due to earlier price changes, would speak in favour of state-dependent pricing as a valid aggregate propagation mechanism of monetary policy shocks.

Coefficient estimates. The resulting coefficients from estimating local projection (1) for PCE inflation, industrial production and the federal funds rate are reported in the panels of Figure 1. The solid black line plots the coefficients of the linear term, $\hat{\beta}_h$, and the dashed-dotted green line plots the ones of the non-linear absolute value interaction term, $\hat{\zeta}_h$, together with their 90% confidence interval bands.

The linear terms in the projection deliver a familiar picture in the top panel. After a positive (contractionary) monetary policy shock the linear coefficients yield an initially muted response of inflation followed by significant decline thereafter. There is an initial positive response which is however not significant on impact and marginally so for just few months in the initial year. Hence, the response displays a slight price puzzle, if we were to consider only the linear effect.

The non-linear effects implied by the absolute value interaction coefficients are supporting the theoretical predictions. Initially the $\hat{\zeta}_h$ coefficients are negative, counteracting the price puzzle as the shock size increases and indicating a quicker negative response of inflation on impact. Moreover, after about two years, the coefficients on the non-linear term turn positive (and increase), meaning that inflation responds less to a large shock at long horizons. Hence, the evidence suggests that inflation has a stronger reaction at short horizons and a weaker one at long horizons. This is consistent with theoretical models of state-dependent pricing as, for large shocks, more firms adjust immediately and hence less adjust later on. Most importantly, the confidence interval shows that the coefficients are statistically significant for most of the horizons.

The other two panels report the results for the linear coefficient in the industrial production and federal funds rate local projection. These results are in line with the theoretical implications from the literature. The linear output coefficient starts to fall (after a small, positive, but not significant, initial response), reaching its trough two and a half years after the shock, and then recovers. The linear federal funds rate coefficient exhibits a hump-shaped response and remains positive for more than two years after the shock before turning negative.

Further, the green dashed-dotted line in the middle panel shows that the output response supports the predictions of state-dependent pricing models. The coefficients on the non-linear term in (1) are negative at the beginning, counteracting the small positive output response, and then turn positive at longer horizons. As such, the $\hat{\zeta}_h$ estimate counteracts the linear response for both short and long horizons, flattening the overall response of output. While the top panel shows that large monetary policy shocks predict a higher degree of price flexibility, the middle panel shows that larger monetary policy shocks have weaker real effects. Again, the confidence band indicates that the non-linear interaction term is statistically significant.

It is crucial to note that these results are not due to a stronger response of monetary policy, because the coefficients of the federal funds rate on the non-linear term in the bottom panel are negative for most of the short to medium horizons, suggesting a proportionally weaker response of monetary policy to a larger shock.

In sum, prices exhibit a non-linear, size-dependent impulse response function, reacting strongly at short and weakly at long horizons for large shocks. Coherently, the output response seems to be smaller, and monetary policy feedback does not seem to drive the above results. Hence, we interpret these findings as evidence in favour of state-dependent prices as an important propagation mechanism of monetary policy, as larger shocks induce more firms to change prices early, thus reducing the real effects of a monetary shock.

Impulse Response Functions. In order to further assess and clarify the importance of the non-linear effect, Figure 2 compares IRFs to different shock sizes of all three headline variables over a four year horizon. It depicts the impulse responses for a 25 (dashed blue line), 100 (solid black line) and 200 (dashed-dotted green line) basis point shock, where each impulse is standardised by dividing it by the respective shock size. The standardised IRFs clearly visualise that inflation responds more strongly to a larger shock at short horizons, but then the reaction is less persistent, so that the response is weaker at long horizons, as theory would predict. First, the larger the shock, the quicker inflation decreases. Second, for a large enough shock, the initial price puzzle on impact tend to disappear.[11]

The response of output to small and large shocks also supports the theoretical prediction. The standardised IRFs show that the trough in output is smaller relative to the size of the shock, consistent with the behaviour of the responses of inflation. Recall that, in state-dependent models, there are two opposite effects on output. First, the larger the shock, *ceteris paribus*, the larger the response of output. This standard effect is the only one present also in time-dependent pricing models. Second, the larger the shock, the greater the number of firms that adjust the price, hence the larger the response of inflation and the smaller the one of output. This second effect is absent in time-dependent models. Therefore, the output response to a large shock is proportionally flatter in state-dependent models, because of this second effect that counteracts the first one. By showing the response relative to the size of the shock, the standardised

[11]Figure H5, discussed in Appendix C.3, depicts the unscaled (i.e., non standardised) impulse responses for a 25 basis point shock and a 200 basis point shock and their 90% confidence interval calculated with the Delta method, respectively (see Appendix for details). The response of inflation to a 200 basis point shock is firstly not significantly different from zero and then significantly negative, while it is positive for some months for a 25 basis point shock (even if only marginally significantly). Consequently, a sufficiently large shock counteracts the small linear coefficient and switches the sign of the overall impulse response of inflation, removing a potential price puzzle.

IRFs isolate the second effect, thus, revealing whether there is a significant effect coming from state-dependent pricing. Finally, the standardised IRFs also highlight the effect of state-dependent pricing both on the scale, i.e., decreasing real impact as the shock gets larger, and on the timing, i.e., arriving sooner with larger shocks, of the response of output and inflation to the policy shocks.

To strengthen our point, Table 1 shows the cumulative effect of a monetary policy shock on inflation (i.e. the PCE deflator) and on output for small and large shock, standardised by the size of the shock, and their significance levels, for different horizons: 1, 12, 24, 36 and 48 months.[12] Coherently with the theoretical predictions, prices move significantly downwards on impact for large shocks, while they do not for small shocks. Due to this lagged and inertial behaviour in case of small shocks, the cumulative response of the price level after four years is almost doubled, relative to the size of the shock, compared with the response to a large shock. The cumulative response of output reflects the one of the price level. The initial response of output is sharper for large shocks, but the cumulative drop in output after four years is about 80% proportionally larger for small shocks.

