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In 2007 all ECB publications feature a motif taken from the €20 banknote.

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Abstract

Macroeconomic data suggest that the New Keynesian Phillips curve is quite flat - despite microeconomic evidence implying frequent price adjustments. While real rigidities may help to account for the conflicting evidence, we propose an alternative explanation: if price markup/cost-push shocks are persistent and negatively correlated with the labor share, the latter being a widely used measure for marginal costs, the estimated pass-through of measured marginal costs into inflation is limited, even if prices are fairly flexible. Using a standard New Keynesian model, we show that the GMM approach to the New Keynesian Phillips curve leads to inconsistent and upward biased estimates if cost-push shocks indeed are persistent. Monte Carlo experiments suggest that the bias is quite sizeable: we find average price durations estimated as high as 12 quarters, when the true value is about 2 quarters. Moreover, alternative estimators appear to be biased as well, while standard diagnostic tests fail to signal a misspecification of the model.

Keywords: Price Rigidities, New Keynesian Phillips Curve, Cost-push shocks, GMM estimation
JEL-Codes: E31, E32, C22

Non-technical summary

This paper suggests a new perspective on the conflicting micro- and macroeconomic evidence on price stickiness. Estimating the New Keynesian Phillips curve (NKPC), which is a widely used structural model of inflation dynamics, macroeconomic evidence suggests that prices do not respond much to changes in demand and supply conditions. Using Generalized Method of Moments (GMM) estimation techniques on U.S. time series Galí and Gertler (1999), for example, find average price durations of about a year. These estimates seem high, or, put differently, the NKPC appears to be quite flat, given recent microeconomic evidence presented by Bils and Klenow (2004) suggesting average price durations of less than 5 months.

Starting from the observation that recent estimates of New Keynesian general equilibrium models on the basis of full information estimation techniques find evidence for autocorrelated cost-push shocks, we assess the implications of these shocks for the single equation, GMM-based approach to the NKPC. With autocorrelated shocks the orthogonality conditions imposed in the GMM estimation are invalid, so GMM estimation is inconsistent. The question remaining is how severe the bias in the estimation actually is and, in particular, in which direction the bias goes. Towards answering this, we conduct Monte Carlo experiments using as data generating process the small-scale New Keynesian general equilibrium model estimated by Galí and Rabanal (2005), which is by and large representative for the recent literature. Our estimates of the degree of price stickiness are inconsistent (as theory suggests) and upward biased, i.e. they imply too much price rigidity, once the cost-push shocks are autocorrelated. We find average price durations estimated as high as 12 quarters, when the true value is about 2 quarters.

We also assess whether alternative approaches are more robust to this mis-specification and consider the two stage minimum distance (MD) estimator suggested by Sbordone (2002, 2005) as well as classical full-information maximum likelihood (FIML) techniques. In both cases, just as with GMM, there is substantial upward bias in the estimated degree of price rigidity whenever the serial correlation of the cost-push shock is neglected.

The economic mechanism that underlies the upward bias is as follows. Positive cost-push shocks induce a negative response of the labor share, the measure for marginal costs used in the literature. To the extent that cost-push shocks play a role in driving the business cycle, a high labor share will frequently come along with a lower than average realization of the cost push-shock and this will be persistently the case if cost-push shocks are autocorrelated. The partial inflationary effect of the labor share (wages) on inflation will thus be underestimated whenever the serial correlation in the markup shock is not accounted for.

1 Introduction

The New Keynesian Phillips curve (NKPC) is a widely used structural model of inflation dynamics. Its key parameter, which governs the pass-through of marginal costs into inflation, is the average time over which prices are kept fixed. This average price duration provides a measure for the degree of price stickiness. If prices are kept unchanged for some time, i.e. if they are sticky, then the pass-through of marginal costs into inflation is limited, or, in other words, the Phillips curve is quite flat. Using Generalized Method of Moments (GMM) estimation techniques on U.S. time series Galí and Gertler (1999), hereafter GG, find average price durations of about 4 quarters. These estimates seem high, or, put differently, the NKPC appears to be quite flat, given recent microeconomic evidence presented by Bils and Klenow (2004) suggesting average price durations of less than 5 months.[1]
Substantial research efforts have been made to reconcile the apparent conflict between micro and macroeconomic evidence on price stickiness. By now it is widely recognized that various kinds of real rigidities allow a reinterpretation of the macroeconomic evidence, making it consistent with much shorter average price durations, see, for instance, Eichenbaum and Fisher (2004) and Altig, Christiano, Eichenbaum, and Lindé (2005). Intuitively, real rigidities induce strategic complementarities in price setting such that price setters change prices by smaller amounts whenever they are able to adjust prices. As a consequence, the pass-through of marginal costs into inflation remains limited, although firms update their prices frequently. The appeal of this route to reconcile micro and macroeconomic evidence remains limited, however, to the extent that microevidence not only suggests frequent, but also large price adjustments, a point stressed, for instance, by Maćkowiak and Wiederholt (2006).
The present paper, therefore, suggests a new perspective on the conflicting evidence on price stickiness. We start from the observation that recent estimates of New Keynesian general equilibrium models on the basis of full information estimation techniques find evidence for autocorrelated cost-push shocks, also labeled ‘price markup shocks’ in the literature; see e.g. Galí and Rabanal (2005) and Smets and Wouters (2006). In the present paper, we assess the implications of these shocks for the single equation, GMM-based approach to the NKPC. Cost-push shocks in general equilibrium are correlated with the labor share, the observable measure typically used for marginal costs. Once these shocks are autocorrelated therefore the orthogonality conditions imposed in the GMM estimation are invalid. As a consequence, we find that GMM estimates of the degree of price stickiness are inconsistent and upward biased, i.e. they imply too much price rigidity.

[1] Galí, Gertler, and López-Salido (2001, 2003, 2005) provide additional evidence. Depending on the particular specification and sample, the average price duration is found to be between 2.4 and 11.8 quarters for U.S. data. Similar results are reported in Sbordone (2002, 2005). Note that average price durations of about 4 quarters seemed to square well with earlier survey evidence, see Blinder, Canetti, Lebow, and Rudd (1998). Also, a more recent assessment of microeconomic data for the U.S. by Nakamura and Steinsson (2006) suggests price durations could be between 8 and 11 month once sales are excluded from the sample.

Clearly, while any omission of actual features of the economy may induce a bias in the estimates, eventually it remains an empirical question how severe these omissions are. To answer this question in the context of the NKPC, we conduct Monte Carlo experiments using as data generating process the small-scale New Keynesian general equilibrium model estimated by Galí and Rabanal (2005). This model stays close in flavor to the New Keynesian variants discussed in Woodford (2003) and is by and large representative of a sizeable literature of small to medium scale estimated DSGE models, such as those of Christiano, Eichenbaum, and Evans (2005) and Smets and Wouters (2006). Moreover, it has been shown to provide a satisfactory account of U.S. time series and thus appears to be an empirically plausible model well suited to take up the issue of misspecification in the single equation approach to the NKPC through Monte Carlo experiments.

Using the simulated data, we find that GMM estimates are precise and consistent as long as the model is correctly specified. However, once we introduce autocorrelated cost-push shocks into the model, the performance of the GMM estimator deteriorates. The procedure signals a precise estimate even though there is substantial bias. In particular, we find that the bias increases drastically as the persistence of cost-push shocks increases. In fact, for a degree of autocorrelation of cost-push shocks of 0.95, the value reported in Galí and Rabanal (2005), the median estimate for the degree of price stickiness implies a mean duration of prices of about 12 quarters although the true value is just over 2 quarters.

We then turn to diagnostic tests asking whether the econometrician would be able to detect the misspecification of the model. We find that this is unlikely given a realistic sample size of about 150 observations. While the orthogonality conditions imposed in the estimation are, in fact, violated by the model, we find that the power of the J-test is too low to reject the null of no violation. Regarding the error terms, there is some evidence for autocorrelation. Yet, it hardly exceeds the extent of autocorrelation in the residuals suggested by tests on U.S. data, say, by Galí et al. (2001). In any case, we illustrate that the widely used Q-Test is a measure of the mere presence of cost-push shocks as much as of the serial correlation properties of those. A more sophisticated test, suggested by GG, gives a mixed picture.

We also assess whether alternative approaches are more robust to misspecification and consider the two stage minimum distance (MD) estimator suggested by Sbordone (2002, 2005) as well as classical full-information maximum likelihood (FIML) techniques. In both cases, just as with GMM, there is substantial upward bias in the estimated degree of price rigidity. In the same vein, casual observations of recent full information estimation results of the NKPC also suggest lower degrees of price rigidities, in case autocorrelated cost-push shocks are allowed for, see Galí and Rabanal (2005) and Smets and Wouters (2006).

The following economic mechanism underlies our results. In the New Keynesian model, which is the data-generating process in the paper, pricing decisions depend not only on current and future marginal costs but also on exogenous cost-push shocks (or equivalently, on exogenous fluctuations in the price markup). Negative realizations of these shocks, i.e. a lower markup, induce a rise in the labor share, the measure widely used as a proxy for marginal costs. A high labor share, for example, may thus partly reflect a fall in markups which by itself, all else equal, would be dis-inflationary. While high labor shares are inflationary in the model keeping all else equal, from an unconditional perspective high labor shares in the model therefore are not always inflationary. If cost-push/markup shocks are not accounted for in the estimation, this limits the estimated average effect of the labor share on inflation. Since the labor share is widely used as a proxy for the full marginal costs, the estimated pass-through of marginal costs to inflation will be too small. Put differently, the NKPC appears to be flat, while in fact the pass-through of marginal costs into inflation is generally quite high.

