Does the cost channel matter for inflation dynamics? An identification robust structural analysis for the US and the Euro area

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PRELIMINARY AND INCOMPLETE

This version: October 2009

Abstract

This paper asks whether interest rate changes have an impact on firms marginal cost and whether this has a direct effect on price setting behavior of firms which translates into aggregate inflation dynamics. Empirical tests of the existence of the cost channel are employed using a structural econometric approach. Estimation and inference is conducted using identification robust methods based on the continuous-updating GMM objective function. We document identification difficulties for some parameters when estimating the general model structure. For the US, a pure forward-looking interest rate augmented Phillips curve is most compatible with data. This suggests that considering the cost channel is a non-negligible aspect for monetary policy.

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Contents

1 Introduction

This paper asks whether costs of external funds affect firms' marginal cost and thus influence the aggregate inflation rate. Recently, many authors – including Christiano, Eichenbaum and Evans (2005), Chowdhury, Hoffmann and Schabert (2006), Ravenna and Walsh (2006) and Tillmann (2008) – provide evidence of a cost channel relevant for inflation dynamics. This cost channel is introduced through the cost of working capital which is motivated by cash-in-advance, i.e. factors of production which have to be paid before the proceeds from sale of output are received. Empirically, the existence of a cost channel can be tested by augmenting the New Keynesian Phillips curve by an interest rate variable as an additional regressor. So the cost channel implies an extension of the standard measure of marginal cost by interest rate effects.

Chowdhury et al. (2006) test such an augmented Phillips curve specification for G7 countries with GMM and find empirical support for this model for most of the countries. Ravenna and Walsh (2006) employ the same method but instead of relaying on the reduced form parameters, they estimate structural parameters of a pure forward-looking specification for the US and draw similar conclusions. The existence of a cost channel is also supported by methods of indirect inference (e.g. Christiano et al., 2005; Huelsewig, Henzel, Wollmershaeuser and Mayer, 2008). But there are also studies that cast doubt about the existence of a cost channel (e.g. Rabanal, 2007; Gabriel, Levine, Spence and Yang, 2008). Their estimation include Bayesian Methods (Rabanal, 2007) as well as GMM (Gabriel et al., 2008).

In this paper we extend the Phillips curve specification of Ravenna and Walsh (2006) to a model which allows for backward looking behavior in price setting due to partial indexation. Then we reexamine the existence of the cost channel by estimating reduced form parameters (similar to Chowdhury et al., 2006) as well as structural parameters for a large group of industrialized countries. Instead of relaying on a standard two-step GMM estimator we use a continuous-updating GMM (CUE) estimator as proposed by Hansen, Heaton and Yaron (1996). This estimator is preferable in terms of small sample properties. Additional, it does not depend on the normalization of the orthogonality conditions. Moreover, joint confidence sets are constructed by using the CUE objective function. Stock and Wright (2000) show that this method for inference is a generalization of the Anderson-Rubin (AR) statistic and that it is fully robust to problems accotiated with weak identification. So this procedure guards against problems induced by weak instruments that might be present in estimates of the new Phillips curve (see Ma, 2002; Mavroeidis, 2005; Dufour, Khalaf and Kichian, 2006). Confidence intervals for the individual parameters are then computed with the projection technique.

The results of this paper indicate that the empirical evidence of the cost channel is mixed. Generally, it is confirmed that weak instrument problems are present in estimating the NKPC model. In particular, it is evident that distinguishing between forward looking and backward looking behavior is very hard. Additionally, the estimates of the slope coefficient of the interest rate variable are very imprecise and often insignificant. However, for the US a pure forward looking model that considers the cost channel performs reasonably well. But once the interest rate in the inflation model is omitted the model is statistically rejected. So the inclusion of a cost channel can indeed improve the reliability of estimates of the New Keynesian Phillips curve by the introduction of the interest rate affecting real marginal costs.

The rest of the paper is organized as follows. Section 2 introduces the theoretical model setup. The empirical strategy is outlined in Section 3. Section 4 presents the estimation results of the interest-rate augmented Phillips curve. Section 5 concludes.

2 The basic model

This section briefly introduces the theoretical model that consists of a standard New Keynesian framework. More detailed derivations may be found in Walsh (2003) and Woodford (2003). We concentrate on aspects necessary to characterize inflation dynamics in the economy. The two basic model features consist of monopolistically competitive goods markets and sticky prices as well as the introduction of the cost channel.

More precisely, the economy consists of a continuum of firms (indexed by $i \in [0,1]$) each producing a differentiated good $Y_t(i)$ according to a standard Cobb-Douglas production function

$$Y_t(i) = A_t \overline{K}_t(i)^{\alpha} N_t(i)^{1-\alpha}, \tag{1}$$

with A_t a common country wide technological factor, $\overline{K}_t(i)$ the (fixed) firm-specific capital stock and $N_t(i)$ denoting the labor factor employed by firm i.

Each firm i faces a demand function characterized by constant elasticity of substitution given by

$$Y_t(i) = \left(\frac{P_t(i)}{P_t}\right)^{-\epsilon} Y_t, \tag{2}$$

where Y_t equals aggregate demand, P_t is the aggregate price level in the economy and $P_t(i)$ is the price of good i that is charged by firm i. The price elasticity of demand for good i is characterized by the parameter ϵ (with $\epsilon > 1$). This determines the constant mark-up (defined as $\mu = \epsilon/(\epsilon - 1)$) that firms require over nominal marginal costs of inputs.