Finally, the last row in Figure 2 again illustrates that the empirical results are not due to the different behaviour of monetary policy after a large shock. The standardised IRFs of the Fed funds rate is milder for larger shocks, relative to shock size, and hence, if anything, it would play against our results.[13]

Figure 2 and Table 1 reinforce our previous empirical results on the significant size-dependent effects of monetary policy shocks. Large shocks induce firms to change prices early on and thus reduce the real effects of such a monetary shock, in accordance with our first theoretical prediction.

[12] We simply cumulate the point estimates from the standard least squares estimation and statistics using the Delta method.

[13] Moreover, the response of the federal funds rate is statistically different (in the first year) between the 25 verus 200 basis points shock (see Figure H5 in the Appendix).

4.2 Implication 2: Trend inflation-dependent effects of monetary policy shocks

Smooth transition local projections. We use smooth transition local projections to test whether the impulse responses after a monetary policy shock are different in high and low inflation regimes. Auerbach and Gorodnichenko (2012b) and Tenreyro and Thwaites (2016) popularised this method, and we follow their approach to a large extent. The impulse response of the variable of interest y_t at horizon h in state $s = HI, LO$[14] to a unitary structural shock e_t is the estimated coefficient β_h^s in:

$$\begin{aligned} y_{t+h} = \tau_h t + F(z_t)(\alpha_h^{HI} + \beta_h^{HI} e_t + \sum_{k=1}^{K} \gamma_{h,k}^{HI} w_{t,k}) \\ +(1 - F(z_t))(\alpha_h^{LO} + \beta_h^{LO} e_t + \sum_{k=1}^{K} \gamma_{h,k}^{LO} w_{t,k}) + u_{t+h}, \end{aligned} \quad (2)$$

for $h = 0, 1, ..., H$. Again, our set of controls $\{w_{t,k}\}_{k=1}^{K}$ includes the contemporaneous values of industrial production and PCE inflation and up to two month lags of industrial production, PCE inflation and the effective federal funds rate. $F(z_t)$ is a smooth transition function which indicates the state of the economy (Granger and Teräsvirta, 1993). We use a logistic function with the following form:

$$F(z_t) = \frac{\exp{(\gamma \frac{(z_t - c)}{\sigma_z})}}{1 + \exp{(\gamma \frac{(z_t - c)}{\sigma_z})}} \in [0, 1]. \quad (3)$$

If state-dependent prices are an important aggregate propagation mechanism we would expect β_h^{HI} to be statistically significantly more negative than β_h^{LO} for the response of the price level and inflation, especially at short horizons. Prices should be more flexible and so react both more quickly and strongly to monetary policy shocks in a high inflation regime.

[14] HI stands for high inflation, LO for low inflation.

In the main specification of our smooth transition function local projection, i.e., equation (2), the state variable z_t represents smoothed personal consumption expenditure (PCE) inflation, so we take a 24 month centered moving average (MA) to capture trend inflation.[15] We set $\gamma = 5$ as this gives an intermediate degree of regime switching intensity. This is relatively standard in the literature and also fits our inflation data well. Finally, c corresponds to the 75th percentile of the historical trend inflation distribution. This is equivalent to assuming that about 70% of the time trend inflation is classified as negligible (i.e. $F(z_t) \in [0, 0.1]$) and 30% of the time there is some trend inflation (i.e. $F(z_t) \in (0.1, 1]$). Figure 3 displays the resulting smooth transition function, $F(z_t)$. The solid black line, measured on the left vertical axis, shows PCE inflation on an annual basis. The green dashed line, measured on the right vertical axis, depicts the smooth transition function based on our MA-filtered measure of PCE inflation. The period of the Great Inflation from around 1974 to 1983 is characterised by two pronounced spikes of inflation of up to 11%. The smooth transition function reaches 1 around these two peaks and stays above 0.4 for the entire period of the Great Inflation, classifying the latter period mostly as a high inflation regime. We take this to be a reasonable approximation for periods of high and low trend inflation in the United States.[16]

Coefficient estimates. The panels of Figure 4 display the coefficient from estimating equation (2) for each response variable (by row). The first column shows the point estimates of the $\hat{\beta}^i_h$ coefficients for the linear (solid black line), the high inflation (dash-dotted green line) and low inflation (dashed blue line) regimes. Column two and three depict the impulse responses conditional on the high inflation and low inflation regime respectively, with their 90% confidence intervals. The last column displays the t-statistic that tests the null of equality of the high and low inflation regime coefficients,

[15] The moving average is our benchmark smoothing procedure. However, the results from the local projections are very similar with a HP-Filter ($\lambda = 14400$) smoothing procedure, see the Appendix.

[16] This calibration is also in accordance with Álvarez *et al.* (2019) who show that the frequency of price changes starts increasing significantly from annual inflation rates of 5%. Our smooth transition function indicates a value of approximately 0.5 with such an annualised inflation rate.

i.e., $\hat{\beta}_h^{HI} = \hat{\beta}_h^{LO}$, where the grey area represents the 90% z-values. A positive value means that the high inflation response is larger whereas a negative value of the t-statistic indicates the opposite. First, the linear terms for PCE inflation in the top left panel show the familiar picture in the literature. PCE Inflation declines eventually, after an initial positive response, which however is not statistically significant. Second, inflation in a high inflation regime declines right away on impact after a contractionary monetary policy shock. On the contrary, the impulse response function in a low inflation regime exhibits a price puzzle for about one year, which is marginally statistically significant only for few months. This suggests that in this regime firms are not willing to change price as frequently, so the price level stays persistently around zero for a longer period. Moreover, at long horizons the response is smaller in a high trend inflation regime. Again, this is consistent with the idea that the effect is less persistent in a high inflation regime, because more firms adjust on impact after the shock. Third, the last column shows that the responses of inflation in a high and low inflation regime are statistically significantly different both at short and at long horizons. We interpret these results as evidence in favour of state-dependent prices models as key propagation mechanism of monetary policy shocks, because they predict a faster, and less persistent, reaction to a monetary disturbance in a high trend inflation regime. This is exactly what the impulse response functions show.