The remainder of the paper is organized as follows. Section 2 outlines the Galí and Rabanal (2005) model used as the data generating process. Section 3 shows that the estimated degree of price rigidity is likely to be upward biased in the presence of autocorrelated cost-push shocks and examines asymptotics of frequently used diagnostic tests. Using Monte Carlo experiments, Section 4 provides a quantitative assessment of this bias considering, in turn, GMM, MD and ML techniques. Concluding remarks follow in Section 5.

2 The Model Economy

In this section we outline a variant of the baseline New Keynesian general equilibrium model. Specifically, we focus on the model estimated by Galí and Rabanal (2005) on post World War II U.S. data. Our exposition of the model economy is brief; for more details on the model and the estimation, we refer the interested reader directly to Galí and Rabanal (2005).

The demand side of the model is represented by a consumption Euler equation:

$$b\Delta y_t = E\left\{\Delta y_{t+1}\right\} - (1-b)\left(r_t - E_t\left\{\pi_{t+1}\right\}\right) + (1-\rho_g)(1-b)g_t. \tag{1}$$

Above, Δy_t denotes output growth, r_t the nominal interest rate, π_t the inflation rate. g_t is a demand shock. Parameters ρ_g and b reflect the autocorrelation of this shock and external consumption habits, respectively.

Firms produce differentiated goods which they sell in monopolistically competitive product markets. Their output is linear in employment. From the production side, up to first order, aggregate output per efficiency unit, $\tilde{y}_t$, is therefore linear in aggregate employment, n_t:

$$\tilde{y}_t = n_t. \tag{2}$$

Productivity shocks in the economy are permanent shocks. A tilde on top of a variable indicates that the respective variable has been normalized by productivity before linearization in order to render the linearized model economy stationary. $\tilde{y}_t$ is linked to Δy_t, and to innovations in permanent technology, ϵ_t^a, by $\Delta y_t = \Delta \tilde{y}_t + \epsilon_t^a$.

Producers of differentiated goods hire bundles of labor in a perfectly competitive market. In each period, which is assumed to be one quarter, each producer of a differentiated good will not be able to reoptimize its price with a certain probability. Producers who do not reoptimize mechanistically index to lagged inflation instead. Linearizing around a zero inflation steady state, the NKPC in this model is given by

$$\pi_t = \gamma_b \pi_{t-1} + \gamma_f E_t \{\pi_{t+1}\} + \kappa_p (l_t + u_t). \tag{3}$$

Above $\gamma_b = \frac{\eta_p}{1+\beta\eta_p}$, $\gamma_f = \frac{\beta}{1+\beta\eta_p}$ and $\kappa_p = \frac{(1-\beta\theta_p)(1-\theta_p)}{\theta_p(1+\eta_p\beta)}$. θ_p is the probability that a firm cannot reoptimize its price in a given period. Below we will be concerned with the estimation of this parameter which, by the law of large numbers, can also be interpreted as the fraction of firms which keep their prices unchanged in a given period. The average price duration, D, is given by $1/(1-\theta_p)$.

The model also allows for price indexation, captured by parameter η_p. Parameter β denotes the time discount factor. According to the model, inflation is driven by current and expected real marginal costs. These can be decomposed into a measure of the labor share, $l_t = \tilde{w}_t$, where $\tilde{w}_t$ denotes the real wage per efficiency unit, and a shock to the markup/a cost-push shock, u_t. It is the autocorrelation of this shock, ρ_u, which crucially influences estimates of the degree of price stickiness in the economy as we will demonstrate below.

Workers supply their labor monopolistically competitive to intermediaries who bundle these labor services and sell them on to goods producers under perfect competition. Like goods prices, also individual wages are subject to a Calvo nominal rigidity. Those workers who do not update their wage in a given period instead partially index their nominal wage to past inflation. The wage equation in the model of Galí and Rabanal (2005) reads as

$$\begin{aligned}\tilde{w}_t \quad = \quad & \tfrac{1}{1+\beta}\tilde{w}_{t-1} + \tfrac{\beta}{1+\beta}E_t\{\tilde{w}_{t+1}\} - \tfrac{1}{1+\beta}\epsilon_t^a + \tfrac{\eta_w}{1+\beta}\pi_{t-1} - \tfrac{1+\beta\eta_w}{1+\beta}\pi_t \\ & + \tfrac{\beta}{1+\beta}E_t\{\pi_{t+1}\} - \tfrac{\kappa_w}{1+\beta}(\mu_t^w - \nu_t).\end{aligned} \tag{4}$$

Wages are driven by endogenous variations in the wage markup, μ_t^w, and by exogenous shocks to the markup, ν_t. Parameter $\kappa_w = \frac{(1-\theta_w)(1-\beta\theta_w)}{\theta_w(1+\epsilon_w\varphi)}$ multiplying the markup term measures the strength with which markups influence wages. θ_w is the probability that a worker cannot change its wage, ϵ_w is the elasticity of labor demand with respect to wages and $1/\varphi$ measures the Frisch elasticity of labor supply. The markup links to the rest of the economy via $\mu_t^w = \tilde{w}_t - \left(\frac{1}{1-b}\tilde{y}_t - \frac{b}{1-b}\tilde{y}_{t-1} - g_t + \frac{b}{1-b}\epsilon_t^a + \varphi n_t\right)$. Like price markup shocks, also wage markup shocks are serially correlated, as captured by the autocorrelation parameter, ρ_ν. The economy is closed by a Taylor type rule for monetary policy. The authority sets interest rates in reaction to inflation and output growth. On top of this monetary policy rates are also subject to a monetary policy shock ϵ_t^m:

$$r_t = \phi_r r_{t-1} + (1-\phi_r)\phi_\pi \pi_t + (1-\phi_r)\phi_y \Delta y_t + \epsilon_t^m. \tag{5}$$

Table 1: Parameter values used in simulation

Consumers		
Time-discount factor	β	= 0.99
External habit parameter	b	= 0.42
Inverse of labor supply elasticity	φ	= 0.80
Calvo-stickiness wages	θ_w	= 0.05
Indexation wages	η_w	= 0.42
Price elasticity of demand	ϵ_p	= 6.00
Producers		
Calvo-stickiness prices	θ_p	= 0.53
Indexation prices	η_p	= 0.00
Wage elasticity of labor demand	ϵ_w	= 6.00
Monetary policy		
Response to lagged interest rate	ϕ_r	= 0.69
Response to inflation	ϕ_π	= 1.35
Response to output growth	$\phi_{\Delta y}$	= 0.26
Autocorrelation of shocks		
Persistence of demand shock	ρ_g	= 0.93
Persistence of productivity	ρ_a	= 1.00
Persistence of price-markup shock	ρ_u	= 0.95
Persistence of wage-markup shock	ρ_ν	= 0.91
Standard deviation of innovations		
Innovation to productivity	σ^a	= 0.009
Innovation to demand shock	σ^g	= 0.025
Innovation to price-markup	σ^u	= 0.011
Innovation to wage-markup	σ^w	= 0.012
Innovation to monetary policy	σ^m	= 0.003

Notes: Parameters used to generate data. Parameters are taken from the mean estimates of Galí and Rabanal (2005). Their mean estimate of price indexation is $\eta_p = 0.02$. We set this parameter to zero in order to enhance expositional clarity.

Table 1 shows the parameterization we choose for the model economy on the basis of the mean estimates reported by Galí and Rabanal (2005). We only deviate with respect to the indexation of prices where their mean estimate is as low as $\eta_p = 0.02$. Following large parts of the GMM literature on the NKPC, we set this parameter to zero.

3 Cost-push shocks and the econometrics of the NKPC

Having outlined a fully specified structural model, we now turn to the econometrics of the NKPC, given by equation (3). This equation takes center stage within the single equation approach based on the Generalized Methods of Moments (GMM). We highlight shortcomings of this approach and how these will affect the economic conclusions about the inferred degree of price rigidity. Intuitively, we expect that the estimated pass-through of the labor share, l_t, which is taken as a measure of real marginal costs into inflation, π_t, is lowered by cost-push shocks if they are negatively correlated with the labor share. The pass-through is captured by the estimate of the slope of the NKPC, $\widehat{\kappa}_p$. Indeed, this is our first point: whenever cost-push shocks are serially correlated, as much of the recent Bayesian estimation literature finds, GMM estimates of price-duration are inconsistent. The serial correlation of cost-push shocks renders the orthogonality conditions exploited in the GMM estimation invalid. In the model at hand we compute, second, that the bias runs in the direction of too much estimated price rigidity. This, in principle, should be detected by the J-Test, but may go unnoticed in realistic sample sizes as we illustrate below. In the following, for the sake of clarity, we restrict ourselves to the estimation of the slope of the Phillips curve, κ_p, and restrict the other parameter in the NKPC, β, to its true value.[2]

Consistency To be slightly more formal note that, absent indexation, NKPC equation (3) can be written as follows

$$y_{t+1} = \kappa_p l_t + \epsilon_{t+1}^{RE} + \tilde{u}_t, \tag{6}$$

where $y_{t+1} := \pi_t - \beta\pi_{t+1}$.