Next, we introduce a liquidity constraint for firms operating in their factor markets. Input factors such as the wage bill has to be paid before revenues for the produced good have been received. To meet these expenditures, firms have to borrow these outlays from a financial intermediary sector. Each period the individual firm i is assumed to borrow the amount $Z_t(i)$ to pay the sum of salaries. So the liquidity constrain is given by

$$Z_t(i) \geq W_t N_t(i),$$

with W_t the nominal wage rate and $N_t(i)$ the utilized labor factor of firm i. At the end of the period when the produced good has been sold, firms have to repay these loans with an interest of the amount of $i_t^l Z_t(i)$. With these liquidity constrains firms marginal costs are equal to

$$MC_t(i) = \frac{R_t^l W_t / P(i)_t}{(1-\alpha) Y_t(i) / N_t(i)} = \frac{R_t^l S_t(i)}{(1-\alpha)}, \tag{3}$$

where $R_t^l = 1 + i_t^l$ and $S_t(i)$ is the firm specific labor share of production.

Further, we assume that firms face nominal price rigidities that can be characterized by Calvo's (1983) model of staggered price setting. This model implies that firms set prices infrequently due to costs of information gathering. The frequency of price re-optimizations is characterized by a stochastic process with a constant probability that a firm changes its price at one particular point in time. So on the aggregate level at each point in time there is a fraction of firms $1-\theta$ that optimally adjusts prices. The expected waiting time is then given by $1/(1-\theta)$.

Price re-optimizing firms that set their optimal price $P_t^*(i)$ are faced with the following dynamic maximization problem

$$E_t \sum_{k=0}^{\infty} (\beta\theta)^k \upsilon_{t,t+k} \left[P_t^*(i) X_{t,t+k} - MC_{t,t+k}(i) \right] \frac{Y_{t+k}(i)}{P_{t+k}}, \tag{4}$$

subject to the demand constraints (2) and

$$X_{t,t+k} = \begin{cases} \prod_{l=0}^{k-1} \overline{\pi}^{1-\xi} \pi_{t+l}^{\xi} & \text{for } k > 0 \\ 1 & \text{for } k = 0. \end{cases} \tag{5}$$

with β a constant discount factor, $\upsilon_{t,t+k} = U^{'}(C_t)/U^{'}(C_{t+k})$ the time-varying portion of the discount factor between t and $t+k$; with $U^{'}(C_t)$ the marginal utility of consumption. $\overline{\pi}$ denotes the long-run average gross rate of inflation. Whenever a firm does not re-optimize its price, it resets its price according to an indexation scheme. $\xi \in [0,1]$ measures the degree of indexation to past inflation rates. Note that this partial indexation scheme nests more specific indexation assumptions as special cases.[1]

As shown by Walsh (2003) and Sahuc (2004) aggregate inflation $\hat{\pi}$ can be related to average real marginal cost $\widehat{mc}$ according to

$$\hat{\pi}_t = \gamma_f E_t \hat{\pi}_{t+1} + \gamma_b \hat{\pi}_{t-1} + \lambda \widehat{mc}_t, \tag{6}$$

[1]This specification is adopted from Smets and Wouters's (2003) and Sahuc (2004). With $\xi = 1$ it equals Christiano et al.'s (2005) dynamic indexation scheme, with $\xi = 0$ it simplifies to a purely forward looking model with an indexation to trend inflation.

where

$$\lambda = \frac{(1-\theta\beta)(1-\theta)}{(1+\beta\xi)\theta}\frac{1-\alpha}{1+\alpha(\epsilon-1)},$$
$$\gamma_f = \frac{\beta}{1+\beta\xi},$$
$$\gamma_b = \frac{\xi}{1+\beta\xi}.$$

This inflation equation is known as the *Hybrid New Keynesian Phillips curve.* Its reduced form coefficients γ_f, γ_b and λ are non-linear functions of the structural parameters β, θ, ξ, α and ϵ.[2] When $\xi = 0$ the equation reduces to the pure forward looking New Keynesian Phillips curve. When the cost channel is introduced real marginal cost do not only depend on the labor share of output (as derived by Galí and Gertler, 1999) but also on the nominal interest rate:

$$\widehat{mc}_t = \hat{R}^l_t + \hat{s}_t,$$

where $\hat{s}_t = \hat{w}_t + \hat{n}_t - \hat{y}_t$ is the log deviation of the labor share around the steady state and $\hat{R}^l_t$ is the percentage point deviation of the nominal interest rate (defined as the lending rate) around its steady state value.

In accordance with Chowdhury et al. (2006) it is assumed that the lending rate R^l_t can deviate from the nominal interest rate set by monetary policy that is denoted by R^m_t. This is motivated by financial market imperfects and is for instance motivated by the likelihood of defaults. Profit maximization of financial

[2]Note, that there exist other versions of the structural Phillips curve as well, that have a slightly different interpretation (e.g. Galí and Gertler, 1999). As shown by Scheufele (2008) Galí and Gertler's (1999) model leads to similar conclusions as the one considered here.

intermediaries leads to the following log linear relationship between the risk free rate (which is assumed to be under the control of monetary policy) and the lending rate

$$\hat{R}_t^l = (1 + \psi_R)\hat{R}_t^m \tag{7}$$

where the coefficient $1+\psi_R$ measures the response of the lending rate $\hat{R}_t^l$ to changes in the monetary policy rate $\hat{R}_t^m$. As illustrated by Chowdhury et al. (2006) for $\psi_R > 0$ indicating the existence of strong financial market imperfections. When the opposite holds ($\psi_R < 0$) then managing costs are very high.

Now we can express the Phillips curve as a function of the labor share as well as of monetary policy rate which is given by

$$\hat{\pi}_t = \gamma_f E_t \hat{\pi}_{t+1} + \gamma_b \hat{\pi}_{t-1} + \lambda \hat{s}_t + \lambda \phi_m \hat{R}_t^m, \tag{8}$$

where $\phi_m = (1 + \psi_R)$. So the idea of the cost channel of monetary transmission follows directly from this equation: whenever the central bank raises its interest rate above its steady state level, it leads to an increase of the current inflation rate over its steady state value. This holds true at least unless this effect is not overcompensated by the response of the labor share through adjustments of aggregate demand.

