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The Price Puzzle and the Cost Channel in Monthly and Quarterly Data

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Abstract
This paper addresses the so-called 'price puzzle'—a positive response of prices to a monetary contraction. Using both reduced-form and structural models, it demonstrates that measures of output gap can be used to estimate the New Keynesian Phillips curve (NKPC) at both monthly and quarterly frequencies. It also examines whether the cost channel of transmission of monetary policy can account for the price puzzle and documents its presence in both monthly and quarterly models. Estimation results indicate that the larger price puzzle in the monthly data cannot be attributed to the stronger presence of the cost channel. Finally, this paper shows that although the estimated extent of cost channel's presence is sensitive to whether the discount factor is estimated or imposed, the cost channel cannot explain the price puzzle in any setting considered below.

JEL Categories: E31; E32; E52

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1 Introduction

Concern with a positive response of prices to a contractionary monetary shock can be traced several decades back. Widely cited in the empirical literature on this subject is the 1970s comment of Congressman Wright Pitman that fighting inflation with higher interest rates was akin to "throwing gasoline on fire." His simile had little impact on the conduct of monetary policy, however, as the standard Keynesian models confidently predicted that an increase in interest rates would reduce aggregate demand and hence the price level, without considering its possible effects on aggregate supply. Starting with the seminal paper by Sims (1992), academic interest in the 'price puzzle'—a positive response of prices to a contractionary monetary policy shock—was resurrected.

The price puzzle is typically addressed in two ways: Finding model specifications that resolve it and imply that the puzzle doesn't exist, or finding modeling devices that provide theoretical substantiation for the puzzle. One such device is the presence of the cost channel of transmission of monetary policy, whereby interest rates enter a representative firm's marginal cost function. In this setup, a contractionary monetary policy shock raises interest rates and hence the firm's marginal cost. In the short term, this increase in cost translates into an increase in prices, which later decline due to the decrease in aggregate demand that results from higher interest rates. Hence models that incorporate the cost channel may be able to explain the price puzzle. More specifically, the cost channel is modeled as the nominal interest rate term entering the New Keynesian Phillips curve. One purpose of this paper is to investigate the strength of this relationship in the monthly and quarterly U.S. data.

There is no consensus in the literature on whether prices exhibit a positive response to a contractionary monetary policy shock or whether the cost channel plays an important role in the business cycles. As the empirical literature review presented in Section 2 documents, different datasets and methodologies seem to point to different conclusions. In particular, it appears that results based on quarterly data tend to resolve the price puzzle more easily, whereas the ones that rely on monthly data generally support it. The present work corroborates this result and investigates whether it can be attributed to the stronger presence of the cost channel in the monthly rather than quarterly data.

The rest of the paper is organized as follows. Section 2 surveys empirical literature that addresses the price puzzle and the cost channel. Section 3 uses the standard structural vector autoregressive (SVAR) evidence to document the existence of the price puzzle in the U.S. data. Section 4 presents empirical results based on a small-scale structural model that demonstrate that the cost channel of transmission of monetary policy cannot be a suitable explanation for the price puzzle. Lastly, Section 5 concludes.

2 Review of the Empirical Literature

Recent interest in the 'price puzzle' began with the Sims' (1992) finding that prices responded positively to an increase in interest rates in several industrialized countries. He also found that introducing an index of commodity prices into the empirical system helped reduce the extent of the price puzzle, which led him to the conjecture that central banks may use 'information variables' that indicate the advent of inflation and allow them to react preemptively. He then suggested that failure to include these variables into an empirical system results in a misspecified model; correcting this misspecification would then remove the 'price puzzle'. These findings produced considerable controversy, prompting researchers to look for explanations of this ostensible anomaly. The proposed solutions draw upon a wide array of methodologies and offer rather different answers. However, one distinction seems to be pervasive: Studies that use monthly datasets find it hard to solve the price puzzle, whereas the ones that use quarterly data frequently suggest that it may not be an issue altogether.

Bernanke and Mihov (1998) developed a general model of monetary policy that nests several alternative assumptions, such as the central bank's targeting the federal funds rate, or total reserves, or non-borrowed reserves. This setup is used as a part of a general model of an economy and describes the variables' responses to monetary policy innovations. They provided evidence that a change in the conduct of monetary policy occurred in the early 1980s.[1] Hanson (2004) employed

[1] Bagliano and Favero (1998) re-examined and largely upheld these findings.

the Bernanke and Mihov (1998) empirical model of monetary policy and surveyed a wide gamut of potential 'information variables' in the monthly US data. Although some of these variables help alleviate the extent of the price puzzle, especially in the more recent sample, several specifications indicate that it still may be a palpable issue, especially in the montly case. More recently, Gallès and Portier (2006) use quarterly and mixed-frequency VAR models to demonstrate that the price puzzle is less pronounced in the quarterly data. Results presented in this paper point to similar conclusions.

On the other hand, studies using quarterly data at the macro-level indicate that the price puzzle may not exist at all. Whereas including 'information variables' may be one way of addressing the potential misspecification problem that could produce the price puzzle, Giordani (2004) points to another potential source of misspecification and shows that the price puzzle disappears once a measure of output is replaced with a measure of the output gap, whose presence is motivated by theory. He rejects the necessity of 'information variables' and claims that they contribute to resolving the price puzzle, only because they are correlated with other targets of a central bank. In particular, he shows that once output is replaced by the output gap in a three-variable VAR (output gap, inflation, and the federal funds rate), the price puzzle disappears in the quarterly data. However, Giordani's (2004) brief inspection of the monthly data acknowledges that the price puzzle is harder to resolve at the monthly frequency. He attributes this effect to measurement errors in the monthly data. This paper questions his results by showing that using the same measure of output gap as his still results in the price puzzle.