The first panel in the second row shows that the IRFs for output exhibit the usual hump shaped dynamics. Output reacts with a larger delay in a low inflation regime compared to a high inflation regime, but this reaction is stronger and reaches a trough after two years which is roughly twice as deep as the one in a high inflation regime. The difference in the IRFs between the two regimes however is not statistically significant, so that we do not find evidence for our second theoretical implication regarding output. Table 2 displays similar results for the cumulated IRFs of inflation and output at different horizons. Prices drop from the outset in a high inflation regime, while the reaction is sluggish in a low inflation regime. However, the latter is more persistent so that it eventually catches up and overtakes the cumulative drop in a high inflation regime, so that the cumulative drop after four years is 50% higher. The difference in the initial response, up to two years, is statistically significant providing supporting evidence for state-dependent pricing. This is not the case for the cumulative response of output, that is both not significant for low inflation regimes and not statistically different between low and high inflation regime. Again, the point estimates are in line with the theoretical prediction, but the standard errors get so large (especially for the low inflation regime, as evident also form Figure 4) that these differences are not significant.

Finally, the panels in the third row show that the interest rate increases after a monetary policy shock and stays positive for about two years, in the linear case and low inflation regime, while for only one year in the high inflation regime. Thus, monetary policy initially reacts differently to a shock in the two regimes, and these differences are statistically significant for the first two years. In a high inflation regime, the nominal interest rate initially reacts more, but then it decreases much faster than in a low inflation regime. While one might argue that this pattern may explain the initially quicker reaction of the price level in the high inflation regime, prices in the low regime react considerably more sluggishly even though the interest rate is positive for a longer period of time. Indeed, the last column shows that for most of the IRFs the coefficients in the low inflation regime are larger than the ones in a high inflation regime, perhaps signalling a stronger endogenous feedback of monetary policy in response to the shock (as also evident from the IRFs in column one). On the one hand, it would be hard to argue that the different monetary policy behaviour is driving the different response of prices between the high and low inflation regimes. On the other hand, the fact that the path of the federal funds rate after the initial monetary contraction is different across the two regimes blurs the comparison, possibly explaining why the evidence for the differences in output responses is not statistically significant.

To conclude, we find evidence in favour of state-dependent prices regarding our second testable implication with respect to the behaviour of inflation. Our results show that in a high trend inflation regime, inflation declines right away after a policy shock, and there is no price puzzle, as theory predicts. Moreover, inflation is more persistent in a low inflation regime, despite the interest rate staying positive for a longer amount of time. Regarding the response of output, however, the point estimates are coherent with the theoretical prediction, but the differences between the high and low inflation regime are not statistically significant.

5 Robustness

This Section reports the results of two particularly important robustness exercises regarding our local projection estimates. We conduct comprehensive robustness checks on many other dimensions, that we confine to the Appendix and we briefly summarise below.

5.1 Stability

It is crucial for our analysis to be relatively confident that the non-linear and state-dependent dynamic behavior of the impulse response functions of output and inflation is not driven by a feedback effect of monetary policy.[17] To show that monetary policy feedback plays a limited role with respect to large and small shocks, we check if inflation coefficients with respect to the linear and absolute value interaction term do not exhibit structural breaks when monetary policy conduct may have changed. If these estimates stay stable over a variety of monetary policy regimes throughout time then we can reject

[17] As discussed in section 4.1, several of our results seem to contradict this possibility. For example, monetary policy seems to react weakly after a large shock whereas prices react by more at the beginning of the horizon. The same applies to a large extent to the test of the second implication in section 4.2 as the funds rate stays positive for longer in the regime where prices seem to be more sticky. These patterns seem to indicate that the dynamics of inflation and output are not defined by the reaction of monetary authorities only.

the thesis of monetary feedback driving the result of size-dependent effects.

We apply the Hansen (1992) stability test to the coefficients in the local projection (1) for PCE inflation. This test has locally optimal power and needs no *a priori* assumption concerning the breakpoint. Furthermore, this test is robust to heteroskedasticity, a potential concern in this analysis. Table 3 shows the results for the two coefficients of interest and the joint test for parameter stability at a 1, 3, 6, 12, 24 and 36 month horizon for our PCE inflation local projection.[18] We cannot reject the null of individual parameter constancy for neither the linear nor the absolute value interaction coefficients at these horizons (with the exception of the linear coefficient at 36 month horizon). However, the joint test statistic does indicate a rejection of the null that all parameters in the local projection are constant. This suggests that, even though the dynamic feedback of monetary policy via lagged control values may have changed throughout time, the shape and non-linearity of the impulse response after a monetary policy shock has stayed relatively constant throughout the sample. The same results hold for the linear and interaction coefficients when the dependent variable is either the industrial production or the Fed Funds rate (see Appendix).[19]

5.2 Excluding the NBR targeting period

Coibion (2012) and others have suggested that the exclusion of the NBR targeting period October 1979 and September 1982 can account for a difference in results between the Romer and Romer (2004) and VAR approach. Critically, the largest absolute monetary shock values lie in this period and may thus play a significant role for the conclusion on

[18]For a test on an individual coefficient, we can reject the null of parameter constancy at the 5% significance level, if the relevant test-statistic is larger than the asymptotic critical value of 0.47. The null hypothesis is that each coefficient in (1) is constant and the respective distribution is non-standard and depends on the number of parameters tested for stability. The intuition is that under the null hypothesis the cumulative sums of the first-order conditions from the estimation will have mean zero and wander around zero. However, under the alternative hypothesis of parameter instability, these first-order conditions will not be mean zero for parts of the sample and so the test statistic will be large, leading us to reject the null (see Appendix).