3 Empirical analysis

3.1 Econometric specification

Next we introduce the strategy for estimating the interest-rate augemnted Phillips curve specification and how one can conduct inference about the parameters of interest. Generally, one can choose among two different econometric methods: full information or limited information methods. Choosing among these categories has a long history in econometrics. Full information methods provide the full range of statistical properties associated with the model under investigation and is preferable in terms of efficiency unless the model is correctly specified. Limited information methods do not require a fully specified model instead setting up certain moment conditions is sufficient for estimating the parameters of interest. So there is the classical trade-off between efficiency and the sensitiveness to model mis-specifications known from simultaneous equations models. Since we are interested solely on the Phillips curve equation and more specifically on the direct impact of interest rates on inflation, we find it more naturally to use limited information methods since we do not want to restrict our results on a particular model structure.[3]

Limited information methods typically require the application of instrumental variable (IV) estimation methods. To get an empirical traceable specification from the theoretical model (8) the unobserved variable $E_t\hat{\pi}_{t+1}$ is replace by its realization assuming the forecasting error $\eta_{t+1} = [E_t\hat{\pi}_{t+1} - \hat{\pi}_{t+1}]$ to be orthogonal

[3]Since we also stress the importance of identification robust inference, full information methods like ML are not immune to that kind of problem. However, there are LI methods that are able to deal with this problems. So once weak identification problems show up, ML with its asymptotic theory is generally unreliable and full information methods that are identification robust do not exist.

to past information. So we obtain the estimable equation

$$\hat{\pi}_t = \gamma_f \hat{\pi}_{t+1} + \gamma_b \hat{\pi}_{t-1} + \lambda \hat{s}_t + \lambda \phi_m \hat{R}_t^m + u_t, \qquad (9)$$

where $u_t = \nu_t + \gamma_f \eta_{t+1}$. We allow the error term to follow a very general structure - so u_t may be autocorrelated and / or heteroskedastic.[4] The natural set up for estimating potentially non-linear dynamic model is to employ GMM as proposed by Hansen (1982). With the assumtion $E_{t-1} u_t = 0$ the moment conditions are given by $E_{t-1} \{f_t(\vartheta)\}$, where $f_t(\vartheta) = u_t(\vartheta)\mathbf{z_{t-1}}$ with $\mathbf{z_{t-1}}$ the vector of instruments including predetermined variables dated $t-1$ or earlier. ϑ denotes the parameter vector of interest. For the reduced form model these parameters are given by γ_f, γ_b, λ and ϕ^i. When we are interested in the structural parameters $\vartheta = (\beta, \theta, \xi, \alpha, \epsilon, \phi^i)$.

For parameter estimation we do not consider the convenient two-step GMM (2GMM) estimator that is frequently used for estimating NKPC models (see e.g. Galí and Gertler, 1999; Galí, Gertler and López-Salido, 2001; Eichenbaum and Fisher, 2007). Instead, we stick to the continues updating GMM (CUE) estimator as proposed by Hansen et al. (1996). This estimator is superior in terms of finite sample properties (Hansen et al., 1996; Stock and Wright, 2000). It is more closely related to LIML than 2SLS (as is the 2GMM estimator).[5] Moreover, it does not share the property of standard GMM that estimation bias increases with the inclusion of irrelevant instruments (as documented by Tauchen, 1986; Kocherlakota, 1990). For non-linear settings, another favorable property is its insensitiv-

[4]When we assume ν_t to be white noise then u_t follows a MA(1) process per construction.

[5]For the estimation of a single equation in the linear simultaneous equation model, the two-step GMM estimator is 2SLS whereas the continuous updating estimator is LIML. The superior characteristics of LIML over 2SLS in finite samples has been well documented in the literature (see e.g. Judge, Griffiths, Hill, Luetkepohl and Lee, 1985, Chapter 15).

ity to the statement of the moment conditions. Since the NKPC in its structural formulation is non-linear in its parameters, it is possible to reformulate the orthogonality conditions for instance through multiplying by a certain parameter. 2GMM estimates may be sensitive to this kind of transformation (see Hall, 2005, for a general discussion and Scheufele, 2008, for this problem in the context of the NKPC).

The CUE estimates can be obtained by minimizing the objective function

$$S(\vartheta) = \left[\frac{1}{T} \sum_{t=1}^{T} f_t(\vartheta) \right]' V(\vartheta)^{-1} \left[\frac{1}{T} \sum_{t=1}^{T} f_t(\vartheta) \right], \tag{10}$$

where $V(\vartheta)$ is a $k \times k$ dimensional covariance matrix of the moment vector. This weighting matrix is computed to be heteroscedastic and autocorrelation consistent (HAC) as proposed by Newey and West (1987). The peculiarity of this estimator is that the covariance is estimated together with the parameter vector ϑ. Instead, the 2GMM computes first an initial estimate of ϑ with a pre-specified weighting matrix (e.g. the identity matrix) and then uses this initial estimate to specify the weighting matrix in the second step.

To estimate the structural form parameters it is necessary to calibrate some parameters since with four variables one can at least identify the same number of parameters. We follow Galí et al. (2001) and choose to calibrate α and ϵ (in accordance with them, we set $\alpha = 0.270$ and $\epsilon = 11$ for the US and $\alpha = 0.175$ and $\epsilon = 11$ for the Euro area). Given these values one can compute the point estimates for β, θ, ξ and ϕ_i.

For conducting inference of the parameters of interest we do not rely on standard Wald-type tests and t-statistics as it is usually done in the standard GMM framework. Instead, we report weak instrument robust confidence intervals for the parameters of the NKPC. As it was early stressed (see Pesaran, 1987, Ch. 6 and 7) the identification issue for forward looking rational expectations models is very important to consider otherwise estimation results get unreliable. In his work, he recommends to pre check the conditions necessary to guarantee identification of the parameters. In contrast, this study takes a different perspective since problems for inference even arise when the parameters are close to be unidentified. This situation is often described as weak identification and is directly connected with the problem of weak instruments.[6] Stock, Wright and Yogo (2002), Dufour (2003) and Andrews and Stock (2005) provide excellent surveys for this situation and discuss possible solutions for this pathology. To put it simple, weak instrument problems arise where the correlation between the right hand side endogenous variables and their instruments is relatively small. In this situation, Wald tests are unreliable since they do not provide the exact test size.