Here ϵ_{t+1}^{RE} is the rational expectations error that ensures $\beta E_t(\pi_{t+1}) \equiv \beta\pi_{t+1} + \epsilon_{t+1}^{RE}$ and $\tilde{u}_t = \kappa_p u_t$. Equation (6) is linear. Estimating it by GMM is thus equivalent to applying two stage least squares estimation (2SLS henceforth).[3] Let the instrument vector be $\boldsymbol{z}_{t-1} = [\pi_{t-1}, l_{t-1}]'$. As Appendix A shows, in probability the estimator converges to

$$\widehat{\kappa}_{p,T}^{2SLS} \xrightarrow{p} \kappa_p + q^{-1}\overline{\boldsymbol{\delta}}'\kappa_p\rho_u E\left\{\left[\begin{array}{c} \pi_{t-1}, l_{t-1} \end{array}\right]' u_{t-1}\right\}. \tag{7}$$

[2] Appendix A presents the case where also β is estimated. This does not have a strong bearing on the bias in the estimated slope of the NKPC, although the upward bias in the implied price duration is somewhat mitigated by an upward bias also in the estimate of β.

[3] Rudd and Whelan (2005) also exploit this property. They assess whether GMM estimates of the NKPC might be biased if the true process is a backward-looking NKPC and the lags are omitted in the estimation.

Here $\overline{\boldsymbol{\delta}}$ is the probability limit of $\widehat{\boldsymbol{\delta}}_T$, the estimator in the first-stage regression, and q is the second cross-moment of regressor and instrumented regressor. Most notably, to the extent that the instruments correlate with the cost-push shock contemporaneously, as is the case in our data-generating model, i.e. $E\left\{\left[\begin{array}{c}\pi_{t-1}, l_{t-1}\end{array}\right]' u_{t-1}\right\} \neq 0$, the estimates will be inconsistent whenever the cost-push shock is persistent ($\rho_u \neq 0$).

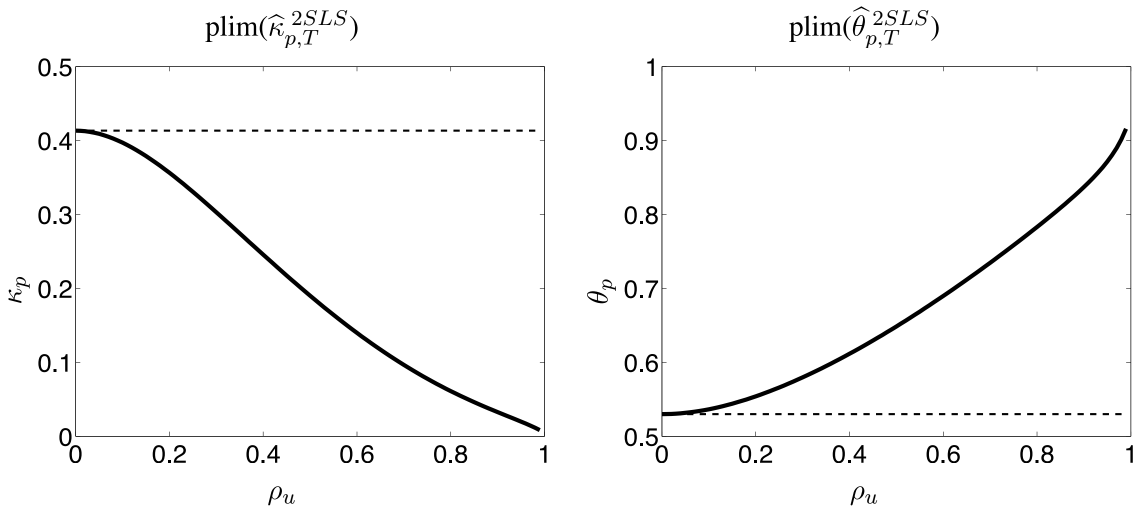

Figure 1: THEORETICAL PROBABILITY LIMITS OF 2SLS ESTIMATORS FOR VARYING DEGREES OF SERIAL CORRELATION, ρ_u, IN THE COST-PUSH SHOCK. *Notes:* Dashed lines show the true values. From left to right: estimator for the slope of the Phillips curve, κ_p, and for the fraction of firms which do not reoptimize their price, θ_p. In the graphs as ρ_u varies, the standard deviation of the cost-push shock, u_t, is left unchanged. We adjust the innovation variance to this shock accordingly.

Assessing the direction of the bias Under the calibration of the model, computing $q, \overline{\boldsymbol{\delta}}$ and the covariance between instruments and the cost-push term, we find that whenever cost-push shocks are persistent, estimates of the slope of the NKPC, κ_p, are inconsistent and converge to too low values. The left panel of Figure 1 plots the probability limit of $\widehat{\kappa}_p$ as a function of the serial correlation of cost-push shocks, ρ_u. We leave the standard deviation of the cost-push shock, u_t, unchanged and, consequently, adjust the standard deviation of the innovation to this shock when ρ_u changes. The downward bias resulting from an increase in ρ_u is apparent. It translates into too large estimates of the mean price duration even as the sample size grows, shown in the right panel of Figure 1.

The economics behind this finding is the following. Consider for the sake of the argument a higher than average realization of the labor share, l_t. According to the NKPC, ceteris paribus a high value of the labor share in period t has an inflationary impact. According to the NK model, all else equal again, a negative realization of the cost-push shock induces a rise in the labor share. To the extent that cost-push shocks are a sizeable and persistent source of fluctuations in the labor share, a persistently higher than average labor share will oftentimes come along with a low realization of the cost-push/markup shock. On these occasions a high labor share is not as inflationary as it would be in the absence of correlated cost-push shocks. In order to reconcile the frequent absence of an inflationary impact of the labor share with the NKPC, the pass-through parameter, κ_p, needs to be small – the slope of the Phillips curve will be underestimated.

Implications of serial correlation for the J-Test Assuming that cost-push shocks are unobserved, the GMM approach to the NKPC exploits the following set of moment conditions:

$$\begin{aligned} E\left\{[y_{t+1}-\kappa_p l_t]\,\boldsymbol{z}_{t-1}\right\} &= E\left\{\left[\epsilon_{t+1}^{RE}+\tilde{u}_t\right]\boldsymbol{z}_{t-1}\right\} \\ &= \rho_u E\left\{\tilde{u}_{t-1}\boldsymbol{z}_{t-1}\right\} \neq 0, \quad \text{if} \quad \rho_u > 0. \end{aligned} \tag{8}$$

In words, the GMM estimation is based on orthogonality conditions which will likely be violated whenever cost-push shocks are serially correlated. In principle, the J-test used in GMM estimation should signal the violation of the moment conditions; in practical applications, however, lack of power leads to a frequent failure to reject the null as we shall demonstrate in Section 4.

Implications of cost-push shocks for the Q-Test The Q-Test for the serial correlation of the error term is frequently used in empirical work to discern whether cost-push shocks are white noise. Evaluated at the true parameter values, the combined residual in the New Keynesian Phillips curve is given by $e_t := \epsilon_t^{\text{RE}} + \tilde{u}_{t-1}$. Evidently $\epsilon_{t-1}^{\text{RE}}$ in the model is not orthogonal to $\tilde{u}_{t-1}$. e_t is therefore serially correlated whenever cost-push shocks are present.[4] Evidence obtained on the basis of the Q-Test which signals serial correlation in the NKPC residual is therefore as much evidence of the mere presence of shocks to the markup/cost-push shocks as a test of the serial correlation properties of cost-push shocks. The Q-Test cannot be used to discern whether cost-push shocks/shocks to the markup are serially correlated as we shall illustrate in the next section.

4 Monte Carlo Experiments

In order to assess the bias induced by the autocorrelation of cost-push shocks quantitatively, we turn to Monte Carlo experiments. As a baseline case we consider the GMM estimator proposed by GG. We first establish that it generally provides consistent estimates in the absence of autocorrelation of cost-push shocks. We then measure the bias induced by a high yet plausible degree of autocorrelation. In order to isolate the effect of autocorrelation on the estimation, we keep the volatility of cost-push shocks constant at the value implied by the estimates reported in Galí and Rabanal (2005), i.e. when varying ρ_u, we adjust σ^u accordingly.

[4]This result is certainly not new to the literature. For example, Mavroeidis (2004) also remarks that the empirical residuals will be serially correlated when conducting GMM with forward-looking models. He does not, however, appear to spell out the consequences for the Q-Test and the empirical practice.

In a similar fashion we afterwards study the behavior of the two step minimum distance (MD) estimator proposed by Sbordone (2002, 2005) and full-information maximum likelihood (FIML) methods in the presence of autocorrelated cost-push shocks. In each case, we simulate the model outlined in Section 2 and generate 1000 random time series. We consider a small yet realistic sample size of 152 observations and a longer time series of 2000 observations to assess the consistency of the estimator. In each case we use 100 additional observations to initiate the model. Throughout, we assume that the value of β is known to be 0.99, the value used in the simulation of the model, and focus on the estimates of the degree of price rigidity, $\hat{\theta}_p$. This in turn is linked to the slope of the Phillips curve by $\kappa_p = \frac{(1-\beta\theta_p)(1-\theta_p)}{\theta_p}$.

