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Abstract

Evidence from vector autoregressions indicates that the impact of interest rate shocks on macroeconomic aggregates can substantially be affected by the so-called cost channel of monetary transmission. In this paper we apply a structural approach to examine the relevance of the cost channel for inflation dynamics in G7 countries. Since firms' costs of working capital increase with interest rates, we augment a (hybrid) New Keynesian Phillips curve by including the short-run nominal interest rate. We find significant and varying direct interest rate effects for the majority of countries, including member countries of the EMU. Simulations further demonstrate that the estimated interest rate coefficients can substantially affect inflation responses to monetary policy shocks, and can even lead to inverse inflation responses, when the cost channel is – relative to the demand channel – sufficiently strong.

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1 Introduction

This paper aims at revealing whether changes in short-run nominal interest rates, which alter costs of working capital, affect pricing decisions of firms and thus matter for inflation dynamics in industrialized countries. Supply side effects of nominal interest rates are already considered in various studies focussing on the transmission of monetary policy shocks, i.e., in the literature on the so-called 'credit channel' (see Bernanke and Gertler, 1995). Their impact, however, on inflation dynamics within the New Keynesian framework, which by now serves as the predominant framework for monetary policy analysis, has rarely been taken into account. Recent empirical contributions to the literature on the New Keynesian Phillips curve presume that price rigidities are the main source for monetary non-neutrality (see e.g., Galí and Gertler, 1999; Galí et al., 2001; Sbordone, 2002; Benigno and López-Salido, 2002). According to this view, monetary policy actions affect inflation dynamics via changes in firms' real marginal costs, solely brought about by shifts in aggregate demand.

Though this approach leads to predictions about price responses to interest rate shocks, which accord to common priors about monetary policy effects, they are not fully consistent with vector autoregression (VAR) based evidence (see Christiano et al., 1999). Thus, it seems that this approach to inflation dynamics does not account for all major aspects of monetary transmission. As shown by Barth and Ramey (2001) for the US, the impact of interest rate shocks on prices and real activity is significantly affected by changes in the cost of working capital.[2] Given that higher nominal interest rates directly raise firms' costs of working capital, the cost alleviating effect of a decline in aggregate demand is counteracted. As a consequence, the inflation response to an interest rate shock is mitigated and the output response is amplified by this "cost-channel of monetary transmission". While the costs of working capital evidently rise with the nominal interest rate, the apparent question is, whether the impact on macroeconomic aggregates and, in particular, on inflation dynamics is in fact of measurable size. Addressing this question, among others, Christiano et al.'s (2004) empirical assessment of a large scale dynamic general equilibrium model, which incorporates a marginal cost based Phillips curve, indicates that the cost channel is non-negligible for the transmission of monetary policy shocks in the US.

In this paper, we provide further (structural) evidence on the cost channel by estimating marginal cost based Phillips curves that account for direct interest rate effects, and show that changes in short-run nominal interest rates exert a substantial direct effect on inflation dynamics in the majority of G7 countries. We build on the evidence by Galí and Gertler (1999) and Galí et al. (2001) and allow for backward-looking elements in the price setting behavior of firms, and estimate a hybrid version of a marginal cost based Phillips curve for the time period 1980-1997. Thereby, we find that changes in short-run nominal interest rates significantly affect the short-run movements of inflation rates in Canada, France, Italy, the UK and the US,[3] while we could not establish a significant cost channel in Germany and

[2] Further empirical evidence on the relevance of working capital for the transmission of interest rate shocks in France, Germany, Italy and in the UK is provided by Dedola and Lippi (2003).

[3] Ravenna and Walsh (2003) provide similar results for a marginal cost based Phillips curve, which is restricted to be entirely forward looking, indicating a significant cost channel in the US.

Japan. The existence and strength of an effective cost channel seems to vary in accordance with differences in financial systems, as summarized in Allen and Gale (2000, 2004). In particular, our results suggest that the cost channel of monetary transmission is hardly effective when the financial intermediary sector is highly regulated and less competitive, as in Germany and Japan. Correspondingly, a high degree of financial market liberalization, such as in the US or in the UK, is associated with an immediate pass-through of changes in the monetary policy rate to the costs of working capital. For these countries, our estimates further imply that firms' marginal costs are raised by more than one for one with changes in the monetary policy rate, indicating the existence of financial market frictions that accelerate the cost channel effects.