As shown by Mavroeidis (2005), identification of the NKPC for economic plausible parameter values is challenging and weak instrument problems are very likely to occur. This view is supported empirically by a multiplicity of studies (Ma, 2002; Dufour et al., 2006; Mavroeidis, 2006; Martins and Gabriel, 2006; Kleibergen and Mavroeidis, 2008b; Nason and Smith, 2008) by comparing weak instrument robust tests with standard Wald type tests obtained with GMM. We basically follow their idea with not imposing a priori the assumption that the parameter are identified. We used S-sets as proposed by Stock and Wright (2000) and applied to the NKPC by Ma (2002) and Mavroeidis (2006) that can be constructed from the CUE ob-

[6]Note that this problem is not specific to GMM estimation. Instead, ML-methods or other Matching Moments methods can be also affected by this pathology (see e.g. Canova and Sala, 2006, for a general overview for DSGE models)

jective function (see eq 10). This method shares important characteristics that are outlined by Dufour (1997) as identification robust. This requires that confidence intervals should be unbounded (and thus uninformative) whenever parameters are unidentified. In the situation where parameters are weakly identified this should translate into confidence sets that are fairly large. As shown by Dufour (1997) this is not the case by standard Wald-type methods that only hold when identification is fully guaranteed and when no weak instrument problems are present; otherwise these methods are unreliable and standard normal approximations provide a very poor guide for inference.

The S-sets used for constructing confidence sets are very close to the well-known overidentification test of Anderson and Rubin (1949). Several authors (including e.g. Dufour, 1997; Stock et al., 2002; Dufour, 2003; Andrews and Stock, 2005; Dufour and Taamouti, 2005; Dufour and Taamouti, 2007) provide evidence that this static is fully robust to weak instrument problems. In linear simultaneous equation models Stock and Wright (2000) have shown that S-sets are asymptotically equivalent to confidence sets obtained by inverting the Anderson-Rubin (AR) statistic. So S-sets can be seen as an extension for the AR test in linear models to GMM as a more general model class. To obtain S-sets that is a joint confidence set for the parameter vector ϑ we use Stock and Wright's (2000) result that $S(\vartheta_0) \xrightarrow{D} \chi^2_k$, where $S(\vartheta_0)$ is the CUE objective function (eq 10) evaluated at the true parameter values ϑ_0 and k is the number of instruments. The joint confidence interval consists of those parameter values for which the test statistic do not reject.[7] This procedure can be applied both to the reduced form parameters γ_f, γ_b, λ and ϕ^i

[7]The construction of joint confidence intervals involves searching for values within a economically plausible range and collecting those values for which the test do not reject.

and to the structural parameters β, θ, ξ and ϕ^i (given the calibrated values for α and ϵ). The resulting S-set is four-dimension for the full model specification.[8]

Confidence intervals for the individual parameters are obtained by using the projection method. The idea is that projection based tests do not reject the individual hypotheses $H_0 : \beta = \beta_0$ when the joint hypothesis $H^* : \beta = \beta_0, \alpha = \alpha_0$ do not reject for some values of α. This test method is proposed by Dufour (1997), Dufour and Jasiak (2001), Dufour and Taamouti (2005) and Dufour and Taamouti (2007). This procedure is fully robust to weak instruments but it has the drawback that projection-based tests are conservative.[9]

A further characteristic of identification robust confidence intervals based on the CUE objective function is that they may be empty. This is the case when the test rejects for all possible parameter values. Thus, S-sets already include a test of overidentified restrictions comparable to a J test as proposed by Hansen. If no parameter vector is compatible with the specified model the corresponding confidence sets will be empty. We interpret this results as a rejection of the empirical model.

[8]There are now additional methods available dealing with weak instrument problems within the GMM setting (see Kleibergen and Mavroeidis, 2008b, for a comparison of different IV robust methods with an application to the Phillips curve). Kleibergen and Mavroeidis (2008b) consider not only S-sets, but also a score Lagrange Multiplier (KLM) test, the difference between S-sets and the KLM statistics (JKLM) and an extension of the conditional likelihood ration test of Moreira (2003) to GMM (MQLR). Their simulation results indicate that the MQLR is at least as powerful as any of the other tests. However, while MQLR dominates the S statistics under some conditions in terms of power, it also imposes additional restrictions on the reduced form models and may be thus more fragile. This may translate into problems when relevant instruments are missing (this point was raised by Dufour, 2008).

[9]Another approach is available that can be applied to parameter subsets (see Stock and Wright, 2000; Kleibergen and Mavroeidis, 2008a). In this case some parameters are assumed to be identified. As long as the assumption are satisfied, these tests are asymptotically non-conservative and are then more powerful then projection based tests. But this method is only partially robust to weak instruments and may break down once the assumed identified parameters turn out to be weakly identified.