[19]These results substantiate the visual impression from the plots of the recursive estimation of the local projection coefficients displayed in the Appendix.

our first implication. In order to account for this suggestion we exclude this part of the shock sample, modifying the non-linear local projection equation (1) by interacting the linear and the non-linear coefficient with a time-dummy that takes a value of 1 for the sample between October 1979 and September 1982 and 0 otherwise. As in Figure 2, the three rows in Figure 5 display the IRFs for this specification for PCE inflation, industrial production and the Fed Funds rate, respectively, for both the coefficients of the linear term, $\hat{\beta}_h$ (green, solid line), and the ones on the non-linear absolute value interaction term, $\hat{\zeta}_h$ (blue, dashed line). The impulse response coefficients are qualitatively similar to the benchmark results. The results, however, change somewhat both in terms of magnitude and, especially, in terms of significance, particularly so for the interaction coefficients. This is hardly surprising. The reduced sample does not include the large shocks of the NBR targeting period, and it features low sample variation with values mostly below 1. Hence, the non-linear effects of large shocks are more difficult to identify, inducing wide confidence intervals.

5.3 Other robustness checks

The Appendix presents comprehensive robustness checks. It displays the plots of the recursive estimates of the local projection coefficients in (1), to further visually inspect their stability. Besides, results are robust to a different specification of the non-linear terms in (1), that includes a squared and a cubed term for the shock value, rather than the absolute value interaction term.

Results for the smooth transition local projections (2) are robust to: 1) changes in γ and c; 2) using HP-filtered inflation for z; 3) using a model based trend inflation measure from Ireland (2007).

Finally, the estimates of the coefficients in both (1) and (2) are also robust to: 1) using the CPI instead of the PCE; 2) using quarterly data with GDP as the output measure; 3) including as controls: (i) the commodity price index, (ii) the corporate bond credit spread by Gilchrist and Zakrajsek (2012) to control for financial frictions, (iii) proxies for fiscal policy using either measure of excess returns on stocks of military contractors from Fisher and Peters (2010) or the exogenous tax changes from Romer and Romer (2010); 3) different measures of shocks obtained from non-linear models: (i) from the Romer and Romer (2004) regression using the smooth transition function; (ii) from a smooth transition VAR; 4) including leads and lags of the shocks to control for potential autocorrelation and misspecification as suggested by Alloza *et al.* (2019).

6 Conclusion

The assumption of sticky prices lies at the very center of the current workhorse model for the analysis of business cycle fluctuations and, particularly, monetary policy effects. The literature features two type of sticky price models: time-dependent and state-dependent prices. A sticky price theory of the transmission mechanism of monetary policy shock based on state-dependent pricing yields two testable implications, that do not hold in time-dependent models; the impulse response function of the aggregate price level and inflation should be more flexible both after a large shock and during high trend inflation regimes. Employing the methodology of local projections, we tested these predictions on aggregate US data. We found some evidence in favour of state-dependent models of price stickiness rather than time-dependent ones. With regards to the response to large shocks, the coefficient of the absolute value interaction shock projections matched our theoretical prior, both in terms of output and inflation. When the NBR targeting period of US monetary policy, October 1979 and September 1982, is taken out of the sample, our results loose statistical significance, because too little variation is present to identify the non-linear effect and the significance bands becomes very wide. The empirical investigation during large trend inflation regimes also showed that inflation reacts significantly more quickly to a monetary policy shock in times of high trend inflation. In this case, the evidence for output is not significant, however, the point estimates behave according to the theoretical predictions. Furthermore, the Appendix shows that the results are robust to a very large variety of robustness tests.

Our results are the first ones (to the best of our knowledge) that point towards a significant presence of state-dependent pricing in the US economy using aggregate data. Our macro evidence is a useful complement to the large empirical literature on individual firm price micro data. Hence, *'Prices are sticky after all'* - just as recent literature has shown (Kehoe and Midrigan, 2015) - but less so when shocks are large or inflation is high. These results are in line with what would be expected if state-dependent pricing played a significant role in the US economy. This supports the theoretical implication that the frequency of changing prices is, at least to some extent, endogenous to the economic environment, as in Álvarez *et al.* (2017). So, although the Calvo (1983) model may work quite well for "normal" times, when considering situations where high-trend inflation is present or large shocks are likely, a state-dependent sticky price framework that accounts for these phenomena seems more appropriate (as e.g., Álvarez and Lippi, 2014; Álvarez *et al.*, 2016; Costain and Nakov, 2011, 2019; Costain *et al.*, 2019).