4.1 Generalized Method of Moments

Our experimental setup is meant to mimic the GMM estimation of the NKPC on U.S. data originally proposed in the seminal work of GG. Specifically, we estimate the parameter θ_p on the basis of the following moment condition[5]

$$E\left\{\left[\theta_p\pi_t - (1-\theta_p)(1-\beta\theta_p)l_t - \theta_p\beta\pi_{t+1}\right]\boldsymbol{z}_{t-1}\right\} = 0, \tag{9}$$

where the instrument vector $\boldsymbol{z}_{t-1}$ contains four lags of inflation, the labor share, and output growth. Our optimal weighting matrix uses the Newey-West correction for the likely serial correlation of the orthogonality conditions. When considering the distribution of point estimates, $\hat{\theta}_p$, recall that the true value is $\theta_p = 0.53$ in all simulations of the model.

The results of our experiments are displayed in Table 2. In the upper panel, we consider a sample size which is typical for macroeconomic studies: 152 usable observations as in Galí and Gertler (1999). The first row gives the estimated degree of price stickiness and the diagnostics if price markup shocks are uncorrelated in the true model economy, while the second row of the first panel shows results obtained on the basis of a higher degree of autocorrelation: $\rho_u = 0.95$, the value reported by Galí and Rabanal (2005). In the first case, the median estimate of $\hat{\theta}_p = 0.53$ corresponds to its true value. The mean duration of prices implied by the median estimate of $\hat{\theta}_p$ is displayed in the second column labeled 'D'. It is about 2 quarters.[6]

In contrast, if cost push shocks are autocorrelated, we obtain quite different estimates for the degree of price stickiness. In the small sample, we find a median estimate of $\hat{\theta}_p = 0.92$, which, in turn, implies average price durations of about 12 quarters. We are thus confronted with a substantial upward bias and – by any measure – an enormous degree of price rigidity. It is noteworthy, though, that such a

[5]Using the alternative moment condition, i.e. dividing equation (9) by θ_p, we find that the estimator for θ_p converges to unity quite frequently. We therefore report results obtained on the basis (9).

[6]Lindé (2005) investigates the properties of GMM in estimating the NKPC focusing on the extent of forward-looking behavior in price setting. He finds that even an uncorrelated cost-push shock induces some bias. In light of our results and the arguments put forward in Galí et al. (2005) this might be the result of his choice of instruments.

Table 2: RESULTS OF GMM ESTIMATION

	Point Estimates		Diagnostics		
	$\hat{\theta}_p$	D	F-Statistic	J-Statistic	Q(4)
152 Observations:					
$\rho_u = 0$	0.53	2.14	3.86	7.54	22.03
	(0.36,0.69)	(1.57,3.22)	[0.96]	[0.00]	[0.94]
$\rho_u = 0.95$	0.92	12.55	109.78	7.99	14.08
	(0.81,0.98)	(5.22,52.19)	[1.00]	[0.02]	[0.88]
2,000 Observations:					
$\rho_u = 0$	0.53	2.13	44.82	9.72	269.32
	(0.49,0.57)	(1.98,2.32)	[1.00]	[0.02]	[1.00]
$\rho_u = 0.95$	0.95	21.59	2131.66	27.35	159.96
	(0.90,0.99)	(9.59,70.53)	[1.00]	[0.81]	[1.00]

Notes: Values are median values over 1000 draws; for every draw of $\hat{\theta}_p$ the corresponding duration, D, and the diagnostics are computed. Values in parenthesis are 2.5% and 97.5% quantiles; except for the diagnostics where rejection frequencies for the null at a 5% level are given in square brackets. The "F-statistic" refers to an F-Test of the joint significance of the instruments when regressing the labor share on the instruments. The "J-statistic" refers to a J-Test for violation of the orthogonality conditions used in the GMM estimation. The "Q(4)" statistic refers to a Ljung-Box test for the presence of serial correlation of the error terms in the GMM estimation – the null being that these are serially uncorrelated.

high degree of price rigidity is also found by Galí and Gertler (1999) for some specifications of the NKPC model estimated on U.S. time series. In our setup, the estimator is also inconsistent in this case. In the lower panel of Table 2 we repeat the experiment but use 2000 observations in each draw in order to evaluate the large sample properties of the diagnostic tests. If 2000 observations are used instead of 152, we still find a median estimate of $\hat{\theta}_p = 0.95$.

To summarize, Figure 2 gives a graphical representation of the results for the small sample case. The left panel shows the distribution of $\hat{\theta}_p$ obtained if cost-push shocks are not persistent. These are well centered around the true value of $\theta_p = 0.53$. The right panel makes clear the bias in the estimates in the presence of serial correlation of the cost-push shock ($\rho_u = 0.95$). For the same realistically small sample size Figure 3 presents the small sample counterparts to the asymptotic bias reported in Figure 1. It plots the median estimate of θ_p against the various degrees of autocorrelation in the markup shock assumed in the simulation of the model. The bias increases monotonically with the degree of autocorrelation in line with the theoretical results obtained for the large-sample limit underlying Figure 1.

Given that the empirical model is misspecified under autocorrelated cost-push shocks, this result and

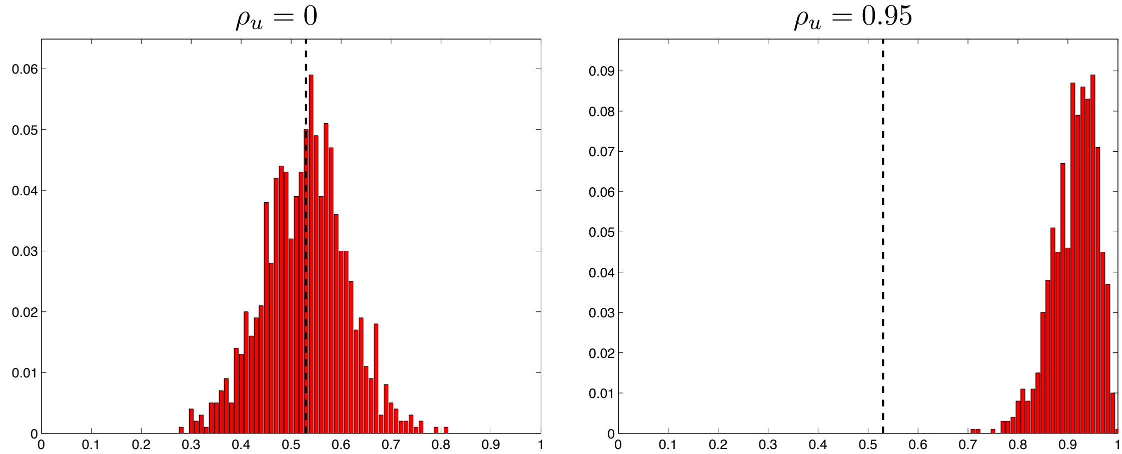

Figure 2: DISTRIBUTION OF GMM ESTIMATES $\hat{\theta}_p$, GIVEN 1000 DRAWS (152 OBSERVATIONS EACH). *Notes:* Vertical dashed line marks the true value, $\theta_p = 0.53$. We simulate 1000 time series of length 252 observations using the model by Galí and Rabanal (2005). For each simulated time series the first 100 observations of these are discarded.

in particular the direction of the bias may not come as a surprise given the arguments put forward in Section 3. The key question is whether standard diagnostic tests are able to detect the violation of the moment condition imposed in the GMM estimation in case of autocorrelated markup shocks.

Before turning to this question, however, we make sure that the we are not confronted with a weak instrument problem as discussed, for example, in the survey by Stock, Wright, and Yogo (2002) and Mavroeidis (2005). A weak instrument problem exists if the instruments have little predictive power for the instrumented variable. As a result, GMM estimates are biased and the estimators are $O_p(1)$, i.e. the dispersion of estimated parameters does not fall when the sample size increases. In our example, we reject the null of no significance for the instruments in a first stage regression of the labor share on the set of instruments in all specifications and especially strongly so if the cost-push shock is serially correlated. The results of the corresponding F-Test are displayed in the third column of Table 2. On top of this the GMM estimator is clearly consistent: If the model is correctly specified and there is no serial correlation in cost-push shocks the estimates converge to their true value. And also for the case of autocorrelated shocks, the standard deviation of the GMM estimates falls as the sample size increases, compare Table 2.