3.2 Data

We use quarterly time series data to estimate Eq (9) for the US and the Euro area. For the US we consider a sample period ranging from 1960q1-2005q4; for the Euro area the estimation period is 1972q1-2005q4. The US data are taken from the OECD Quarterly National Accounts database and the IMFs International Financial Statistics (IFS). Whereas for the Euro area we employ the data set of the Area Wide Model (AWM). Inflation is defined as the quarterly log difference of the GDP deflator. Real marginal cost is proxied by the labor share of output which is defined as the ratio of total compensation to nominal GDP. As a measure for the short-run nominal interest rate two definitions are considered: 3-month treasury bill rates and bank lending rates.[10] Both explanatory variables – labor share and interest rates – are defined as percentage deviations of a steady state value while inflation rate are expressed as percentage point deviations.[11]

The potential instrument set is composed of lags of inflation, the labor share and short term interest rates (up to four lags). Additional instruments consist of a yield spread, $(r^l - r^m)_t$, defined as the 10-year government bond yield minus the 3-month Treasury bill rate, wage inflation Δw_t and a quasi-real time detrended output gap $\tilde{y}_t$ (which is computed recursively and contains only information up to period t). These additional instruments are followed up to two lags. To get

[10] Bank lending rates for the US are taken from the IMFs International Financial Statistics (IFS). Whereas for the Euro area we use the constructed series from Calza, Manrique and Sousa (2003) which is available from 1980q1-2001q3.

[11] For the US we assume constant steady state values for inflation, the labor share and the nominal interest rate. For the Euro area we assume a broken linear time trend for inflation (where a falling trend is assumed until 1998.4, afterwards the steady state is assumed to be constant). For the the labor share a linear falling trend is presumed and for the nominal interest rates three different periods can be distunguished: a constant nominal rate until 1979, then a falling trend until 1998 and since 1999 again a constant steady state value for the interets rate.

rid of redundant instruments (those which are only marginally correlated with the variables they instrument) we estimate a three dimensional system similar to a VAR consisting of inflation $\hat{\pi}_{t+1}$, the labor share $\hat{s}_t$ and interest rates $\hat{R}_t$ as left hand side endogenous variables. These variables are regressed on their own lags and on the lags of the remaining endogenous variables. The additional instrumental variables are treated as supplemental exogenous variables (with one and two lags). After estimating the full model a model reduction procedure is applied that sequentially eliminates blocks of regressors. This procedure is based upon the Schwarz criterion and stops when the lowest value for the information criterion is obtained.[12] The exclusion of redundant instruments can avoid the well documented power loss of the AR statistic when the number instruments gets large (Andrews and Stock, 2005).

After applying this reduction technique, the instrument sets consist of

$$\mathbf{z}_{t-1}^{sc,us} = \left[c \ \hat{\pi}_{t-1} \ \hat{\pi}_{t-2} \ \hat{\pi}_{t-3} \ \hat{s}_{t-1} \ \hat{R}_{t-1}^{i} \ \hat{R}_{t-2}^{i} \ \hat{R}_{t-3}^{i} \ \hat{R}_{t-4}^{i} \ \tilde{y}_{t-1} \ (r^{l} - r^{m})_{t-1}\right]' \tag{11}$$

for the US and

$$\mathbf{z}_{t-1}^{sc,eu} = \left[c \ \hat{\pi}_{t-1} \ \hat{\pi}_{t-2} \ \hat{s}_{t-1} \ \hat{s}_{t-3} \ \hat{R}_{t-1}^{i} \ \hat{R}_{t-3}^{i} \ \Delta w_{t-1} \ \Delta w_{t-2}\right]' \tag{12}$$

for the Euro area; where c denotes a constant term. Overall, we consider ten and eight instrumental variables (plus a constant) for the US and the Euro area, respectively.

[12]See Scheufele (2008) for an application of this procedure to the standard NKPC.

4 Estimation results

In the following we present the results of the identification robust estimation and test procedure. First, we estimate the linear, reduced form version of the interest rate augmented hybrid Phillips curve similar to Chowdhury et al.'s (2006) specification. We start by considering the most general specification and then test whether further parameter restrictions can be imposed. Besides the treasury bill rate, we also use the bank lending rates as an interest measure relevant for the cost channel. Estimates are provided for the US and the Euro Area. Finally, we estimate the structural version of the Phillips curve including the cost channel, where a generalization of Ravenna and Walsh's (2006) model is presented. All variables included in the Phillips curve are taken as deviations from steady state values (which are given by sample means and deterministic time trends). In addition, a Newey-West heterostedasticity and autocorrelation-consistent (HAC) covariance estimator is used for all specifications, with a lag length of 5. For the construction of S-sets a ten percent test level is presumed (this coincides with Stock and Wright, 2000).

4.1 Reduced form estimates

Table 1 reports the estimation and test results for the reduced form parameters of eq 9 for the US. This model is linear both in parameters and in variables. When the treasury bill rate is taken as measure for short run interest rates the parameter estimates for $(\gamma_f, \gamma_b, \lambda, \lambda\phi^m)$ are given by $(0.60, 0.34, 0.04, 0.03)$. The estimates for γ_f, γ_f and λ are basically in line with those reported in the literature on specifications without considering the cost channel (see e.g. Galí and Gertler, 1999).

More specifically, these results indicate that inflation is mainly forward looking. The backward looking coefficient is smaller and turns out to be insignificant (at least for the full model specification). The coefficient of the labor share as well as the coefficient of the interest rates are both insignificant. This finding also holds true even when further restrictions are imposed, i.e. when $\phi^m = 1$ and $\gamma_b = 1 - \gamma_f$ is assumed. Given this specification, the labor share does not seem to be the driving variables for inflation dynamics since this variable mostly turns out to be unimportant. Only under the restriction $\phi^i = 1$ and $\gamma_b = 0$ the slope of the Phillips curve is significant. This also implies that the cost channel does matter for explaining inflation dynamics. This result is in line with Ravenna and Walsh (2006). However, for the most general hybrid form we cannot confirm Chowdhury et al.'s (2006) results of a significant cost channel and as well as a significant labor share in the hybrid specification. However, their results are based upon a standard 2GMM estimation strategy using asymptotic theory for conducting inference that is not robust against weak instrument problems. When we take the lending rate as interest rate, the results hardly change and basic findings still hold.