Castelnuovo and Surico (2005) examine the VAR models of the price puzzle price at different time periods and show that, in the quarterly data, the price puzzle existed in the 1966-1979 sample but was absent in the 1979-2002 sample. Samples that span these two time periods are likely to produce some behavior consistent with the price puzzle. The authors explain the discrepancy in the results from these two subsamples by the difference in the conduct of monetary policy: Insufficiently tight monetary policy may result in indeterminacy, which, as they demonstrate with the simulated data, produces the "price puzzle" impulse responses.[2] Contrary to their VAR findings, the present paper shows that the price puzzle still exists even in the more recent sample.

Micro-level studies also suggest that the price puzzle indeed exists and propose the cost channel of transmission of monetary policy as its solution. Gaiotti and Secchi (2006) find that it plays a significant role in the data from 2,000 Italian manufacturing firms and that the parameter describing the extent of the cost channel is large and significant. Working with the U.S. industry-level manufacturing data, Barth and Ramey (2000) find that a contractionary monetary policy shock produces lower output and higher price-wage ratios.[3] They also show that this effect is much stronger in the time period up to 1979.[4] Therefore, it appears that manufacturing firms indeed tend to respond to a contractionary shock by raising prices.

Small-scale structural explanations of the price puzzle frequently rely on the cost channel of transmission of monetary policy whereby the nominal interest rate enters a representative firm's marginal cost function. Thus, a monetary contraction raises the firm's marginal cost and, if able to reset the price on its product in the Calvo (1983) pricing setup, the firm will raise it, which leads to behavior consistent with the cost channel. Ravenna and Walsh (2006) find that the cost channel has significant presence in the U.S. quarterly data and Chowdhury, Hoffman, and Schabert (2005) provide further support from the international quarterly data. Both of these studies employ the generalized method of moments to show this significance.[5] Castelnuovo (2006), however, questions the validity of their results, as their estimated models lack persistence in aggregate demand and interest rate rules; the present paper incorporates his criticism. Rabanal (2003) uses Bayesian methods to estimate a model with sticky prices and a real wage rigidity and finds that the extent of

[2]Tightness of monetary policy is measured in the spirit of Clarida, Galí and Gertler (1999): To produce determinacy in the New Keynesian model, the coefficient on expected future inflation should be greater than 1 in the interest rate rule. Clarida, Galí, and Gertler (1999) also show that the this coefficient was less than 1 in the pre-1980 subsample, which is consistent with the results of Castelnuovo and Surico (2005).

[3]They also provide evidence that this effect is primarily due to higher prices rather than lower wages.

[4]Hanson (2004) reaches a similar conclusion. Arguably, much of the puzzle in that dataset can be driven by the accommodating monetary policy the Fed was pursuing prior to the 1980s, which is documented by Clarida, Galí, and Gertler (1999).

[5]Both of these papers rely on a New Keynesian model with Calvo pricing where the probability of prices being reset is determined exogenously. Ceniz (2006) shows that similar dynamics can be obtained in a model where the frequency of price adjustments endogenously increases in the nominal interest rate.

the cost channel's presence is negligible. He advances the idea that the cost channel may be confused with countercyclical mark-up shocks. Barth and Ramey (2000), however, counter that argument by questioning the ad hoc nature of such mark-up shocks and stressing that the cost channel produces behavior that is consistent with the theory of countercyclical mark-ups. Kaufmann and Scharler (2006) estimate a model that features incomplete pass-through of monetary shocks to producers and conclude that the extent of the cost channel is small in both European and U.S. data. In a related paper that also features a similar form of incomplete pass-through, Hülsewig, Mayer and Wollmerhäuser (2006) find evidence in favor of the cost channel's presence that bears, however, quantitatively small impact. Furthermore, Tillmann (2006) presents evidence that augmenting the Phillips curve to include the cost channel significantly reduces the difference between actual inflation and estimates of fundamental inflation in the U.S., U.K., and Euro area data. However, he presents no estimates of the cost channel's extent.

Conclusions regarding cost channel's ability to produce the price puzzle are markedly different. At one end of the spectrum, Ravenna and Walsh (2006) and Chowdhury, Hoffman, and Schabert (2005) produce estimates that support this possibility. On the other hand, the results of Rabanal (2003) strongly reject it. Section 4 demonstrates how these different estimates can arise, given the differences in assumptions that the two sets of authors use. However, this paper uses methodology similar to those of Ravenna and Walsh (2006) and Chowdhury, Hoffman, and Schabert (2005) to show that reasonable estimates of the cost channel are low and even if they were higher, it would still not be able to account for the price puzzle. Furthermore, the cost channel cannot provide an explanation for the price puzzle showing up stronger in the monthlay rather than quarterly data.

3 Documenting the Price Puzzle

This section offers a brief documentation of the price puzzle typically found in the SVAR literature along the lines of Hanson (2004) and Giordani (2004). The model takes the following form:

$$\mathbf{A_0 x_t} = \mathbf{A(L) x_{t-k}} + \epsilon_{\mathbf{t}}, \tag{1}$$

where $\mathbf{x_t}$ is the $n \times 1$ vector of variables, $\mathbf{A_0}$ is the matrix describing the contemporaneous relationship between the variables (with the lead diagonals of 1's) , L is the lag operator, and $\epsilon_\mathbf{t}$ is the vector of uncorrelated structural errors that are assumed to be white noise. In practice, it is convenient to estimate the reduced-form version of this model:

$$\mathbf{x_t} = \mathbf{A_0^{-1}} \mathbf{A(L)} \mathbf{x_{t-k}} + \mathbf{A_0^{-1}} \epsilon_\mathbf{t}, \tag{2}$$

and then impose necessary restrictions to recover the structural parameters. Errors from the reduced form model, $\mathbf{e_t}$, are related to the structural shocks by: $\mathbf{A_0 e_t} = \epsilon_\mathbf{t}$, or, defining $\mathbf{\Lambda} = \mathbf{A_0^{-1}}$:

$$\mathbf{e_t} = \mathbf{\Lambda} \epsilon_\mathbf{t}. \tag{3}$$

Imposing restrictions on $\mathbf{\Lambda}$ makes it possible to identify the SVAR system that consists of measures of output gap, inflation, and nominal interest rate.