In the last part of the paper, we illustrate the impact of the cost channel on inflation responses to monetary policy shocks, by integrating the interest-rate-augmented Phillips curve in a general equilibrium framework. Interest rate shocks raise the costs of working capital and reduce unit labor costs by a decline in aggregate demand. As a consequence, the direct interest rate effect on firms' marginal costs can drive a wedge between the responses of output and inflation to a monetary policy tightening. Applying the estimated parameter values for the aggregate supply relation, our simulation results indicate that the cost channel substantially alters the inflation path. Moreover, when direct interest rate effects are sufficiently strong – compared to the demand channel – the inflation response to an interest rate hike can even be positive, which, as stressed by Barth and Ramey (2001) might serve as an explanation for the so-called "price puzzle", often found in monetary VARs (see Christiano et al., 1999; Hanson, 2004). Our results imply that a significant cost channel, which is found to vary substantially between G7 countries, can serve as a major source for differences in the transmission of monetary policy shocks. In particular, heterogenous financial structures and different degrees of interest rate pass-through between countries can lead to asymmetric inflation responses to interest rate shocks. In our analysis we excluded the recent past (due to data availability), where financial market heterogeneities in EMU member countries have been reduced, but still exist (see ECB, 2003). Thus, differences in financial structures and, therefore, in direct interest rate effects, as disclosed in this paper, are likely to impose a burden for the conduct for monetary policy in the EMU.

The remainder is set out as follows. Section 2 provides a structural description of the pricing behavior of firms that rely on working capital. Empirical evidence for an interest-rate-augmented Phillips curve is presented in the first part of Section 3. In the second part of Section 3 we assess the robustness of direct interest rate effects on inflation via the cost channel. Section 4 discusses the effects of interest rate shocks on inflation. Section 5 concludes.

2 Working capital, marginal costs, and inflation

In this section we develop a simple theoretical framework, which provides a structural description for the firms' price setting behavior that is based on the specification developed in New Keynesian macroeconomics (see Goodfriend and King, 1997; Clarida et al., 2000). In our specification, we explicitly take into account that firms face liquidity constraints, which cause them to hold working capital, that is defined as the difference between current assets and current liabilities. Higher interest rates raise the opportunity costs of working capital and, therefore, the costs of production. As a consequence, a monetary tightening exerts pressure on firms' costs due to a *supply effect*, which accompanies the cost alleviating effect of a decline in aggregate demand, i.e., the conventional New Keynesian *demand effect* (see Goodfriend and King, 1997). In equilibrium, firms' price adjustments are therefore jointly determined by both effects.

To induce firms' to hold working capital, we follow Christiano et al.'s (1997) approach and introduce a liquidity constraint for firms in the factor markets. According to this assumption, firms have to pay for production factors before the goods market opens. Firms, therefore, borrow funds to finance their outlays for production inputs, such that the interest rate on external funds raises the marginal costs of production. While this approach focuses on the costs of external funds, the logic of interest rate effects on firms' costs also applies when firms are primarily financed by internal funds, as stressed by Barth and Ramey (2001). Thus, direct cost effects of interest rates are not particular to economies with a high ratio of external to internal funds, as the opportunity costs of working capital, i.e., net current assets, increase with the interest rate regardless whether funds are internally or externally generated. Moreover, interest rate effects on firms' costs are likely to be accelerated by adverse effects on firms' balance sheets and on their net worth, as pointed out in the literature on the "credit channel" or "financial accelerator" (see Bernanke and Gertler, 1995; Bernanke et al., 1999). Thus, a specification that solely considers interest rate payments on debt is likely to underestimate direct interest rate effects on firms' costs. We, therefore, allow for a friction in the financial intermediary sector, which is not explicitly derived for simplicity.

The details of the firms' problem unfold as follows. There is a continuum of monopolistically competitive firms indexed with $i \in [0,1]$, which are owned by households. Firm i produces differentiated goods y_{it} with the production technology: $y_{it} = a_t l_{it}^{1-\alpha} x_{it}^{\alpha}$, where l_{it} is the firm specific labor input and a_t denotes the productivity level. The second production factor x_{it} denotes raw materials or commodities that are owned by the households.[4] In order to hire labor and to purchase commodities, we assume that firms have to pay outlays for wages and for commodities in advance, i.e., before production takes place. Put differently, they face a liquidity constraint on the factor markets, such that production relies on a sufficiently large amount of liquid funds. Given that profits are transferred at the end of each period to their owners, firms rely on external funds to meet this liquidity constraint. In particular, firm i is assumed to borrow the amount Z_{it} from financial intermediaries before it enters the factor markets, in order to meet the following liquidity constraint:

$$Z_{it} \geq P_t w_t l_{it} + P_t q_t x_{it}, \tag{1}$$

where w_t denotes the economy-wide real wage rate, q_t the economy-wide real price for

[4] It should be noted that our benchmark specification of the aggregate supply relation even applies for more general production functions, which for example feature physical capital k_{it}, e.g., $y_{it} = a_t l_{it}^{1-\alpha} x_{it}^{\alpha\mu} k_{it}^{\alpha(1-\mu)}$, where $\mu \in [0,1]$.