For the Euro area the point estimates are similar to the ones for the US. However, in the full model all regressors turn out to be insignificant. The obtained confidence intervals are very wide and always include zero. In this case, the typical weak instrument pathologies show up and confidence intervals are so large that no economic meaningful implications of the model can be obtained. When restrictions are imposed ($\phi^m = 1$ and $\gamma_b = 1 - \gamma_f$) the Phillips curve for the Euro Area is again compatible with an important forward looking element, whereas the backward looking part seems to be of minor importance. As opposed to the US, there is never a significant role for the marginal cost variables – irrespectively whether

Table 1: Reduced form estimates for the US

	Restrictions	γ_f	γ_b	λ	$\lambda\phi^i$	p-value
I. Interest rate measure: treasury bill rate						
			A. CHS specification:			
(1)	no	0.6040	0.3414	0.0434	0.0303	0.9827
		[0.35,1.00]	[0.00,0.55]	[-0.10,0.23]	[-0.04,0.14]	
(2)	$\gamma_b = 1 - \gamma_f$	0.6927	0.3073	0.0028	0.0173	0.7501
		[0.50,1.00]	–	[-0.10,0.12]	[-0.04,0.10]	
(3)	$\phi = 1$	0.6131	0.3348	0.0344	0.0344	0.9804
		[0.35,0.95]	[0.00,0.55]	[-0.03,0.14]	–	
			B. Standard hybrid Specification:			
(4)	$\phi = 0$	0.5896	0.3701	0.0543	0	0.8810
		[0.35,0.95]	[0.05,0.55]	[-0.07,0.22]	–	
(5)	$\phi = 0$,	0.6714	0.3286	0.0170	0	0.7395
	$\gamma_b = 1 - \gamma_f$	[0.50,0.95]	–	[-0.05,0.11]	–	
			C. Pure forward looking Specifications:			
(6)	$\gamma_b = 0$	0.9643	0.0000	0.0131	0.0630	0.1402
		[0.95,1.00]	–	[-0.05,0.07]	[0.04,0.10]	
(7)	$\gamma_b = 0$,	0.9549	0.0000	0.0487	0.0487	0.1219
	$\phi = 1$	[0.95,0.96]	–	[0.04,0.06]	–	
(8)	$\gamma_b = 0$, $\phi = 0$		empty set			0.0643
II. Interest rate measure: Bank lending rate						
			A. CHS specification:			
(1)	no	0.6218	0.3317	0.0361	0.0261	0.9915
		[0.35,1.00]	[0.00,0.55]	[-0.12,0.23]	[-0.04,0.12]	
(2)	$\gamma_b = 1 - \gamma_f$	0.6995	0.3005	-0.005	0.0237	0.8298
		[0.50,1.00]		[-0.12,0.12]	[-0.04,0.10]	
(3)	$\phi = 1$	0.6300	0.3259	0.0290	0.0290	0.9908
		[0.40,0.95]	[0.00,0.55]	[-0.02,0.11]	–	
			B. Standard hybrid Specification:			
(4)	$\phi = 0$	0.5854	0.3725	0.0565	0	0.9195
		[0.35,0.90]	[0.10,0.55]	[-0.06,0.23]	–	
(5)	$\phi = 0$,	0.6628	0.3372	0.0179	0	0.6826
	$\gamma_b = 1 - \gamma_f$	[0.50,0.90]	–	[-0.05,0.11]	–	
			C. Pure forward looking Specifications:			
(6)	$\gamma_b = 0$	0.9674	0	-0.0026	0.0678	0.1825
		[0.90,1.00]	–	[-0.08,0.07]	[0.04,0.10]	
(7)	$\gamma_b = 0$,	0.9495	0	0.0505	0.0505	0.1398
	$\phi = 1$	[0.94,0.95]	–	[0.04,0.06]	–	
(8)	$\gamma_b = 0$, $\phi = 0$		empty set			0.0449

Notes: Point estimates are obtained using CUE. Projection based confidence intervals in squared brackets. P-values report the test for the joint confidence set evaluated at the CUE point estimates. A 5-lag Newey-West HAC estimate is used. Sample period: 1960:1-2005:4. SC based instrument set (see eq 11).

Table 2: Reduced form estimates for the Euro area

	Restrictions	γ_f	γ_b	λ	$\lambda\phi^m$	p-value
			A. CHS specification:			
(1)	no	0.8056	0.1961	-0.0281	0.0563	0.8847
		[-0.40,2.00]	[-0.40,0.40]	[-0.37,0.14]	[-0.06,0.78]	
(2)	$\gamma_b = 1 - \gamma_f$	0.8038	0.1962	-0.0276	0.0561	0.8847
		[0.52,1.35]	–	[-0.26,0.09]	[-0.5,0.28]	
(3)	$\phi = 1$	0.7188	0.2077	0.0217	0.0217	0.7533
		[0.25,1.45]	[-0.25,0.45]	[-0.06,0.12]	–	
			B. Standard hybrid Specification:			
(4)	$\phi = 0$	0.7267	0.2296	0.0115	0	0.6115
		[0.30,1.70]	[-0.20,0.40]	[-0.25,0.14]		
(5)	$\phi = 0$,	0.7767	0.2233	-0.0006	0	0.5962
	$\gamma_b = 1 - \gamma_f$	[0.55,1.20]	–	[-0.11,0.09]		
			C. Pure forward looking Specifications:			
(6)	$\gamma_b = 0$	1.0317	0	-0.0375	0.0798	0.6900
		[0.80,1.80]	–	[-0.37,0.12]	[-0.05,0.28]	
(7)	$\gamma_b = 0$,	0.9363	0	0.0315	0.0315	0.4961
	$\phi = 1$	[0.64,1.40]	–	[-0.05,0.10]	–	
(8)	$\gamma_b = 0$, $\phi = 0$	1.019	0	0.0057	0	0.3484
		[0.75,1.65]	–	[-0.22,0.12]		

Notes: See Table 1.

further restrictions are imposed. Here, the cost channel (as well as the labor share) does not turn out as a significant driver of inflation.