[Insert Figure 1 about here]

The variables in the model have the standard lower-triangular ordering, with a measure of the output gap first and the nominal interest rate last. This specification assumes that the nominal interest rate is the most endogenous variable that can contemporaneously respond to output gap and inflation, inflation can contemporaneously respond only to the output gap, and the latter responds to both other variables with a lag. An alternative identification scheme, used by Bernanke and Mihov (1998) and Bagliano and Favero (1998) among others, models the monetary shocks more explicitly but yields very similar quantitative and qualitative results.[6] Two measures of output gap available at the monthly frequency are used: linearly detrended log index of industrial production

[6] In their setup, variables are subdivided into two blocks. First, non-policy variables, $\mathbf{Y_t}$, include a measure of output gap and inflation. The second block consists of policy variables, $\mathbf{P_t}$: federal funds rate, total reserves, and non-borrowed reserves. As was mentioned in Section 2, they argue that this setup produces a monetary policy shock that originates in the reserves market rather than as the portion of nominal interest unexplained by macroeconomic variables as in Sims (1992) or Giordani (2004). Recursive structure of the policy block detailed below is consistent with a FFR-targeting regime. Matrix $\mathbf{\Lambda}$ is partitioned with its segments representing the relationships between different blocks of variables:

$$\mathbf{\Lambda} = \begin{bmatrix} \mathbf{\Lambda_{YY}} & \mathbf{\Lambda_{YP}} \\ \mathbf{\Lambda_{PY}} & \mathbf{\Lambda_{PP}} \end{bmatrix}. \tag{4}$$

(IP) and the percentage deviation of capacity utilization (CU) from its mean.[7] This comparison is interesting because Giordani (2004) claims that the latter is a better proxy of the output gap and using it instead of the IP resolves the price puzzle. Figure 1 plots the time series for these measures of the output gap and suggests some differences between the two. Inflation is measured as the annualized difference in log CPI. Finally, the Federal Funds rate is the measure of nominal interest rates. The data cover the 1980:1—2006:3 time period. Given the results in Castelnuovo and Surico (2005), this should limit the scope of the price puzzle. To make results comparable to Giordani (2004), models with monthly data are estimated with 6 lags, whereas quarterly models are estimated with 2 lags. Note that the results discussed below are robust to the lag structure of these VARs.

[Insert Figure 2 about here]

Figure 2 describes the difference in how the price puzzle appears in the monthly and quarterly data. Both measures of the output gap deliver similar predictions about the initial effect of a monetary contraction. In the quarterly data, inflation initially increases by about 0.7% and its impulse response stays positive and significant for about 3 quarters. Monthly results are also close for both measures: Inflation initially increases by 1.2%, the response stays significantly positive for about 4 months, and becomes insignificant thereafter.[8] Therefore, both of these models unequivocally point that the price puzzle exists in this sample and the Giordani (2004) suggestion that switching

Traditionally, $\mathbf{\Lambda_{YY}}$ has lower triangular ordering (see Bagliano and Favero (1998)). Hanson (2004) also imposes lower-triangular structure on $\mathbf{\Lambda_{PP}}$, with the federal funds rate placed first. FFR targeting implies that the federal funds rate is placed first, whereas the ordering of the two measures of reserves is unimportant. Policy is assumed to respond to current macroeconomic conditions, hence parameters in $\mathbf{\Lambda_{PY}}$ need to be estimated freely. Conversely, macroeconomic variables are assumed to respond to policy innovations with a lag; therefore, $\mathbf{\Lambda_{YP}}$ is constrained to zero. For simplicity and without significant change to results, however, only the results from the three-variable SVAR are considered here.

[7] All data are from the Saint Louis Fed FRED II database (http://research.stlouisfed.org/fred2/). CPI comes from the CPIAUCSL series, FFR from FEDFUNDS series, IP from INDPRO series, CU from MCUMFN series, short rate from TB3MS series, and long rate from GS10 series. Quarterly data are averages of the monthly data. Earlier work on this paper indicated that if the quarterly time series were to be constructed by removing two monthly observations out of three in each quarter, the results would be quantitatively and qualitatively similar.

[8] The sample in Giordani (2004) is longer and earlier work on this paper that covered his sample indicates that the price puzzle maybe slightly smaller in the CU data than in the IP data, with the initial response of about 0.9% for the data starting in 1972:1. The impulse response is still significantly positive, however, hence the price puzzle is not eliminated in the monthly series even when capacity utilization is used as a measure of the output gap.

to capacity utilization as a measure of output resolves the price puzzle does not hold. The question that the next section attempts to answer is whether empirical estimates of the cost channel's presence using a specification of the New Keynesian Phillips curve can account for these results.