In the fourth column of Table 2 we report the J-Statistic which provides a test of the orthogonality conditions exploited in GMM estimation. For the small sample, the J-Test generally fails to detect the violation of the orthogonality conditions which were imposed in the GMM estimation: the null is rejected only for 2 percent of the draws. This squares well with Mavroeidis (2005). In his simulations, lags of inflation and the labor share are incorrectly not included as regressors in the GMM estimation

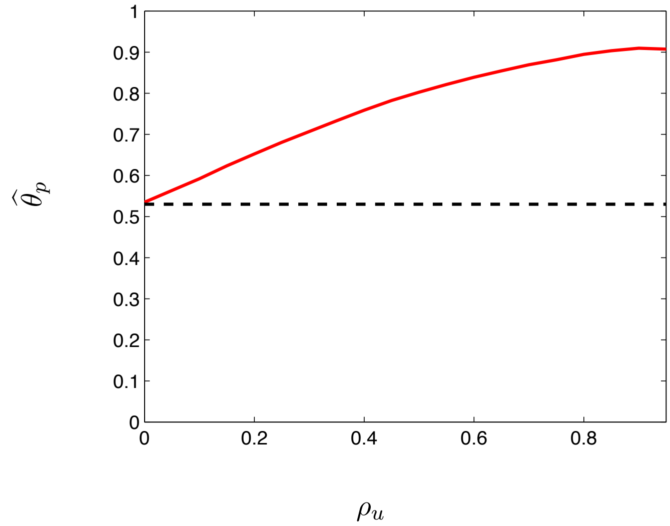

Figure 3: MEDIAN ESTIMATE FOR θ_p USING GMM, FOR INCREASING VALUES OF ρ_u. *Notes:* We simulate 1000 time series of length 252 observations using the model by Galí and Rabanal (2005). For each simulated time series the first 100 observations are discarded, so 152 observations are left for the estimation.

of a hybrid NKPC. While his focus is thus somewhat different from ours,[7] it is noteworthy that also in his simulations the J-Test has little power to detect the misspecification. Turning to the hypothetical sample size of 2000 observations in our simulations, the power of the J-Test increases considerably. We find the null correctly rejected for about 80 percent of the draws in case of serially correlated cost-push shocks ($\rho_u = 0.95$).

Turning to the Ljung-Box Q-test for autocorrelation of the residuals, we generally find high rejection frequencies for all specifications. Given the argument put forward in Section 3 this is not surprising either. The Q-test is a test on the correlation of the combined residual which reflects both the rational expectations forecast error and the lagged cost-push shock.[8] In empirical work, for instance, Galí et al. (2001) report a test statistic of $Q(4) = 10.2$ with a p-value of 4 percent, when estimating the NKPC on U.S. data.

As a final test of the model, we consider the forecasting performance of the estimated NKPC following GG and Galí et al. (2001). Solving the NKPC forward gives

$$\pi_t = \kappa_p \sum_{k=0}^{\infty} \beta^k E_t \{l_{t+k} + u_{t+k}\} = \kappa_p \sum_{k=0}^{\infty} \beta^k E_t \{l_{t+k}\} + \frac{\kappa_p}{1 - \beta\rho_u} u_t. \tag{10}$$

[7]He analyzes the resulting bias in estimates of the relative weight of the forward- and backward-looking component of a hybrid NKPC and is not concerned with the slope of the NKPC. For generating data for the simulations, Mavroeidis uses a reduced form equation for the labor share and the NKPC equation for inflation. In this setup the correlation between cost-push shocks and the labor share, which induces a downward bias in the slope estimates of the NKPC as highlighted above, cannot be examined.

[8]If one assumes that there are no cost-push shocks at all (σ_u=0), a case not reported in Table 2, the null of no autocorrelation in the residuals cannot be rejected.

In order to model the forecasts of the labor share, we employ a bivariate VAR(3) model in the labor share and inflation. Let $\boldsymbol{A}$ denote the companion matrix and $\boldsymbol{x}_t = [\ \pi_t, l_t, \ldots, \pi_{t-2}, l_{t-2}\ ]'$, then

$$\boldsymbol{x}_t = \boldsymbol{A}\boldsymbol{x}_{t-1} + \boldsymbol{\epsilon}_t, \tag{11}$$

where $\boldsymbol{\epsilon}_t$ denotes a vector of unidentified shocks. Let $\boldsymbol{\iota}_l$ be a vector with zeros except for a 1 in the position of the labor share. Then $l_{t+k|t} = \boldsymbol{\iota}_l' \widehat{A}^k \boldsymbol{x}_t$ is the VAR based forecast for the labor share. Substituting this for rational expectations in (10), we obtain

$$\pi_t = \kappa_p \boldsymbol{\iota}_l'(\boldsymbol{I} - \beta\hat{\boldsymbol{A}})^{-1}\boldsymbol{x}_t + \frac{\kappa_p}{1-\beta\rho_u}u_t, \tag{12}$$

where $\hat{\boldsymbol{A}}$ denotes the OLS estimate of the companion matrix $\boldsymbol{A}$. Assuming that the markup shock, u_t, is not observed, the following expression gives the NKPC-*cum*-VAR prediction for inflation, referred to as fundamental inflation by GG:

$$\pi_t^* = \kappa_p \boldsymbol{\iota}_l'(\boldsymbol{I} - \beta\hat{\boldsymbol{A}})^{-1}\boldsymbol{x}_t. \tag{13}$$

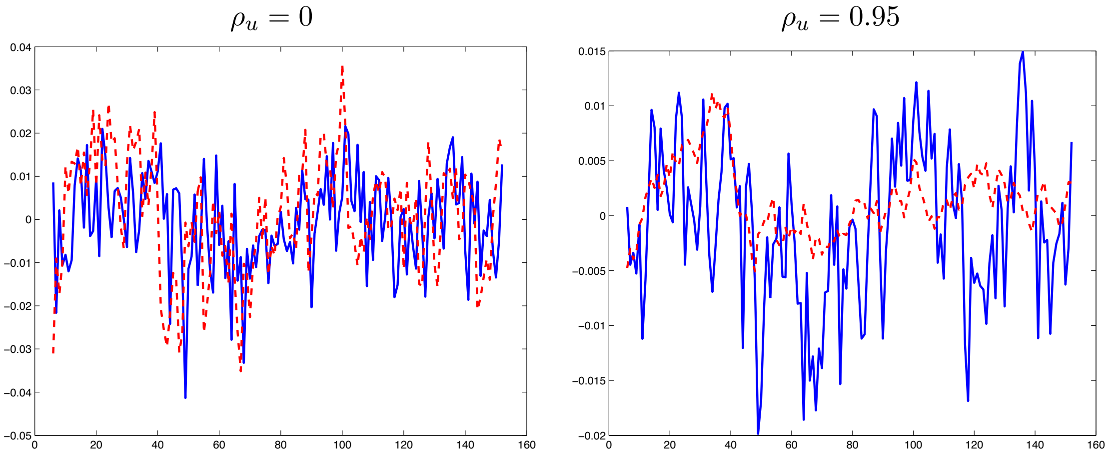

Figure 4: ACTUAL (SOLID LINE) VS. FUNDAMENTAL (DASHED LINE) INFLATION. *Notes:* We simulate randomly one set of time series using the model by Galí and Rabanal (2005). Fundamental inflation is obtained according to (13). We repeat the experiment twice: once assuming away the persistence of cost-push shocks (left panel) and once using the model with serially correlated cost-push shocks as the data-generating process.

Figure 4 plots actual and fundamental inflation for the case of uncorrelated cost-push shocks (left panel) and the case of highly autocorrelated cost-push shocks (right panel). Clearly the model performance is only fully convincing in the first case. Moreover, casual inspection suggests that the NKPC model may do a better job if confronted with actual data as in GG and Galí et al. (2001).

In sum, the autocorrelation of cost-push/markup shocks induces a substantial upward bias in the estimated degree of price rigidity. At the same time it is unclear whether such a bias would be detectable in actual data, given the ambivalence of the signals obtained from standard diagnostic tests.

4.2 Minimum Distance estimation

We now turn to an alternative approach to estimate the NKPC which has generally led to similar results for the degree of price rigidity as the GMM approach. In a seminal paper, Sbordone (2002) suggests a two stage minimum distance (MD) procedure to estimate the NKPC. The idea is to find an estimate for the degree of price rigidity by matching the actual path of the price-unit-labor-cost-ratio and the corresponding ratio predicted by the model under the auxiliary assumption of a VAR-based forecasting model. In the following, we draw on Sbordone (2005) in modifying the two stage MD estimator to match inflation dynamics directly.