4.2 Structural estimates

More interesting than the linear Phillips curve models is the specification where the structural form parameters are estimated directly. Given plausible values for ϵ and α direct estimates of the deep parameters – β the discount factor, θ the Calvo parameter that measures the degree of nominal price rigidities, ξ the degree of price indexation of firms that do not re-optimize as well as ϕ^m indicating the relevance of the cost channel – can be obtained. For the full model we get an estimated parameter vector of (β , θ, ξ, ϕ^m) equal to $(0.85, 0.61, 0.48, 0.70)$. The estimates values for β, θ and ξ are more or less comparable with other studies (Galí et al., 2001; Sahuc, 2004; Eichenbaum and Fisher, 2007).

Table 3 reports the results for the structural parameters of the interest rate augmented Phillips curve. Estimates for the full model (1) confirms the findings for the reduced form parameters. The point estimate of the subjective discount factor β is 0.85 and the corresponding confidence intervals are within an economically plausible range near one. The parameter θ, that can be interpreted as a measure of nominal price rigidity, is positive and significant. The estimated parameter $\theta = 0.61$ translates into an average frequency of price re-optimization of firms between 2 and 3 quarters. This is in line with other empirical studies for the US that do not consider the role of a cost channel (see e.g. Galí et al., 2001; Eichenbaum and Fisher, 2007). However, the null hypothesis that $\theta = 1$ cannot be rejected which implies that the model is also consistent with perfect price rigidity where firms never re-optimize their prices. This also translates into a slope coefficient of the marginal cost variable $\lambda = 0$ that is consistent with the results based on the reduced form parameters. Estimates for the degree of indexation ξ is complete uninformative - so one cannot decide whether full indexation $\xi = 1$ (indexation to past inflation rates) or no indexation $\xi = 0$ (which is equivalent to an indexation scheme to trend inflation) applies. For the additional variable ϕ^m that measures the impact of the cost channel no meaningful conclusions can be obtained. The confidence set of this parameter is completely uninformative and include the hole parameter space. So it is easy to see that weak identification is an issue for the interest rate augmented Phillips curve, in particular when the structural parameters are of interest.

In a second specification the model is restricted to the pure forward looking

Table 3: Structural Estimates for the US

	Restrictions	β	θ	ξ	ϕ	Freq.	p-value
I. Interest rate measure: treasury bill rate							
			A. Full model specification:				
(1)	no	0.8517	0.6083	0.4815	0.6980	2.5530	0.9827
		[0.50,1.00]	[0.37,1.00]	[0.00,1.00]	[$-\infty$,∞]	[1.59,∞]	
(2)	$\phi = 1$	0.8618	0.6445	0.4705	1.0000	2.8129	0.9804
		[0.43,1.00]	[0.45,1.00]	[0.00,1.00]	–	[1.82,∞]	
(3)	$\phi = 0$	0.8693	0.5588	0.5457	0	2.2665	0.8810
		[0.45,1.00]	[0.35,1.00]	[0.05,1.00]	–	[1.54,∞]	
			B. Full indexation scheme:				
(4)	$\xi = 1$	1.0000	0.7500	1	0.0000	4.0000	0.1319
		[0.96,1.00]	[0.59,1.00]	–	[$-\infty$,∞]	[2.43,∞]	
(5)	$\xi = 1$, $\phi = 1$	1.0000	0.9100	1	1	11.1111	0.1263
		[1.00,1.00]	[0.82,1.00]	–	–	[5.56,∞]	
(6)	$\xi = 1$, $\phi = 0$	1.0000	0.7300	1	0	3.7037	0.1316
		[1.00,1.00]	[0.64,1.00]	–	–	[2.78,∞]	
			C. RW specification (static indexation):				
(7)	$\xi = 0$	0.9643	0.7858	0	4.8401	4.6685	0.1402
		[0.91,1.00]	[0.59,0.99]	–	[0.60,992.0]	[2.44,100]	
(8)	$\xi = 0$, $\phi = 1$	0.9548	0.6220	0	1	2.6455	0.1219
		[0.92,1.00]	[0.57,0.71]	–	–	[2.33,3.45]	
(9)	$\xi = 0$, $\phi = 0$			empty set			0.0643
II. Interest rate measure: Bank lending rate							
			A. Full model specification:				
(1)	no	0.8769	0.6327	0.4677	0.7228	2.7226	0.9915
		[0.50,1.00]	[0.37,1.00]	[0.00,1.00]	[$-\infty$,∞]	[1.59,∞]	
(2)	$\phi = 1$	0.8855	0.6654	0.4582	1	2.9886	0.9908
		[0.48,1.00]	[0.53,1.00]	[0.00,1.00]	–	[1.82,∞]	
(3)	$\phi = 0$	0.8625	0.5535	0.5486	0	2.2396	0.9195
		[0.45,1.00]	[0.34,1.00]	[0.10,1.00]	–	[1.51,∞]	
			B. Full indexation scheme:				
(4)	$\xi = 1$	1.0000	0.7500	1	0.0000	4.0000	0.1482
		[0.96,1.00]	[0.57,1.00]	–	[$-\infty$,∞]	[2.33,∞]	
(5)	$\xi = 1$, $\phi = 1$	1.0000	0.9100	1	1	11.1111	0.1263
		[1.00,1.00]	[0.82,1.00]	–	–	[5.56,∞]	
(6)	$\xi = 1$, $\phi = 0$	1.0000	0.7300	1	0	3.7037	0.1479
		[1.00,1.00]	[0.64,1.00]	–	–	[2.78,∞]	
			C. RW specification (static indexation):				
(7)	$\xi = 0$	0.9650	0.9850	0	455.00	66.6667	0.1823
		[0.90,1.00]	[0.54,0.99]	–	[0.60,2000]	[2.17,200]	
(8)	$\xi = 0$, $\phi = 1$	0.9496	0.6175	0	1	2.6144	0.1398
		[0.91,1.00]	[0.56,0.70]	–	–	[2.27,3.33]	
(9)	$\xi = 0$, $\phi = 0$			empty set			0.0449