4 Estimates of the Cost Channel's Extent

As was detailed in Section 2, the cost channel of transmission of monetary policy is a theoretical mechanism that gives rise to the price puzzle. If firms need to finance their labor expenditures with credit, the nominal interest rate will enter the marginal cost function, giving rise to an additional term in the New Keynesian Phillips curve. It becomes possible then for a positive response of inflation to a contractionary monetary policy shock. In particular, as in Ravenna and Walsh (2006), given the standard production function:

$$Y_t = Z_t N_T^{\nu}, \tag{5}$$

where Y_t is output, Z_t is the technological factor, N_t is the labor input, and ν is labor's share of output, the firm's marginal cost function becomes:

$$MC_t = \frac{I_t^{\chi} W_t^r}{MPN_t}, \tag{6}$$

where I_t is the gross interest rate paid to secure credit for financing expenses, W_t^r is the real wage rate paid to the labor input, and MPN_t is the marginal product of labor. The parameter $\chi \geq 0$ measures the extent of cost channel's presence and can be interpreted as a combination of the share of firms that are exposed to it and the premium over the Federal Funds rate they need to pay to obtain credit. Chowdhurry, Hoffman, and Schabert (2005) also provide basis for interpreting it as the cost of financial market imperfections.

The general form of the Phillips curve that allows for both forward- and backward-looking dynamics can be posited as:

$$\pi_t = \frac{\beta}{1+\beta\omega} E_t \pi_{t+1} + \frac{\omega}{1+\beta\omega} \pi_{t-1} + \frac{\kappa}{1+\omega\beta} mc_t, \tag{7}$$

where mc_t is the log-deviation of the firm's real marginal cost from their deterministic steady state value. This specification can be formally derived by assuming the standard Calvo pricing mechanism where a share of firms θ cannot reset prices in a given time period. Endogenous persistence in inflation is introduced as in, for instance, Giannoni and Woodford (2004) and Rabanal and Rubio-Ramírez (2005), using the indexation mechanism, so that a fraction ω of firms that are not able to reset their prices, index them using last period's inflation. Sbordone (2002) has shown that the slope of the Phillips Curve is $\kappa = \frac{(1-\theta\beta)(1-\theta)}{\theta} \frac{\nu}{1+(1-\nu)\epsilon}$, where ϵ is the elasticity of substitution between differentiated goods that make up the composite consumption product.

Galí and Gertler (1999), Galí, Gertler, and Lópoez-Salido (2001, 2005) discuss estimates of this specification of the Phillips curve and find that the forward-looking component is quantitatively more important than the backward-looking one, although the latter typically is statistically significant. Their estimates of θ vary depending on the specification and typically fall within the 0.5-0.85 range. They emphasize that reliable estimates that fit the standard theory can be obtained with a measure of marginal (unit labor) costs but not with a measure of the output, which is what theoretical models routinely use. The coefficient on the output gap term they report is negative, which contradicts the standard theory. One contribution that the present paper makes is producing estimates of the Phillips curve that rely on measures of the output gap with stickiness parameter estimates that can fit the standard theory using both quarterly and monthly data.

4.1 A Reduced-form Model

Persistence in output is a well-documented empirical phenomenon that has several theoretical explanations. Introducing this form of a real rigidity implies that the firm's marginal cost function relates not only to a concurrent measure of the output gap but also to its lag and, possibly, to the current expectation of its future value. Therefore, as a starting point, I estimate the reduced-form version of the Phillips curve as:

$$\pi_t = \hat{a}E_t\pi_{t+1} + \hat{b}\pi_{t-1} + \hat{c}x_t + \hat{d}x_{t-1} + \hat{e}E_t x_{t+1} + \hat{f}i_t, \tag{8}$$

where x_t is the log output gap term. Following Chowdhury, Hoffman, and Schabert (2005), equation (8) is estimated together with an interest-rate monetary policy rule to account for the interplay between the Phillips curve and the monetary policy rule. I choose the standard specification from Clarida, Galí, and Gertler (1999), viz.:

$$i_t = (1-\rho)(\gamma_\pi E_t\pi_{t+1} + \gamma_x x_t) + \rho i_{t-1}, \tag{9}$$

where the lagged interest rate term adjusts for the observed persistence in the nominal interest rate targeted by the Federal Reserve.[9] Note that unlike Chowdhury, Hoffman, and Schabert (2005), this specification does include a lagged interest rate term and hence is not subject to the critique of Castelnuovo (2006) who shows that highly persistent interest rate rule dampen the effect of the cost channel. The system of equations (8) and (9) is estimated using the generalized method of moments with a set of instruments that is standard in the literature: lags of the Federal Funds rate, measure of the output gap, inflation, and the spread between interest rates on long- and short-term Treasury bonds. Four lags are used in quarterly estimation and twelve in monthly.

[Insert Table 1 about here]

Table 1 presents the reduced-form results. The coefficients on the lagged and led inflation terms, as well as the interest rate term, are similar to those reported in Chowdhury, Hoffman, and Schabert (2005). The output gap terms cannot be compared to this study because it uses a measure of unit

[9] In this equation, the following restrictions are motivated by theory (and common sense): $0 < \rho < 1$, γ_π, $\gamma_x > 0$. To produce stability in a system of equations with (8), (9), and an aggregate demand equation, γ_π should be greater than 1. See Clarida, Galí and Gertler (1999) for a discussion and estimates of this parameter for the U.S. data. The results reported below assume that the Fed targets inflation one quarter ahead in both monthly and quarterly results; hence the expected inflation term in one period ahead in the quarterly data and three periods ahead in the monthly data. Estimation results are not sensitive to the choice of this horizon in the monthly data. They are also not sensitive to the extending the lag on the right-hand-side nominal interest rate term, which is set to 2 to reflect the frequency of the Board meetings.

labor costs instead of the three output gap terms. All of the coefficients are significant at the 1% level in all models. The estimates of the "cost channel" coefficient are significant at the 1% level but quantitatively small, even if slightly larger than those reported in Chowdhury, Hoffman, and Schabert (2005). Note, however, that $\hat{f}$ captures the interaction between χ and κ adjusted for the degree of inflation persistence, hence when κ adjusted for inflationary persistence is small, χ can be large. Since κ decreases in θ and ω can be expected to be larger in the monthly data, the extent of the cost channel implied by this reduced-form coefficient does not necessarily have to be smaller. Finally, estimation results for the monetary policy rules indicate that all coefficients are significant at the 1% level and that γ_π is significantly larger than 1. These estimates are in lines with those available in existing literature.