Specifically, defining $\boldsymbol{\iota}_\pi'$ such that $\pi_t = \boldsymbol{\iota}_\pi' \boldsymbol{x}_t$, and assuming that i) $\rho_u = 0$ and ii) that the forecasts based on the VAR model (11) provide a good approximation to the actual rational expectations forecasts of the labor share, we can write the forward solution of the NKPC (10) as follows

$$\boldsymbol{\iota}_\pi' \boldsymbol{x}_t \approx \kappa_p \boldsymbol{\iota}_l' (\boldsymbol{I} - \beta \hat{\boldsymbol{A}})^{-1} \boldsymbol{x}_t + \tilde{u}_t.$$

Assuming that $E(\tilde{u}_t | \boldsymbol{x}_{t-1}) = 0$, conditioning on the information in $\boldsymbol{x}_{t-1}$ and the VAR model (11) gives

$$\boldsymbol{\iota}_\pi' \hat{\boldsymbol{A}} \boldsymbol{x}_{t-1} \approx \kappa_p \boldsymbol{\iota}_l' (\boldsymbol{I} - \beta \hat{\boldsymbol{A}})^{-1} \hat{\boldsymbol{A}} \boldsymbol{x}_{t-1}. \tag{14}$$

Sbordone (2005) emphasizes that as the VAR provides an unrestricted interpretation of the data, it serves as a natural benchmark against which the restrictions imposed by the NKPC can be tested while estimating the degree of price rigidity. More formally, rewriting (14) gives rise to the following vector function,

$$F(\theta_p, \hat{\boldsymbol{A}})' = \boldsymbol{\iota}_\pi' \hat{\boldsymbol{A}} - \kappa_p(\theta_p) \boldsymbol{\iota}_l' (\boldsymbol{I} - \beta \hat{\boldsymbol{A}})^{-1} \hat{\boldsymbol{A}}, \tag{15}$$

which is approximately zero under the null of the NKPC. Given an appropriately defined weighting matrix, $\boldsymbol{\Sigma_A}^{-1}$, the two step minimum distance (MD) estimator is defined as[9]

$$\hat{\theta}_p = \arg\min F(\theta_p, \hat{\boldsymbol{A}})' \boldsymbol{\Sigma_A}^{-1} F(\theta_p, \hat{\boldsymbol{A}}). \tag{16}$$

Sbordone (2005) also employs a Wald statistic to test formally the restrictions imposed on the data by forcing the function (15) to be as close to zero as possible:

$$W = F(\widehat{\theta}_p)' \boldsymbol{\Sigma_F}^{-1} F(\widehat{\theta}_p), \tag{17}$$

where $\boldsymbol{\Sigma_F}$ denotes the covariance matrix of $\hat{F}$.[10]

[9] In practice we follow Sbordone (2005) and use a diagonal weighting matrix containing the inverse of the variance of the VAR coefficients $\boldsymbol{\iota}_\pi' \widehat{\boldsymbol{A}}$. As in the GMM specification we also multiply the vector function with θ_p.

[10] Below we compute the covariance matrix $\boldsymbol{\Sigma_F}$ on the basis of the sequence $\{F(\hat{\theta}_p)_i\}$ with draws $i \in \{1, 1000\}$. For each i we compute a Wald statistic using $F(\hat{\theta}_p)_i$ and $\boldsymbol{\Sigma_F}$.

In order to assess the performance of the two stage MD estimator we repeat the Monte Carlo experiment, i.e. we solve problem (16) for data simulated under the assumption of two different data-generating processes. One of these maintains that cost-push shocks are uncorrelated ($\rho_u = 0$). The other one entertains correlated cost-push shocks ($\rho_u = 0.95$).

Table 3: RESULTS OF MD ESTIMATION

	Point Estimates		Diagnostics
	$\hat{\theta}_p$	D	Wald-Test
152 Observations:			
$\rho_u = 0$	0.52 (0.34,0.67)	2.10 (1.51,3.04)	5.11 [0.06]
$\rho_u = 0.95$	0.97 (0.89,1.00)	37.94 (9.42,240.93)	5.34 [0.05]
2,000 Observations:			
$\rho_u = 0$	0.50 (0.45,0.55)	2.00 (1.81,2.21)	5.12 [0.06]
$\rho_u = 0.95$	0.98 (0.97,0.99)	59.09 (36.22,98.13)	5.36 [0.04]

Notes: Values are median values over 1000 draws. Values in parentheses are the corresponding 2.5% and 97.5% quantiles; except for the Wald-Test. There the values in square brackets refer to the rejection frequency of the null at a 5% level.

Table 3 shows the results. In case of uncorrelated cost-push shocks the performance of the estimator is satisfactory. Interestingly, for the large sample case, the median estimate is somewhat lower than the true value (lower panel, 2000 observations).[11]

In the case of autocorrelated cost-push shocks, the estimates are strongly upward biased; in the small sample we find a median estimate of $\hat{\theta}_p = 0.97$, implying a mean duration of more than 30 quarters, when, in fact, the true value is about two quarters. Importantly, however, the model passes the Wald-Test (last column) in 95% of the times also in case of autocorrelated cost-push shocks although the Wald-Test should firmly reject the null.

Before turning to a brief assessment of full information estimation methods, we note that the NKPC has also frequently been estimated on the basis of an alternative minimum distance approach, namely

[11]Given that the estimation is based on the auxiliary forecasting model (11) which does not nest the exact VARMA process underlying the expectation formation, the two stage MD estimator need not deliver consistent estimates. More generally, results by Kurmann (2005) suggest that the model performance depends on the exact specification of the VAR model used to model the forecasting process.

by matching impulse responses to a monetary policy shock. In this case it is necessary to modify the baseline New Keynesian general equilibrium model such that it conforms with the recursive identification assumptions generally imposed in structural VAR models.[12] We also performed a Monte Carlo experiment using an appropriately modified model. It turns out that in this case estimates of θ_p are not biased even if cost-push shocks are autocorrelated.[13] Eventually one's stand regarding the identifying assumptions is therefore crucial. As this is a topic beyond the scope of the present paper, we just note that results regarding the degree of price rigidity obtained within this literature tend to be quite heterogenous. Christiano et al. (2005), for instance, find $\hat{\theta}_p = 0.6$, where the model is explicitly designed to imply low price rigidities by muting movements in marginal costs. In contrast, the paper by Rotemberg and Woodford (1997), for example, finds that the pass-through of marginal costs into prices is very low in fact.

4.3 Maximum Likelihood estimation

We now turn to full-information maximum likelihood (FIML) estimation and ask once more to which extend the omission of an autocorrelated cost-push term may induce a bias towards too much estimated price stickiness. At the same time, we investigate whether this bias would be identifiable in at least a subset of the estimation statistics.

In the first experiment, we assume that estimation is conducted with a misspecified model. The model is estimated without allowing for serial correlation in the cost-push shock, $\rho_u = 0$, while for the actual data generating process we assume that $\rho_u = 0.95$. In the estimation we give the econometrician more than a fair chance by assuming that she has full knowledge about all the parameters of the model but for the degree of price stickiness, θ_p, (or equivalently the slope of the Phillips curve, κ_p) and the stochastic structure, i.e. the standard deviations of the innovations to the shocks. In addition, apart from the serial correlation of the cost-push shock, the corresponding serial correlation parameters of all shocks are assumed to be known and are set to their true values. For each draw then, the degree of price stickiness, θ_p, and the standard deviations of the innovations to the shock processes are estimated. Also in this example, we leave the standard deviation of the cost-push shock, u_t, in the data-generating process unchanged as we vary the autocorrelation of the shock, ρ_u.

Figure 5 shows the histograms of the distribution of the estimates for θ_p and σ^u over the different draws when the estimation is conducted with the misspecified model which does not feature serial correlation in the cost-push shock. A dashed vertical line marks the true values. The sample length is 152 quarters. Also with FIML, there is a considerable downward bias in the estimates of the

[12] In this context, Kehoe (2006) stresses that most monetary models do not satisfy the recursiveness assumption. However, without the assumption of predetermined prices and output, the responses of the baseline New Keynesian cannot be successfully matched with those obtained from a VAR estimated on data generated by the model.

[13] This illustrates the robustness of this limited information approach with respect to a misspecification of the stochastic structure of the model, see Meier and Müller (2006) for further discussion of this point.

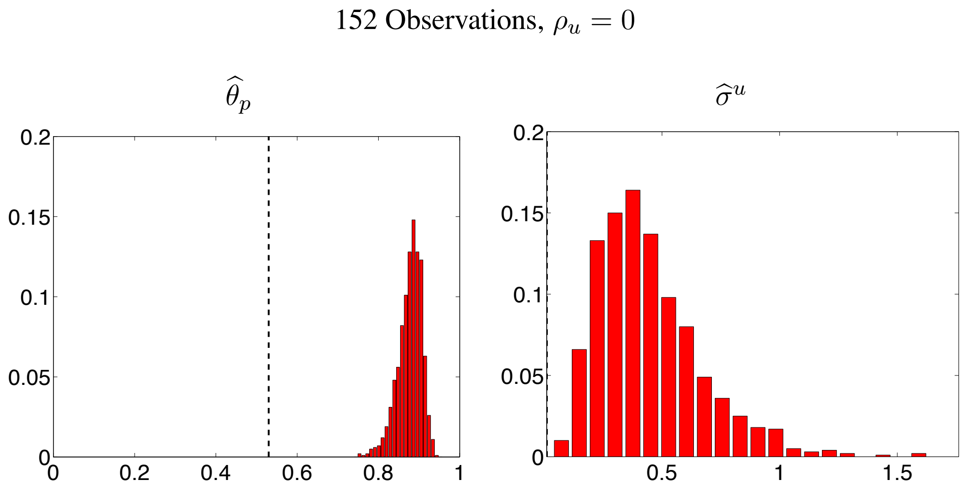

Figure 5: DISTRIBUTION OF ESTIMATED CALVO PARAMETER AND STANDARD DEVIATION OF INNOVATION TO THE COST-PUSH SHOCK. *Notes:* The histograms are obtained conducting FIML estimation using 1000 draws with sample length of 152 quarters each in a misspecified model which sets $\rho_u = 0$. For details on the estimation, see the notes to Table 4. In each of the plots, a vertical black dashed line marks the true parameter value which had been used to generate the data. From left to right, the parameters are: price stickiness θ_p and the standard deviation of the innovation to the cost-push shock. For the estimate of the standard deviation of innovations to the cost-push shock, $\widehat{\sigma}_u$, the dashed line signalling the true value is adjacent to the left vertical axis.

slope of the Phillips curve. As a result of omitting serially correlated cost-push shocks from the estimated model, the econometrician will be lead to infer that the economy features considerable price stickiness and that, as a consequence, the NKPC is flat. The procedure significantly overestimates price stickiness, θ_p, as is evident from the left panel of Figure 5.