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Figures

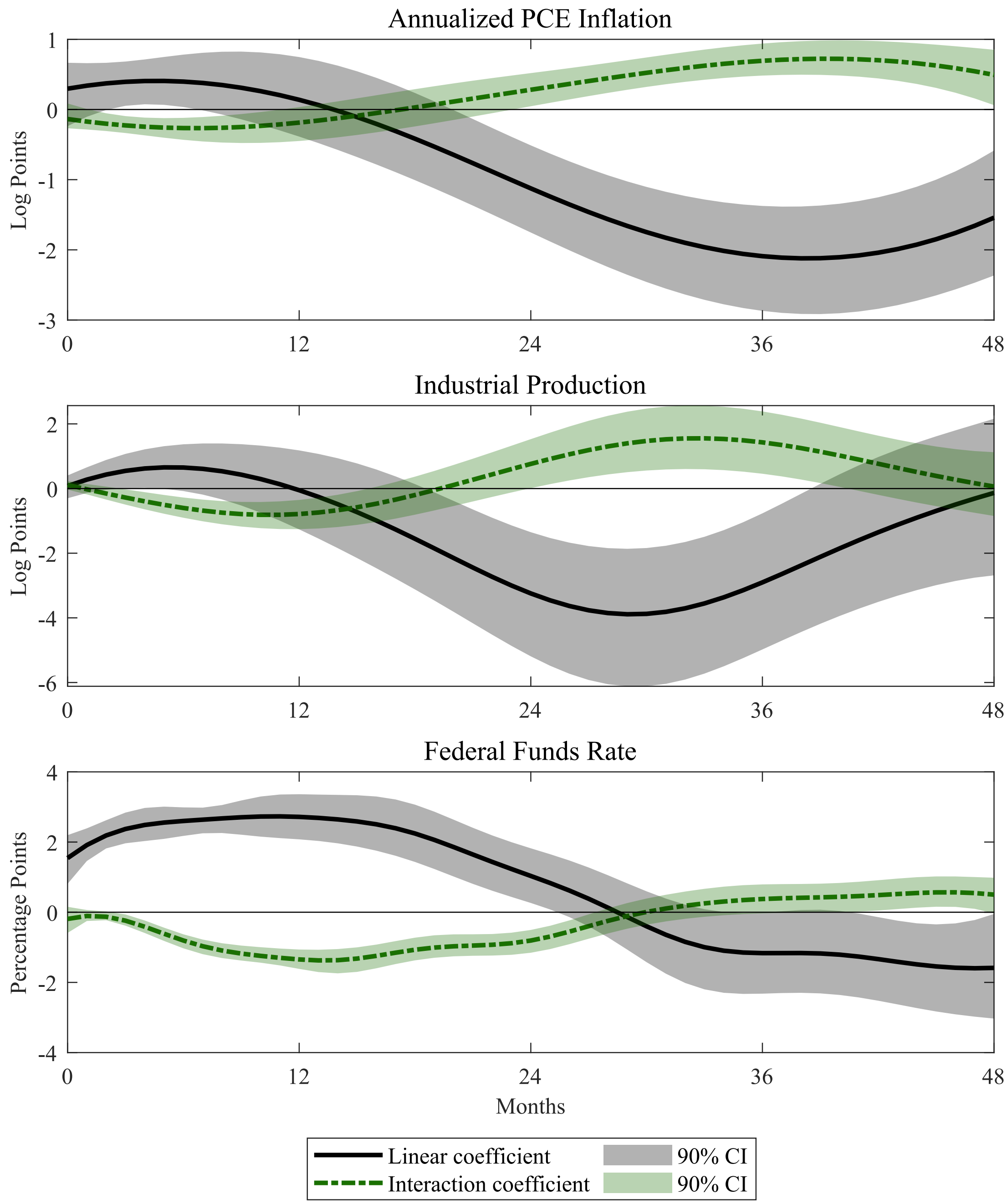

Figure 1: Panel of smooth local projection coefficients for annualised PCE inflation, industrial production and the federal funds rate. Every panel depicts both the point estimates of the linear coefficient (solid line) and the absolute value interaction coefficient (dashed-dotted), together with their 90% confidence intervals for the various response variables. The coefficients are depicted over a four year horizon.