4.2 Structural Explanation of the Reduced-form Estimates

The reduced-form estimates of (8) suggest that a structural model that would explain inflationary dynamics would need to incorporate forward- and backward-looking persistence in both output gap and inflation. The presence of the forward-looking output gap term makes it hard to use some devices for modeling output persistence, such as the real wage rigidity due to Blanchard and Galí (2005), Steinsson's (2003) rule-of-thumb pricing behavior, or external habit formation in consumption.

A popular explanation for the relationship between a firm's marginal cost and the backward- and forward-looking persistence in output is the assumption of internal habit formation in either consumption or labor supply. For instance, Amato and Laubach (2004) show that internal habit formation allows to relate marginal cost to output gap through the definition of the real wage that emerges from the household's first order condition with respect to labor supply in the following fashion:

$$mc_t = \zeta x_t - \xi \left(x_{t-1} + \beta E_t x_{t+1}\right) + \chi i_t, \tag{10}$$

where $\zeta = \frac{(\sigma-1)(1+\beta h^2)}{1-\beta h} + \frac{1+\eta}{\nu}$ and $\xi = \frac{h(\sigma-1)}{1-\beta h}$.[10] Note that this specification has been adjusted to allow for decreasing returns to labor as in equation (5) and that the presence of forward- and backward-looking output gap terms is motivated by the fact that households derive utility from some growth in consumption rather than its levels. This feature allows to model persistence in output observed in the data. Note that the output gap terms appear with the same sign as on the reduced-form estimates in Table 1. Before the results in this table are rendered structural interpretation, however, a discussion of the preference parameters that feature in this specification of the Phillips curve is in order.

4.2.1 Issues with Aggregate Demand

The aggregate demand relation derived in Amato and Laubach (2004) can be rewritten as:

$$x_t = \delta^{-1} x_{t-1} + \delta^{+2} E_t x_{t+2} + \delta^{+1} E_t x_{t+1} - \phi(i_t - E_t \pi_{t+1} - r_t^n), \tag{11}$$

where $\delta^{-1} = \frac{h(\sigma-1)}{(\sigma-1)(1+h+\beta h^2)+1-\beta h}$, $\delta^{+2} = -\frac{\beta h(\sigma-1)}{(\sigma-1)(1+h+\beta h^2)+1-\beta h}$, $\delta^{+1} = \frac{(\sigma-1)(1+\beta h+\beta h^2)+1-\beta h}{(\sigma-1)(1+h+\beta h^2)+1-\beta h}$, and the sensitivity of the current output gap to the real interest rate is $\phi = \frac{1-\beta h}{(\sigma-1)(1+h+\beta h^2)+1-\beta h}$. Problems with using GMM to estimate such an equation have been known at least since the work of Fuhrer and Rudebusch (2004) who, in particular, report the wrong sign on the real interest rate for a variety of specifications. For all empirical specifications that can be considered using the data described in Section 3, the estimate of this coefficient is very small in absolute value (typically less than 0.01), statistically insignificant, and, depending on the exact specification, frequently negative. Therefore, since the main focus of the my investigation is on the Phillips curve, I choose to impose the preference parameters exogenously, using parameter values that are consistent with the ones reported in existing literature.

[10] They assume that the representative household's utility function is

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[\frac{1}{1-\sigma} \left(\frac{C_t}{C_{t-1}^h} \right)^{1-\sigma} - \frac{1}{1+\nu} N_t^{1+\nu} \right],$$

. Hence the labor-leisure tradeoff that emerges from the first-order conditions implies that the real wage is related to labor supply and led, lagged, and concurrent consumption terms.

[Insert Table 2 about here]

For instance, Cho and Moreno (2005) use the full-information maximum likelihood method to estimate a small-scale New Keynesian model that is similar to the one used in the present paper using quarterly data and report that the interest rate sensitivity of output gap is small quantitatively at about 0.005 and not significantly different from zero.[11] I calibrate the parameters for the quarterly case so that this parameter matches their estimates. The choice of β at 0.99 and 0.997 for the quarterly and monthly cases, respectively, is standard. I assume that the intertemporal elasticity of substitution is three times higher in the monthly case and set σ at 4 and 1.33 for the quarterly and monthly calibration, since this parameter is the inverse intertemporal elasticity of substitution. To obtain the requisite quarterly value of ϕ, h should be equal to 0.97. This is close to the value used by Giannoni and Woodford (2004) who set it equal to 1 and to the estimates of Dennis (2005) that lie in the 0.85-0.90 range. I assume that for the reference points that are closer in time to the present h should be higher and set it at 1 for the monthly case. The resulting value of ϕ is lower than in the quarterly case, which is plausible and, as will be discussed below, sets the parameterization in favor of the cost channel appearing more strongly in the monthly case. Table 2 summarizes this parameterization. Note that the two parameters that characterize the production process are also listed there, as they will be needed to back out the degree of Calvo price stickiness θ: $\nu = 2/3$ and $\epsilon = 11$. The former value implies that the share of labor in output is two thirds and the latter that the steady state markup of a competitively monopolistic firm is 10%.