commodities, and P_t the aggregate price level. After goods are produced and sold in the goods market, firms repay loans with the nominal interest $i_t^l Z_{it}$ at the end of the period. Hence, these loans are supplied and repaid within a period and are not accumulated. Accordingly, total costs of firm i in period t consist of wage payments $P_t w_t l_{it}$, payments for commodities $P_t q_t x_{it}$, and interest payments on loans $i_t^l Z_{it}$. Cost minimization subject to the production technology and to the liquidity constraint (1) for given prices implies

$$R_t^l w_t = mc_{it} \left(1-\alpha\right) a_t l_{it}^{-\alpha} x_{it}^{\alpha}, \tag{2}$$

and $R_t^l q_t = mc_{it} \alpha a_t l_{it}^{1-\alpha} x_{it}^{\alpha-1}$, where R_t^l denotes the gross lending rate $R_t^l \equiv 1 + i_t^l$. By applying both first-order conditions, there exist various ways to express real marginal costs. According to the labor demand condition (2), real marginal costs can be expressed as a function of the lending rate and real unit labor costs $s_{it} = w_t l_{it} / y_{it}$:

$$mc_{it} = \left(1-\alpha\right)^{-1} R_t^l s_{it}. \tag{3}$$

Equation (3), which reveals that real marginal costs increase with the lending rate and with unit labor costs, will be applied for our benchmark specification. The final good is an aggregate of the differentiated goods. The aggregator of differentiated goods is given as $y_t^{(\epsilon-1)/\epsilon} = \int_0^1 y_{it}^{(\epsilon-1)/\epsilon} di$, with $\epsilon > 1$, where y denotes the number of units of the final good and ϵ the constant elasticity of substitution between these differentiated goods. Let P_i denote the price of good i set by firm i. Then, the cost minimizing demand for each differentiated good is given by $y_{it} = \left(P_{it}/P_t\right)^{-\epsilon} y_t$, where $P_t^{1-\epsilon} = \int_0^1 P_{it}^{1-\epsilon} di$.

Firms are further characterized by Calvo's (1983) staggered price setting, modified to allow for a history dependent evolution of inflation, as in Galí et al. (2001). In particular, we assume that firms may reset their prices with the probability $1-\phi$, independent of the time elapsed since the last price setting. A fraction ω of the latter firms is assumed to set their prices according to the following simple rule-of-thumb: $\widetilde{P}_{it} = \pi_{t-1} P_{t-1}$, where π_t denotes the inflation rate $\pi_t = P_t / P_{t-1}$. The fraction $1-\omega$ is assumed to set their prices in an optimal way. These firms maximize their market value, which equals the expected sum of discounted profits $E_t \sum_{s=0}^{\infty} \eta_{t,t+s} \Delta_{it+s}^f$, where $\Delta_{it}^f \equiv \left(P_{it} - P_t mc_{it}\right) y_{it}$. Future profits are weighted with the (stochastic) discount factor $\eta_{t,t+s}$, which originates in the households' savings decision, as the managers of the firms are assumed to act on behalf of the firm owners, i.e., the households (see Danthine and Donaldson, 2002). In each period these firms set new prices P_{it}^* according to $\max_{P_{it}^*} E_t \sum_{s=0}^{\infty} \phi^s \eta_{t,t+s} (P_{it}^* y_{it+s} - P_{t+s} mc_{it+s} y_{it+s})$, s.t. $y_{it+s} = (\overline{\pi}^s P^*{}_{it})^{-\epsilon} P_{t+s}^{\epsilon} y_{t+s}$. The remaining fraction $\phi \in (0,1)$ of firms adjust their prices with the average inflation rate $\overline{\pi} \geq 1$, which allows to consider different values for the steady state inflation rate (see Woodford, 2003). Now suppose that there exists a steady state, and use that firms only differ with regard to the price setting behavior. Then, the log-linearized version of the first-order condition and the price aggregator $P_t^{1-\epsilon} = (1-\phi)\left[(1-\omega)\left(P^*\right)^{1-\epsilon} + \omega \widetilde{P}_t^{1-\epsilon}\right] + \phi \overline{\pi} P_{t-1}^{1-\epsilon}$ can be combined to give

$$\widehat{\pi}_t = \gamma_f E_t \widehat{\pi}_{t+1} + \gamma_b \widehat{\pi}_{t-1} + \chi \widehat{mc}_t, \tag{4}$$

where $\widehat{k}_t$ denotes the percent deviation from the steady state value $\overline{k}$ of a generic variable k, $\widehat{k}_t = \log(k_t) - \log(\overline{k})$ and the composite parameter in (4) are given by $\gamma \equiv \phi + \omega\left[1 - \phi(1-\beta)\right]$, $\gamma_f \equiv \beta\phi/\gamma$, $\gamma_b \equiv \omega/\gamma$, $\chi \equiv (1-\omega)(1-\phi)(1-\beta\phi)\,\xi\gamma^{-1}$, and $\xi \equiv \frac{1-\alpha}{1+\alpha(\epsilon-1)}$ (see Galí et al., 2001), where $\beta \in (0,1)$ denotes the constant discount rate of households. Equation (4), which summarizes the evolution of the inflation rate, is also known as the *hybrid* marginal cost based Phillips curve, as it introduces a backward-looking element into an otherwise entirely forward-looking New Keynesian Phillips curve.