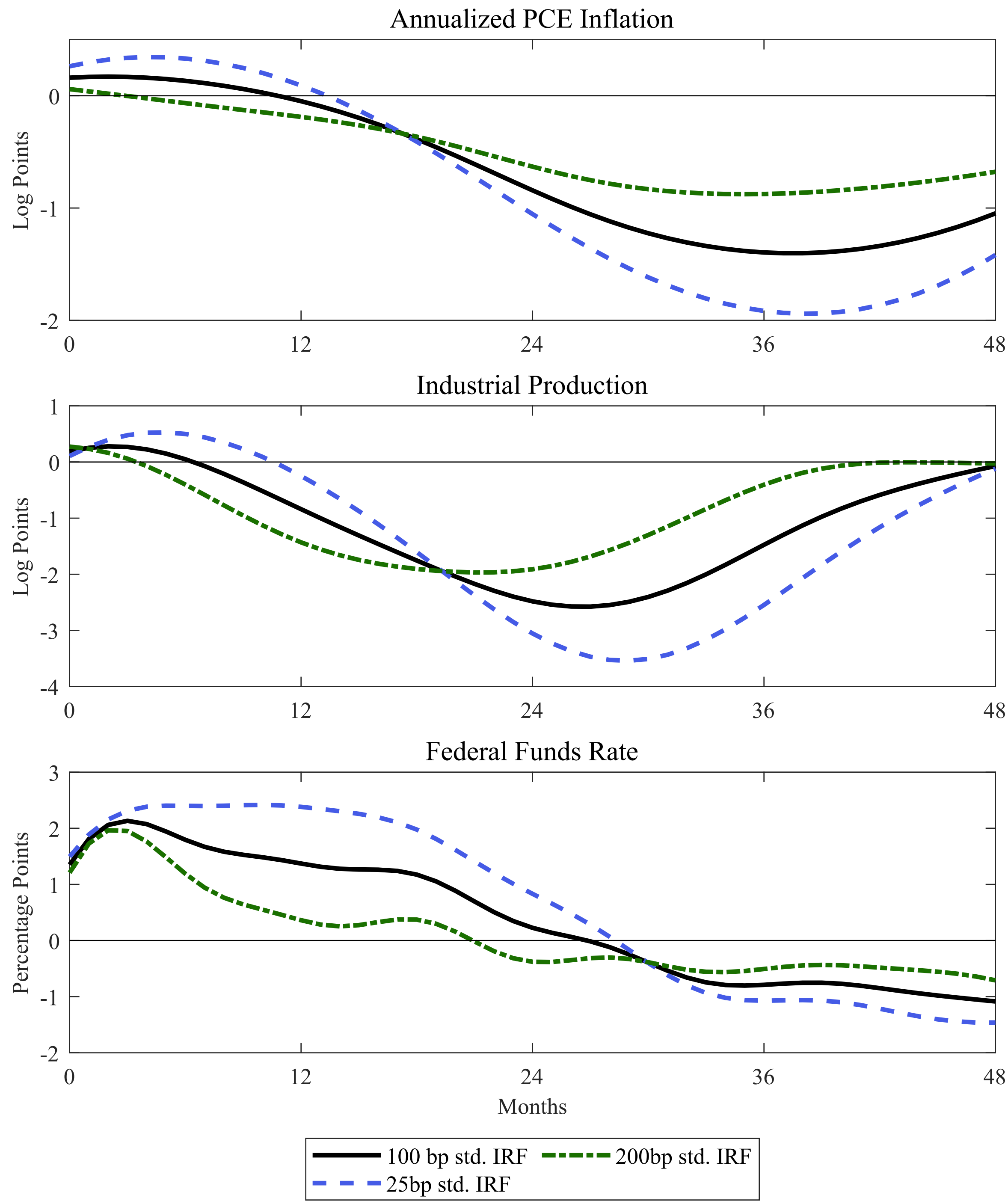

Figure 2: Panel of simulated size-dependent impulse responses for annualised PCE Inflation, Industrial Production and the Federal Funds Rate over a four year horizon. The Figure depicts the impulse response for a 25 (dashed line), 100 (solid) and 200 (dashed-dotted) basis point shock, rescaled by dividing by the size of the shock. The impulse responses are depicted over a four year horizon.

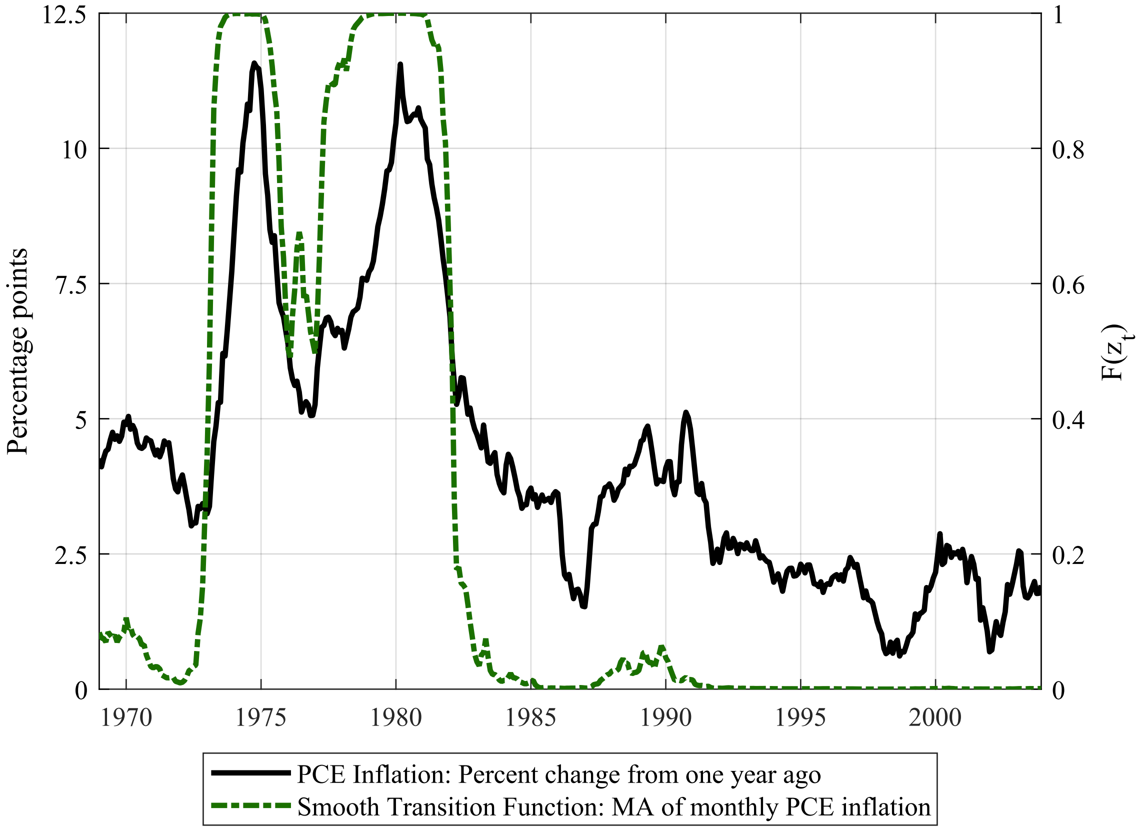

Figure 3: The solid line plots PCE Inflation, measured in percentage change from one year ago on the left axis. The dashed line shows the benchmark smooth transition function based on the two year centred moving average of monthly PCE inflation, measured on the right axis. Both are depicted from January 1969 until December 2003.