4.2.2 Estimates of the Cost Channel and Nominal Rigidities

Existence of endogenous habit formation in consumption, allows to relate the real wage, that enters a firm's marginal cost function, to a sequence of output gap terms: lag, concurrent, and lead. The

[11] It is possible to find specifications of aggregate demand that, in the absence of the cost channel as discussed in Section 4.2.3 produce similar estimates of this parameter. This result, however, is sensitive to specification and hence unreported.

Phillips curve then becomes:

$$\pi_t = \beta\tilde{\omega}E_t\pi_{t+1} + \beta^{-1}(1-\tilde{\omega})\pi_{t-1} + \kappa\tilde{\omega}\zeta x_t - \kappa\tilde{\omega}\xi\left\{x_{t-1} + \beta E_t x_{t+1}\right\} + \kappa\tilde{\omega}\chi i_t, \tag{12}$$

where $\tilde{\omega} = \frac{1}{1+\beta\omega}$. This equation is estimated jointly with (9). Table 3 reports the results. Products of parameters with 'widehats' over them are estimated as one coefficient and then disentangled given other estimates and imposed parameters.

[Insert Table 3 about here]

The parameterization outlined in the previous subsection implies the following biases: (a) lower interest sensitivity of output gap in the monthly data should make it easier to generate the price puzzle at that frequency; (b) higher value of h dominates the effect of a lower value of σ, resulting in a higher value of ξ at the monthly frequency, which, given a certain estimate of $\widehat{\kappa\tilde{\omega}\xi}$, should produce a lower value of κ and hence imply stickier prices in the monthly data; and (c), given lower κ and an estimate of $\widehat{\kappa\tilde{\omega}\chi}$, the implied value of χ that describes the extent of the cost channel should be higher. Hence the price puzzle should be generated more easily at the monthly frequency. Even so, Table 3 reveals that these estimation and parameterization produce *lower* estimates of the cost channel in the monthly data for both IP and CU samples. In all cases, the implied value of χ is too low to generate the price puzzle. Note that these estimates of χ are just slightly higher than Rabanal's (2003) of 0.24, lower than those of Chowdhury, Hoffman, and Schabert (2005) who obtain values between 0.8 and 1.5 or those of Ravenna and Walsh (2006) who also present quarterly estimates that, depending on the instrument set, range from 1.2 to 11.8. Therefore, neither in the monthly nor the quarterly case can the cost channel explain the price puzzle. Furthermore, the cost channel cannot account for the fact that the price puzzle appears more strongly in the monthly rather than quarterly data.

The results in Table 3 point to several other interesting issues. First, although inflationary persistence shows up more strongly in the monthly case for the IP model, it is about the same for the CU model. Second, the implied values for both κ and θ are about the same in all monthly and quarterly cases. They suggest that the average duration of a pricing contract is about 3.5 time periods, which means that prices appear to be much more flexible in the monthly as opposed to quarterly case. The possibility that quarterly estimates overstate the true extent of price stickiness should be further investigated in future research. Finally, a casual comparison of the coefficients on output gap and interest rate terms in Tables 1 and 3 reveals that the former have not changed much in absolute value or statistical significance, whereas the latter is much smaller in value and loses its significance. Notice also that the specification of the Phillips curve (12) has enough moments to estimate the value of β as well, although imposing it exogenously is a common strategy in the literature pursued, for instance, by Giannoni and Woodford (2003) and Rabanal (2003). The last part of this section briefly investigates the implications of this parameterization.

4.2.3 The Relationship Between β and the Cost Channel: Another Puzzle?

Hülsewig, Mayer and Wollmerhäuser (2006) point out that there are no precise guidelines established in the literature for which parameters to estimate and which to impose exogenously. At the same time, implausibly low estimated values of β are fairly common.[12] For instance, Galí and Gertler (1999) produce estimates of about 0.77 for the most recent dataset considered in their study for a Phillips curve with persistent inflation and unit labor costs as a measure of marginal costs. Their results indicate that removing the lagged inflation term from the specification or extending the sample farther into the past can increase the value of these estimates closer to the steady state value. In a purely forward-looking model and with a much longer sample that does not allow to control for stability of monetary policy conduct, Ravenna and Walsh (2006) report estimates of β that range from 0.85 to 0.97. Interestingly, when they account for the cost channel—and this result is robust for different specifications—they obtain *lower* estimates of β then when they omit the cost channel's presence. Their highest estimate of χ is produced in the specification where accounting for the cost channel produces the largest decline in the estimate of β.

[Insert Table 4 about here]

[12] All of the estimates in alí and Gertler (1999) and Ravenna and Walsh (2006) are obtained using quarterly data.

Table 4 presents regression results where β is also estimated. Again, the coefficients on the output gap terms are about the same as before, whereas those on the interest term are higher in absolute value and significant at the 1% level, except in the quarterly IP case. The estimated values of β are implausibly low; interestingly, they seem to be inversely related to the interest rate coefficients. Trying to tease out the structural estimates is problematic because the low β values imply either inordinately large sensitivity of output gap to interest rate and/or implausibly high values of σ, as well as dramatically low degrees of price stickiness. It is possible to manipulate the preference parameters to obtain implied values of χ that range between 1 and 2, although they are still not sufficient to generate the price puzzle and, again, the quarterly estimates continue to be higher than the monthly ones.

[Insert Table 5 about here]

Finally, Table 5 reports results where χ is restricted to zero and there is no cost channel. The β estimates are much higher than in Table 4 and the null hypotheses that they equal their steady state values cannot be rejected. Under this null, the structural parameters that describe nominal frictions take on roughly the same values as in Table 3. This exercise illustrates an interesting "tradeoff" between the cost channel and the discount factor supported by the evidence in Ravenna and Walsh (2006): The presence of the nominal interest rate term in the Phillips curve reduces the estimate of β, whereas fixing the latter at its steady state value reduces the estimated significance of the cost channel.