Notes: Using the calibrated values $\alpha = 0.27$ and $\epsilon = 11$. See also Table 1.

specification where firms that do not re-optimize are assumed to index their prices to trend inflation, so $\xi = 0$. This also corresponds to the specification of Ravenna and Walsh (2006) who estimate their model with 2GMM. In this case a significant coefficient for the nominal interest rate can be obtained with $\phi^m = 4.8$. This point estimate is roughly in the middle of the parameter estimates obtained by Ravenna and Walsh (2006), who find quite different estimates depending on how the orthogonality conditions are normalized. Since we use CUE, our results are not sensitive to that kind of problem. But, we get a confidence interval that is again very large. Another remarkable result for this specification is that the estimate for θ gets more precise. Now, full price rigidity can be rejected and marginal cost seems to be a relevant source for inflation dynamics (once the cost channel is considered). When the parameter ϕ^m is set equal to one, which implies that ψ is assumed to be equal to zero. For this model, confidence sets for the remaining parameters turn out to be small and well in line with theoretical aspects. When the interest rate is completely omitted from the equation ($\phi^m = 0$), the model is totally rejected. No parameter is compatible with this model. This can be interpreted as a omitted variable test for the cost channel. Since the P-value is now below 10% the confidence sets are empty. Taken together, once the pure forward looking specification is considered, there is indeed evidence for a cost channel that is supported by our structural estimation and test methodology. So, for the US we can confirm the results of Ravenna and Walsh (2006) and Tillmann (2008) with our robust structural approach.

Using the lendig rate as adequate interest rate measure hardly changes the results. Instead the previous results are confirmed and a pure foreward looking model including the cost channel seems most faviourable.

Table 4: Structural estimates for the Euro area

	Restrictions	β	θ	ξ	ϕ	Freq.	p-value
			I. Interest rate measure: treasury bill rate				
			A. Full model specification:				
(1)	no	0.9305	0.9957	0.2373	572.4373	232.5581	0.8550
		[0.00,1.00]	[0.64,1.00]	[0.00,0.90]	$[-\infty,\infty]$	$[2.78,\infty]$	
(2)	$\phi = 1$	0.8794	0.7803	0.2540	1.0000	4.5517	0.7533
		[0.30,1.00]	[0.64,1.00]	[0.00,0.90]	–	$[2.78,\infty]$	
(3)	$\phi = 0$	0.9218	0.8291	0.2912	0	5.8514	0.6115
		[0.35,1.00]	[0.55,1.00]	[0.00,0.90]	–	$[2.22,\infty]$	
			B. Full indexation scheme:				
(4)	$\xi = 1$			empty set			0.0665
(5)	$\xi = 1$, $\phi = 1$			empty set			0.0639
(6)	$\xi = 1$, $\phi = 0$			empty set			0.0639
			C. RW specification (static indexation):				
(7)	$\xi = 0$	0.9733	0.9930	0	910.5027	142.8571429	0.6456
		[0.70,1.00]	[0.57,1.00]	–	$[-\infty,\infty]$	$[2.33,\infty]$	
(8)	$\xi = 0$, $\phi = 1$	0.9364	0.7435	0	1	3.8986	0.4961
		[0.63,1.00]	[0.60,1.00]	–	–	$[2.50,\infty]$	
(9)	$\xi = 0$, $\phi = 0$	1.0000	0.8634	0	0	7.3206	0.3484
		[0.71,1.00]	[0.57,1.00]	–	–	$[2.33,\infty]$	

Notes: Using the calibrated values $\alpha = 0.27$ and $\epsilon = 11$. See also Table 1.

4.3 Summary

In sum, this analysis discovers several interesting findings. First, the results confirm that weak identification is an issue in estimating the New Keynesian Phillips curve as often discussed in the literature (it also holds true for the interest rate augmented version). So identification-robust methods can help to determine reliable inference of the parameters of interest. Second, it turns out that it is very difficult to discriminate between forward-looking and backward-looking behavior empirically once identification robust methods are applied (this finding is consistent with Mavroeidis, 2006; Kleibergen and Mavroeidis, 2008b). Third, when a pure forward looking specification is employed, the point estimates for the coefficient on the nominal interest rate is significant for the US and seems to impact firms marginal costs. However, it is very difficult to pin down the exact magnitude of the cost channel effect since confidence intervals for the parameter ϕ^m are extremely wide. So the parameter ψ_R is also difficult to determine. However, our results are consistent with $\psi_R > 0$ which indicates that monetary policy rate effects on firms' costs for working capital (but also a one-to-one relationship cannot be rejected) are amplified. This points to the existence of considerable financial market imperfections. Finally, these results reveal that the cost channel matter for inflation dynamics. So marginal costs do not solely depend on real wages relative to marginal productivity (measured by the labor share) but also on the nominal interest rate. The exclusion of the cost channel leads to a statistical rejection of the model when a pure forward looking Phillips curve is specified. Given these results, important consequences for monetary policy emerge (see e.g. Chowdhury et al., 2006; Ravenna and Walsh, 2006). In general, the empirical evidence of the cost channel suggests that it has to be included into new Keynesian Models for monetary policy analysis, at least for the US.

5 Conclusions

This paper investigates whether interest rate changes have an impact on firms marginal costs and whether this has a direct effect on the price setting behavior of firms that translates into aggregate inflation dynamics. Empirical tests of the existence of this cost channel effect using a structural approach are employed. Estimation and inference is conducted using identification robust methods based on the continues-updating objective function. We document identification difficulties for some parameters when estimating the general model structure. For the US, a pure forward-looking interest rate augmented Phillips curve is indeed compatible with data. This suggests that considering the cost channel is a non-negligible aspect for monetary policy.

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