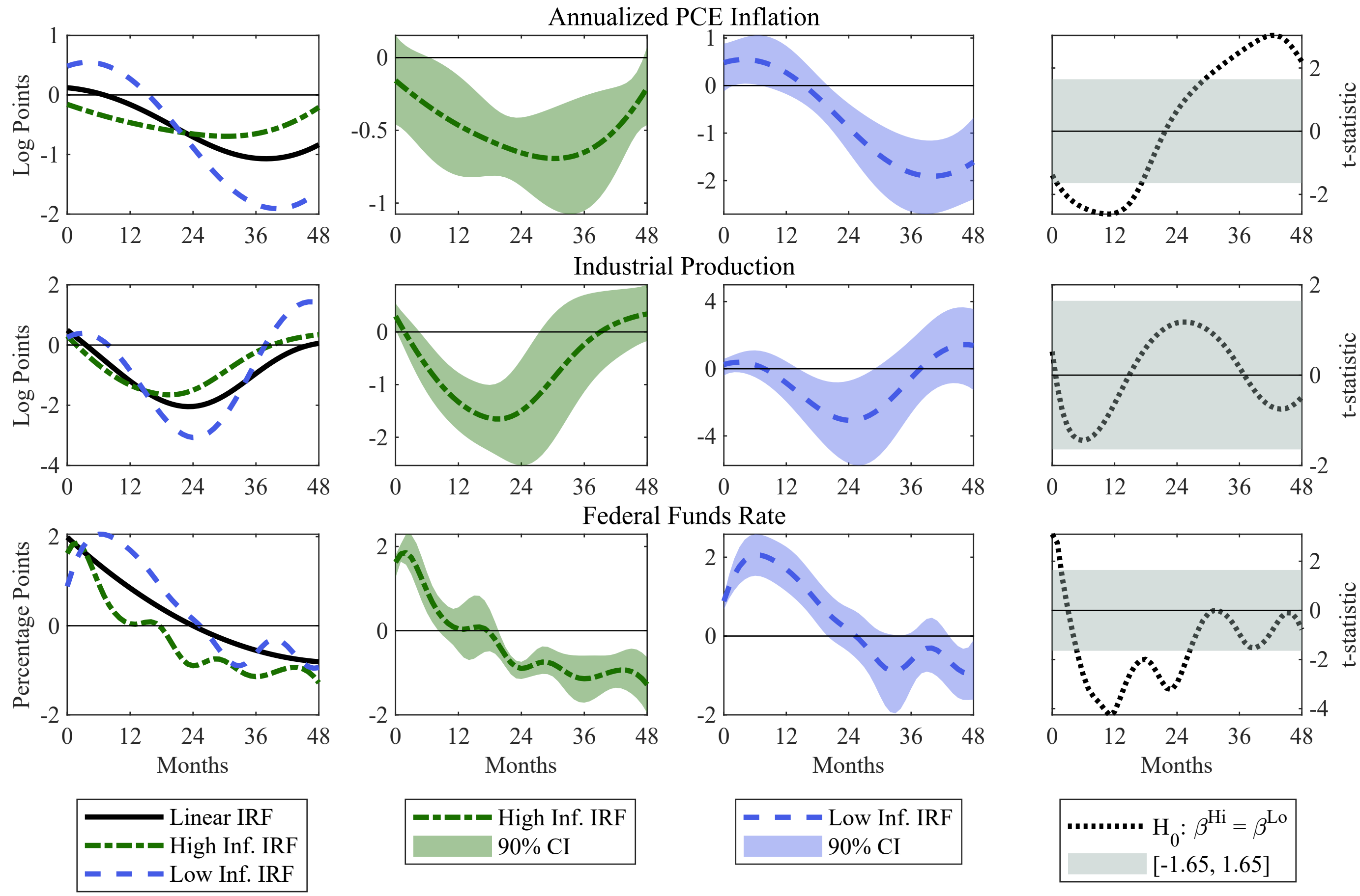

Figure 4: Panel of smooth impulse responses in different inflation states for annualised PCE inflation, industrial production and the federal funds rate. Column 1 depicts the point estimates for the linear (solid line), high inflation (dash-dotted) and low inflation (dashed) impulse response. Column 2 and 3 depict the high inflation and low inflation impulse responses, together with their 90% confidence bands. Column 4 depicts the t-statistic for the null hypothesis of equality of the high and low inflation responses (dotted), together with 90% z-values (shaded area). The impulse responses are depicted over a four year horizon.