Table 4 summarizes these results and provides more details about the estimation procedure. In the misspecified model, the median price stickiness parameter is estimated to be equal to $\hat{\theta}_p = 0.88$ with a narrow empirical standard deviation. This parameter value would imply that prices on average remain non-optimized for slightly longer than two *years* instead of the two *quarters* in the true data generating process. Not surprisingly perhaps, a noticeable bias also emerges for the standard deviation of the innovation to the cost-push shock (right panel in Figure 5 and first row, fourth column in Table 4).[14]

In a second experiment, we consider the case in which the econometrician allows for serial correlation in the cost-push shock and thus estimates a correctly specified model. In this case the bias vanishes altogether, as the second row of Table 4 illustrates.

We next turn to the question whether this bias would be recognizable to an economist/econometrician

[14]It is interesting to observe that in estimating a similar model using Bayesian techniques Rabanal and Rubio-Ramírez (2005) do not allow for autocorrelation of markup shocks. As in Smets and Wouters (2003) their estimates point to a large standard deviation of cost-push shocks which dwarfs those of the other shocks to a similar extent as in our simulations when the data feature serially correlated cost-push shocks but this is not taken into account in the estimation.

Table 4: PARAMETER ESTIMATES WITH MAXIMUM LIKELIHOOD

	$\widehat{\theta}_p$	$\widehat{\sigma}^a \cdot 100$	$\widehat{\sigma}^g \cdot 100$	$\widehat{\sigma}^u \cdot 100$	$\widehat{\sigma}^w \cdot 100$	$\widehat{\sigma}^m \cdot 100$	$\widehat{\rho}_u$
False perception	0.88 (0.81,0.92)	0.90 (0.80,1.00)	2.33 (2.04,2.61)	40.43 (14.01,97.70)	1.23 (1.07,1.38)	0.30 (0.26,0.33)	0.00 –
Correct perception	0.53 (0.48,0.58)	0.90 (0.80,1.00)	2.50 (2.21,2.75)	1.10 (0.96,1.27)	1.19 (1.06,1.35)	0.30 (0.26,0.33)	0.95 (0.90,0.98)
True values	0.53	0.90	2.50	1.10	1.20	0.30	0.95

Notes: Median parameter estimates using 1000 runs of the model with a sample length of 152 observations. An additional 100 observations at the beginning of each simulated series were discarded. Values in in parenthesis refer to the 2.5% quantile and the 97.5% quantile of the distribution of the estimates. From left to right: estimate of the Calvo probability of not updating, estimates of the standard deviations of the innovations to the technology shock, the demand shock, the cost-push shock, the wage-markup shock and the monetary shock (all estimates of innovations are multiplied by a factor of 100). Final column: estimated autocorrelation of the cost-push shock. All other parameters are assumed to be known by the econometrician. First row: The econometrician does not account for persistence in the cost-push shock and sets $\rho_u = 0$ in the estimation process. Second row: ρ_u is left free in the estimation process. Final row: true parameter values underlying the data-generating process. The choice of observable data for the estimation follows Galí and Rabanal (2005): the nominal interest rate, inflation, real wage to output ratio, hours worked and output growth.

Table 5: RMSE WITH MAXIMUM LIKELIHOOD

	r_t	π_t	$\tilde{w}_t - \tilde{y}_t$	n_t	Δy_t
False perception	0.27	0.66	1.35	1.21	1.22
Correct perception	0.22	0.51	1.29	1.17	1.18
Relative RMSE (in %)	22.35	29.35	5.00	3.41	3.22
% of draws for which false model has lower RMSE	0.00	0.00	3.40	4.90	5.90

Notes: Mean root mean squared errors (in sample, multiplied by 100) using 1000 runs with a sample length of 152 observations. The true correlation of the cost-push shock is ρ_u =0.95. The perceived (false) correlation of the cost-push shock is ρ_u =0.00. See Table 4 for details. From left to right: nominal interest rate, inflation, real wage output ratio, hours worked and output growth.

who is conducting estimation and inference on just one set of data. Towards this end Table 5 shows mean root-mean squared forecast errors (in sample RMSE) for both estimated variants of the model. Evidently, the correctly specified model dominates the misspecified model in both the inflation forecast and the interest rate forecast dimension. Yet, while the RMSEs speak a very clear language and so would likelihood ratio tests, most economists who firmly believe that their estimated model should feature no serial correlation in the cost-push shock will, we suppose, be quite hesitant to opt for the larger model despite the econometric evidence.[15]

Summarizing, autocorrelated cost-push shocks, whenever they are not properly accounted for in the estimation process, induce a considerable upward bias in the estimated degree of price rigidity. As with the GMM and MD single-equation approaches therefore, also when using FIML estimation of the model, the estimates of the slope of the NKPC are biased downwards. The same generic economic mechanism is behind the bias in all of theses cases: serially correlated cost-push shocks induce a persistent opposite reaction of the labor share in the data-generating model (i.e. whenever cost-push shocks are positive, the labor share all else equal is lower). Assume that an econometrician believes that cost-push shocks are white noise and incorporates this view into his model of the economy. A labor share which persistently deviates from its steady state value then must be driven by other factors apart from cost-push shocks. That this persistently, say, high labor share according to the observed data is non-inflationary at times is therefore reconciled with the NKPC embedded into the model by underestimating the slope of the Phillips curve – or, equivalently, by estimating too much price rigidity.

5 Conclusion

Is the New Keynesian Phillips curve actually flat? In this paper we suggest a new interpretation of the macroeconometric evidence, which is consistent with microeconometric studies implying frequent and sizeable price adjustments. We start from the observation that full information estimation of medium-scale New Keynesian general equilibrium models provides evidence in favor of highly autocorrelated cost-push/markup shocks, see, e.g. Galí and Rabanal (2005) and Smets and Wouters (2006).

We show that in this case the orthogonality conditions exploited in the GMM approach to the New Keynesian Phillips Curve, which was pioneered by Galí and Gertler (1999), are likely to be invalid. In fact, estimates can be shown to be inconsistent and upward biased. We assess the quantitative implications of these complications through Monte Carlo experiments. As the data generating process

[15]On the economic side, current generation DSGE models are often criticized for lacking internal propagation, see, for instance, Cogley and Nason (1995). Serially correlated shocks can partially make up for the resulting lack of persistence. This may, however, have little appeal for many economists since economic theory provides little guidance with regard to the autocorrelation structure of shocks.

we use a variant of the standard New Keynesian model estimated and shown to provide a satisfactory account of U.S. time series by Galí and Rabanal (2005).

In the data generating process prices have a mean duration of about 2 quarters and the degree of autocorrelation of cost-push shocks is 0.95. Using a sample size of 152 observations, which reflects the typical sample size available to macroeconomists, we find that the average estimate for the degree of price rigidity implies a mean duration of prices of about 12 quarters. Interestingly, standard tests fail to detect the misspecfication of the model. Moreover, the bias is not limited to GMM estimation, but also shows if the two stage minimum distance estimator suggested by Sbordone (2002, 2005) or classical ML techniques are employed.

An intuitive interpretation of our results is as follows. Positive cost-push/markup shocks induce a negative response of the labor share, the measure for marginal costs used in the literature. To the extent that cost-push shocks play a role in driving the business cycle, a high labor share will frequently come along with a lower than average realization of the cost push-shock and this will be persistently the case if cost-push shocks are autocorrelated. The partial inflationary effect of the labor share (wages) on inflation will thus be underestimated whenever the serial correlation in the markup shock is not accounted for. Put differently, if cost-push shocks are an important source of business cycle fluctuations, but their autocorrelation is not considered in the estimation, the New Keynesian Phillips curve will appear to be implausibly flat, while, in fact, it is quite steep as would be suggested by microeconomic evidence.

In principle, there are two interpretations of our results each of which implies an avenue to test the arguments of the present study against actual time series data. First, if cost-push/markup shocks are interpreted as a deep or fundamental feature of the data generating process, full information estimation techniques are adequate. Using these techniques one can assess whether removing the restriction of no-autocorrelation of cost-push shocks leads to increases in the estimated slope coefficient of the NKPC and thus to smaller degrees of implied price rigidity. Alternatively, one may interpret the estimated shock process as merely semi-structural, such that when the process of marginal costs is modeled in greater detail these 'shocks' disappear from the data. In this regard, it is interesting to note that Batini, Jackson, and Nickell (2005) estimate a variant of the NKPC on U.K. data considering that in addition to the labor share, employment adjustment costs, oil and import prices may drive inflation.[16] If this is indeed the case and better proxies for marginal costs can be found, the slope estimates from both the traditional single-equation NKPC approach and the full-information simultaneous-equation approaches would likely need to be revised.