5 Conclusion

This paper has reviewed evidence from the monthly and quarterly U.S. data on inflationary dynamics, the price puzzle, and the cost channel of transmission of monetary policy. It has demonstrated that measures of output gap can be used to obtain sensible estimates of the New Keynesian Phillips curve using both quarterly and monthly data. Evidence that relies on two different measures of output gap points that the price puzzle exists in the U.S. data at both frequencies. A structural model employed here indicates that the presence of the cost channel may be significant at both monthly and quarterly frequencies but only in the case where the discount factor β is freely estimated. However, including a nominal interest rate term in the Phillips curve to describe the cost channel leads to unreasonably low estimates of the discount factor. On the other hand, once the interest rate term is dropped from the specification very reasonable estimates of β can be obtained at both frequencies. Finally, regardless of the specification, when the cost channel is assumed to exist its estimated extent is not large enough to produce the price puzzle.

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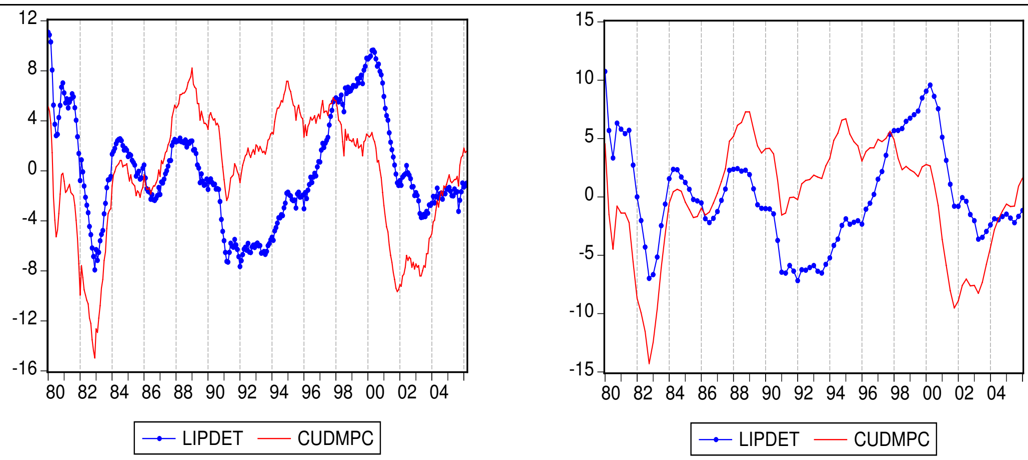

(a) Monthly (b) Quarterly

Figure 1: Measures of Output Gap in the Monthly and Quarterly Data

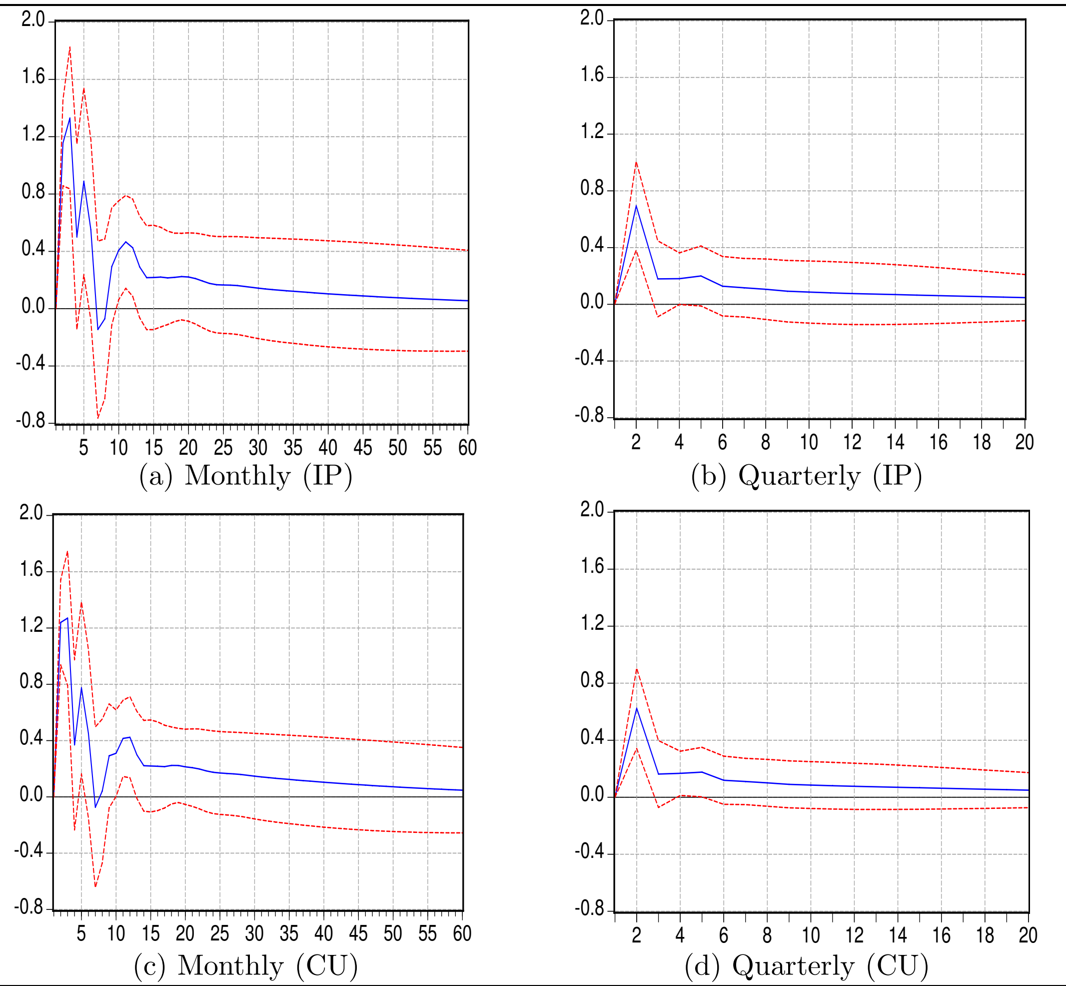

Figure 2: Empirical impulse responses of inflation to a 1% FFR shock

Table 1: Reduced-form Parameter Estimates (standard errors in parentheses)