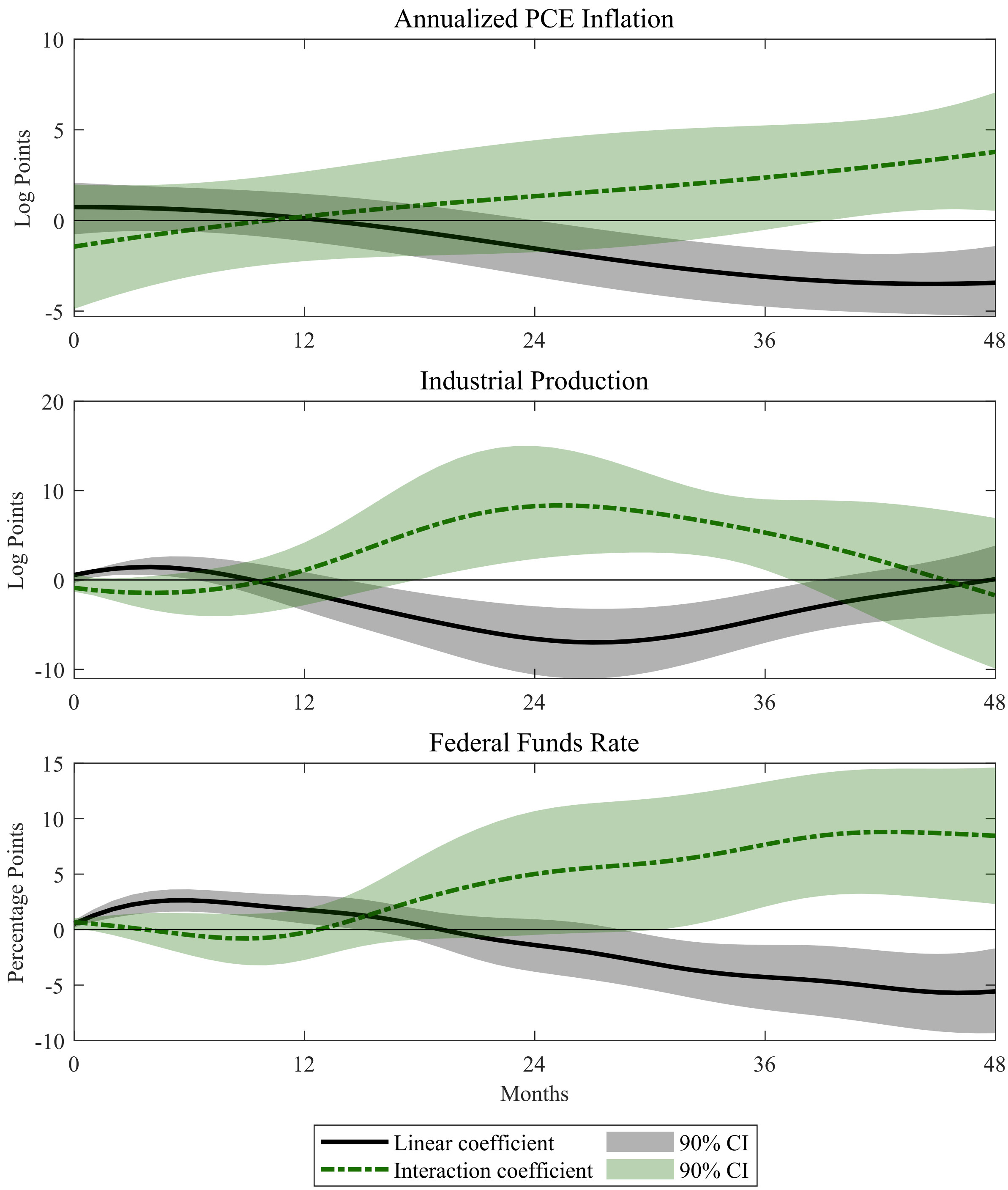

Figure 5: Exclusion of the NBR targeting period between October 1979 and September 1982. Panel of smooth local projection coefficients for annualised PCE Inflation, Industrial Production and the Federal Funds rate. Every panel depicts both the point estimates of the linear coefficient (solid line) and the absolute value interaction coefficient (dashed-dotted), together with their 90% confidence intervals for the various response variables. The coefficients are depicted over a four year horizon.

Tables

Standardized cumulative impulse responses for a 25bp and 200bp shock

PCE Deflator					
Horizon	$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
$\hat{\beta}_h + \hat{\zeta}_h \vert 0.25\vert$	0.02	0.23	-0.16	-1.75^{**}	-3.57^{***}
	(0.03)	(0.21)	(0.51)	(0.81)	(1.13)
$\hat{\beta}_h + \hat{\zeta}_h \vert 2.00\vert$	-0.03^{***}	-0.17^{*}	-0.43^{**}	-1.20^{***}	-1.83^{***}
	(0.01)	(0.09)	(0.20)	(0.41)	(0.57)

Cumulative Industrial Production					
Horizon	$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
$\hat{\beta}_h + \hat{\zeta}_h \vert 0.25\vert$	0.18	3.29	-16.32	-57.63^{**}	-69.67^{**}
	(0.15)	(4.54)	(14.31)	(25.43)	(34.13)
$\hat{\beta}_h + \hat{\zeta}_h \vert 2.00\vert$	0.25^{***}	-7.75^{***}	-30.13^{***}	-40.53^{***}	-38.98^{***}
	(0.07)	(1.64)	(5.37)	(8.52)	(11.54)

Table 1: Impulse responses of cumulative PCE inflation, ie. the PCE Deflator, and cumulative industrial production after a 25bp and a 200 bp point shock, standardized by the respective size of the shock. Newey-West standard errors in parentheses.
****: Significant at the 1% level; **: Significant at the 5% level; *: Significant at the 10% level*

Cumulative impulse responses in high and low inflation regimes

			PCE Deflator		
Horizon	$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
$\hat{\beta}_h^{HI}$	-0.06^{***}	-0.40^{**}	-0.92^{***}	-1.60^{***}	-2.06^{***}
	(0.02)	(0.19)	(0.25)	(0.20)	(0.24)
$\hat{\beta}_h^{LO}$	0.08^{*}	0.49	0.39	-1.31^{**}	-3.10^{***}
	(0.05)	(0.34)	(0.51)	(0.57)	(0.83)
t-stat: $\beta_h^{HI} - \beta_h^{LO}$	-2.60^{**}	-2.21^{**}	-2.61^{**}	-0.47	1.20

			Cumulative Industrial Production		
Horizon	$h = 1$	$h = 12$	$h = 24$	$h = 36$	$h = 48$
$\hat{\beta}_h^{HI}$	0.06	-8.49^{***}	-27.14^{***}	-36.40^{***}	-34.92^{***}
	(0.07)	(2.75)	(5.35)	(9.74)	(13.32)
$\hat{\beta}_h^{LO}$	0.25	1.87	-26.22	-54.52	-41.12
	(0.29)	(5.99)	(18.03)	(35.77)	(41.27)
t-stat: $\beta_h^{HI} - \beta_h^{LO}$	-0.60	-1.39	-0.05	0.51	0.15

Table 2: Impulse responses of cumulative PCE inflation, ie. the PCE Deflator, and cumulative industrial production in the high and low inflation regimes. The last rows report the t-statistic with the null hypothesis of the two coefficients being equal. Newey-West standard errors are in parentheses.
****: Significant at the 1% level, **: Significant at the 5% level; *: Significant at the 10% level*

PCE Inflation Local Projection - Hansen (1992) test statistic

Horizon	$h = 1$	$h = 3$	$h = 6$	$h = 12$	$h = 24$	$h = 36$
Linear coeff.	0.10	0.04	0.08	0.04	0.09	0.54^{**}
Interaction coeff.	0.01	0.03	0.02	0.01	0.01	0.06
Joint: all coeffs.	3.13^{**}	3.04^{**}	3.68^{***}	5.88^{***}	9.43^{***}	9.76^{***}

Table 3: Estimated Hansen (1992) test statistics for parameter constancy of the PCE inflation local projection with both a linear and an absolute value interaction shock term. The first row reports the individual test statistic for the linear coefficient $\hat{\beta}$ at different horizons , the second row reports those for the absolute value interaction coefficient $\hat{\zeta}$ and the final row reports the test statistic for the joint hypothesis of all parameters (ie. regression coefficients and variance) to be constant.
Individual critical values are $c_{1,1\%} = 0.75$, $c_{1,5\%} = 0.47$ and $c_{1,10\%} = 0.35$. Joint critical values for a model with $K = 12$ parameters are $c_{12,1\%} = 3.51$, $c_{12,5\%} = 2.96$ and $c_{12,10\%} = 2.69$.
****: Significant at the 1% level; **: Significant at the 5% level; *: Significant at the 10% level*