While future research will show which of these interpretations will prevail, the paper has one unambiguous conclusion: it appears inconsistent to base ones reasoning about price rigidity and about

[16]In line with the second interpretation of our results, they report a coefficient on the labor share which is considerably higher than what is typically found for standard specifications of the model.

the slope of the Phillips curve on both the single-equation GMM approach and minimum distance estimates existing in the literature and, at the same time, on the results of the estimation of DSGE models which feature highly serially correlated cost-push shocks. At most one of the approaches is valid.

Appendix: Properties of the GMM estimator

This appendix derives the probability limit of the estimator in equation (7). To assess this more formally we follow Hamilton (1994, p. 238f.). We consider the convergence properties when both the discount factor, β, and the slope of the Phillips curve, κ_p, are estimated. The formulae in Section 3, which restrict the estimation to κ_p can be obtained following a modification of the line of exposition below.

First, we note that absent indexation equation (3) can be written as follows

$$y_t = \boldsymbol{\alpha}' \boldsymbol{x}_{t+1} + \epsilon_{t+1}^{RE} + \tilde{u}_t,$$

where $\boldsymbol{x}_{t+1} = [\pi_{t+1}, l_t]'$, $\boldsymbol{\alpha} = [\beta, \kappa_p]'$ and ϵ_{t+1}^{RE} is the rational expectations error that ensures $E_t(\pi_{t+1}) \equiv \pi_{t+1} + \epsilon_{t+1}^{RE}$ and $\tilde{u}_t = \kappa_p u_t$. As the above equation is linear, estimating it by GMM is equivalent to applying two stage least squares estimation (2SLS henceforth). Let the instrument vector be $\boldsymbol{z}_{t-1} = [\pi_{t-1}, l_{t-1}]'$.

The 2SLS estimator of $\boldsymbol{\alpha}$ when the sample size is T is

$$\widehat{\boldsymbol{\alpha}}_{2SLS,T} = \left[\sum_{t=1}^{T} \widehat{\boldsymbol{x}}_{t+1} \boldsymbol{x}_{t+1}' \right]^{-1} \left[\sum_{t=1}^{T} \widehat{\boldsymbol{x}}_{t+1} y_t' \right].$$

Above $\widehat{\boldsymbol{x}}_{t+1} = \widehat{\boldsymbol{\delta}}_T' \boldsymbol{z}_{t-1}$, the OLS fitted value when regressing $\boldsymbol{x}_{t+1}$ on the instruments with

$$\widehat{\boldsymbol{\delta}}_T = \left[\sum_{t=1}^{T} \boldsymbol{z}_{t-1} \boldsymbol{z}_{t-1}' \right]^{-1} \left[\sum_{t=1}^{T} \boldsymbol{z}_{t-1} \boldsymbol{x}_{t+1}' \right].$$

Using the Phillips curve to substitute for y_t in the 2SLS estimator, we obtain

$$\widehat{\boldsymbol{\alpha}}_{2SLS,T} = \left[\sum_{t=1}^{T} \widehat{\boldsymbol{x}}_{t+1} \boldsymbol{x}_{t+1}' \right]^{-1} \left[\sum_{t=1}^{T} \widehat{\boldsymbol{x}}_{t+1} (\boldsymbol{\alpha}' \boldsymbol{x}_{t+1} + \epsilon_{t+1}^{RE} + \tilde{u}_t)' \right],$$

So

$$\widehat{\boldsymbol{\alpha}}_{2SLS,T} = \boldsymbol{\alpha} + \left[\frac{1}{T} \sum_{t=1}^{T} \widehat{\boldsymbol{x}}_{t+1} \boldsymbol{x}_{t+1}' \right]^{-1} \left[\frac{1}{T} \sum_{t=1}^{T} \widehat{\boldsymbol{x}}_{t+1} (\epsilon_{t+1}^{RE} + \tilde{u}_t) \right].$$

Assuming that $\boldsymbol{x}_{t+1}$ and $\boldsymbol{z}_{t-1}$ are jointly covariance-stationary and ergodic for second moments, we have that

$$\widehat{\boldsymbol{\delta}}_T \xrightarrow{p} \left[E(\boldsymbol{z}_{t-1} \boldsymbol{z}_{t-1}') \right]^{-1} \left[E(\boldsymbol{z}_{t-1} \boldsymbol{x}_{t+1}') \right] =: \overline{\boldsymbol{\delta}}.$$

Under the same assumptions

$$\frac{1}{T} \sum_{t=1}^{T} \widehat{\boldsymbol{x}}_{t+1} \boldsymbol{x}_{t+1}' \xrightarrow{p} Q,$$

where $Q = \overline{\boldsymbol{\delta}}' \left[E(\boldsymbol{z}_{t-1} \boldsymbol{x}_{t+1}') \right]$.

In addition,

$$\left[\frac{1}{T}\sum_{t=1}^{T}\widehat{\boldsymbol{x}}_{t+1}(\epsilon_{t+1}^{RE}+\tilde{u}_t)\right]=\left[\frac{1}{T}\sum_{t=1}^{T}\widehat{\boldsymbol{x}}_{t+1}\epsilon_{t+1}^{RE}\right]+\left[\frac{1}{T}\sum_{t=1}^{T}\widehat{\boldsymbol{x}}_{t+1}\tilde{u}_t\right],$$

the first term of which will typically converge to zero in probability by means of a suitable law of large numbers. The final term, however, is given by

$$\left[\frac{1}{T}\sum_{t=1}^{T}\widehat{\boldsymbol{x}}_{t+1}\tilde{u}_t\right]=\widehat{\boldsymbol{\delta}}_T'\left[\frac{1}{T}\sum_{t=1}^{T}\boldsymbol{z}_{t-1}\tilde{u}_t\right].$$

Taking the probability limit we thus have that

$$\left[\frac{1}{T}\sum_{t=1}^{T}\widehat{\boldsymbol{x}}_{t+1}(\epsilon_{t+1}^{RE}+\tilde{u}_t)\right]\xrightarrow{p}\overline{\boldsymbol{\delta}}'\left[E(\boldsymbol{z}_{t-1}\tilde{u}_t)\right].$$

Note that in our example

$$E(\boldsymbol{z}_{t-1}\tilde{u}_t)=E\left(\left[\begin{array}{c}\pi_{t-1},l_{t-1}\end{array}\right]'\tilde{u}_t\right)=\rho_u\kappa_p E\left(\left[\begin{array}{c}\pi_{t-1},l_{t-1}\end{array}\right]'u_{t-1}\right)\neq 0 \text{ unless } \rho_u=0.$$

Summarizing,

$$\widehat{\boldsymbol{\alpha}}_{2SLS,T}\xrightarrow{p}\boldsymbol{\alpha}+Q^{-1}\overline{\boldsymbol{\delta}}'\kappa_p\rho_u E\left(\left[\begin{array}{c}\pi_{t-1},l_{t-1}\end{array}\right]'u_{t-1}\right),$$

illustrating the inconsistency of the GMM approach with serially correlated price-markup shocks. In Figure 6, we illustrate the inconsistency of the estimates for β and κ_p in the calibrated model outlined in Section 2 for various values of ρ_u, confirming the results discussed in the main text. Note that as β is left unrestricted in the estimation, the bias in the slope of the Phillips curve is hardly affected. Since estimates of β are upward biased, the implied price stickiness shows a smaller bias than in the case presented in Section 3.

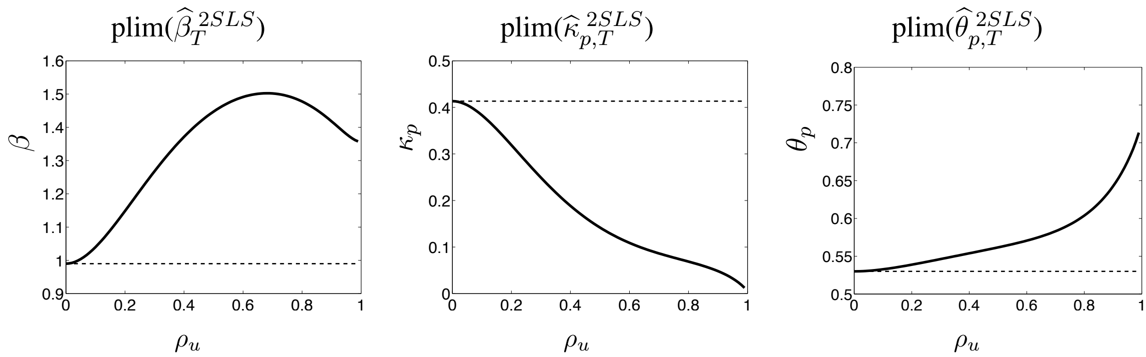

Figure 6: β AND κ_p ARE ESTIMATED. THEORETICAL PROBABILITY LIMITS OF 2SLS ESTIMATORS FOR VARYING DEGREES OF SERIAL CORRELATION, ρ_u, IN THE COST-PUSH SHOCK. *Notes:* Dashed lines show the true values. From left to right: estimator for the discount-factor, β, for the slope of the Phillips curve, κ_p, and for the fraction of firms which do not reoptimize their price, θ_p.

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