	IP Model		CU Model	
Estimated Parameters	Monthly	Quarterly	Monthly	Quarterly
$\hat{a}$	0.4951	0.5569	0.5547	0.4685
	(0.0466)	(0.0821)	(0.0560)	(0.0708)
$\hat{b}$	0.3568	0.1851	0.3095	0.2470
	(0.0277)	(0.0661)	(0.0312)	(0.0516)
$\hat{c}$	2.0859	1.9190	2.1317	1.3474
	(0.2397)	(0.2614)	(0.2357)	(0.1423)
$\hat{d}$	-1.2891	-0.9394	-1.2784	-0.7014
	(0.1479)	(0.1241)	(0.1390)	(0.0746)
$\hat{e}$	-0.8025	-1.0472	-0.8517	-0.6721
	(0.1326)	(0.1640)	(0.1355)	(0.0865)
$\hat{f}$	0.0778	0.1250	0.0736	0.1448
	(0.0213)	(0.0308)	(0.0242)	(0.0310)
Interest Rate Rule				
$\hat{\rho}$	0.8740	0.8497	0.8940	0.8768
	(0.0074)	(0.0170)	(0.0072)	(0.0138)
$\hat{\gamma_\pi}$	1.7508	1.7867	1.6693	1.6492
	(0.0493)	(0.0980)	(.0586)	(0.1056)
$\hat{\gamma_x}$	0.2756	0.2899	0.3323	0.4017
	(0.0365)	(0.0662)	(0.0504)	(0.0792)

All coefficients are significant at the 1% level

Table 2: Imposed Structural Parameters

	Monthly	Quarterly
	Preference	
β	0.997	0.99
σ	1.33	4
h	1	0.97
ϕ	0.003	0.005
	Production	
ν	2/3	2/3
ϵ	11	11

Table 3: Structural Parameter Estimates (s.e.)

	IP Model		CU Model	
Estimated Parameters	Monthly	Quarterly	Monthly	Quarterly
$\widehat{\tilde{\omega}}$	0.6302	0.7355	0.6633	0.6692
	(0.0264)	(0.0490)	(0.0328)	(0.0421)
$\widehat{\kappa\tilde{\omega}\zeta}$	2.0495	1.8894	2.1871	1.3267
	(0.2444)	(0.2535)	(0.2361)	(0.1586)
$\widehat{\kappa\tilde{\omega}\xi}$	1.0291	0.9710	1.0975	0.6815
	(0.1224)	(0.1316)	(0.1183)	(0.0821)
$\widehat{\kappa\tilde{\omega}\chi}$	0.0029*	0.0059*	0.0037*	0.0072*
	(0.0048)	(0.0094)	(0.0051)	(0.0079)
Implied Parameters				
ω	0.5886	0.3632	0.5091	0.4994
κ	0.0147	0.0180	0.0149	0.0139
θ	0.7231	0.6913	0.7216	0.7214
χ	0.3136	0.4485	0.3715	0.7755

All coefficients are significant at the 1% level except marked by *, which is not significant at the 10% level.

Table 4: Structural Parameter Estimates with Free β (s.e.)

	IP Model		CU Model	
Estimated Parameters	Monthly	Quarterly	Monthly	Quarterly
$\widehat{\beta}$	0.6525	0.9140	0.6961	0.5310
	(0.0705)	(0.0916)	(0.0809)	(0.0795)
$\widehat{\tilde{\omega}}$	0.7699	0.7541	0.7835	0.8600
	(0.0278)	(0.0458)	(0.0290)	(0.0268)
$\widehat{\kappa\tilde{\omega}\zeta}$	2.0948	1.7484	2.1693	0.8241
	(0.2212)	(0.2745)	(0.2222)	(0.1318)
$\widehat{\kappa\tilde{\omega}\xi}$	1.2711	0.9336	1.2778	0.5426
	(0.1474)	(0.1301)	(0.1388)	(0.0883)
$\widehat{\kappa\tilde{\omega}\chi}$	0.0760	0.0244*	0.0769	0.1460
	(0.0199)	(0.0228)	(0.0230)	(0.0269)

All coefficients are significant at the 1% level except marked by *, which is not significant at the 10% level.

Table 5: Structural Parameter Estimates without Cost Channel (s.e.)

	IP Model		CU Model	
Estimated Parameters	Monthly	Quarterly	Monthly	Quarterly
$\widehat{\beta}$	0.9460	1.0019	0.9681	1.0086
	(0.0354)	(0.0382)	(0.0304)	(0.0451)
$\widehat{\tilde{\omega}}$	0.6423	0.7258	0.6749	0.6612
	(0.0231)	(0.0498)	(0.0295)	(0.0407)
$\widehat{\kappa\tilde{\omega}\zeta}$	2.0722	1.9159	2.1840	1.3456
	(0.2461)	(0.2587)	(0.2403)	(0.1521)
$\widehat{\kappa\tilde{\omega}\xi}$	1.0639	0.9785	1.1109	0.6849
	(0.1295)	(0.1327)	(0.1258)	(0.0823)

All coefficients are significant at the 1% level.