We assume that there is a continuum of identical and perfectly competitive financial intermediaries of mass one. They receive deposits D_t from households and supply loans $Z_t = \sum_{i=0}^{1} Z_{it}$ to firms at the nominal interest rate i_t^l. At the end of each period, deposits D_t are repaid to the households together with interest earnings $i_t D_t$. It should be noted that $R_t = 1 + i_t$ further equals the risk free interest rate on one-period riskless government bonds, which is assumed to be set by the central bank (see Appendix A). Any profits are paid to the owners, i.e., the households. Financial intermediaries face costs of managing loans, which amount a constant value $\kappa \geq 0$ per unit of loans. We further consider a financial market imperfection by which interest rate effects on firms' lending costs can be accelerated. Instead of providing an explicit microfoundation, we introduce, for simplicity, a continuously differentiable function $\Psi(R_t)$, that summarizes adverse effects of the risk-free nominal interest rate on the return of risky investments. This function can be interpreted as a measure for the likelihood of defaults on loans, which increases with the interest rate R_t. This property, $\Psi'(R_t) \geq 0$, can, for example, be rationalized by the willingness of firms to invest in risky projects under asymmetric information and debt financing, when the interest rate on risk free investments R_t is high (see, e.g., Stiglitz and Weiss, 1981). The profits of financial intermediaries are, therefore, given by $\Delta_t^b = R_t^l\left[1 - \Psi(R_t)\right] Z_t - R_t D_t - \kappa Z_t$, where we assume that $\Psi(R_t) \in (0,1)$. Maximizing profits subject to the balance sheet constraint $Z_t = D_t$, leads to a first-order condition that relates the risk-free interest rate $R_t \equiv 1 + i_t$ to the lending rate $R_t^l \equiv 1 + i_t^l$. Its log-linearized version is given by

$$\widehat{R}_t^l = (1 + \psi_R)\,\widehat{R}_t, \tag{5}$$

where the coefficient ψ_R consists of two components, $\psi_R = \psi_1 - \psi_2$, which are non-negative and given by $\psi_1 \equiv \frac{\bar{\Psi}'\bar{R}}{1-\bar{\Psi}} \geq 0$ and $\psi_2 \equiv \frac{\kappa}{\bar{R}+\kappa} \geq 0$. As a consequence, $1 + \psi_R$ can either be smaller or larger than one, depending on whether costs of financial market imperfections, measured by Ψ and Ψ', or managing costs κ are more pronounced. Hence, the effects of a change in the monetary policy rate R_t on the lending rate R_t^l are accelerated for $\psi_1 > \psi_2 \Leftrightarrow \psi_R > 0$, indicating the existence of strong financial market imperfections. When managing costs κ are sufficiently high, $\psi_1 < \psi_2$, the coefficient ψ_R becomes negative, such that the lending rate rises by less than one for one with the monetary policy rate. Put differently, a change in the risk-free interest rate is then not completely passed through to the lending rate.[5] Hence, our simple reduced form specification of the cost structure in the financial intermediary sector suffices to allow for differences in the impact of interest rate

[5] A more elaborate analysis of interest rate pass-through should account for incomplete competition in the banking sector and loan price rigidities (see Hannan and Berger, 1991).

changes on the lending costs of firms.

The model's aggregate supply behavior is characterized by the log-linearized version of the labor demand condition (3), $\widehat{mc}_t = \widehat{R}^l_t + \widehat{s}_t$, the marginal costs based Phillips curve (4), and the first-order condition of financial intermediaries (5), which can be combined to give the following *interest-rate-augmented* Phillips curve:

$$\widehat{\pi}_t = \gamma_f E_t \widehat{\pi}_{t+1} + \gamma_b \widehat{\pi}_{t-1} + \chi \widehat{s}_t + \chi \left(1 + \psi_R\right) \widehat{R}_t. \qquad (6)$$

It should be noted that percentage deviations of inflation from its steady state value $\widehat{\pi}_t = \log(\pi_t) - \log(\overline{\pi})$ can be replaced by the log of inflation, if the steady state inflation rate is assumed to be equal to one, as in Galí et al. (2001). However, the steady state of the model, which further features optimizing households with rational expectations, is characterized by the restriction $\overline{\pi} = \beta \overline{R}$, which originates in the households' consumption Euler equation (see Appendix A). We, therefore, abstain form setting $\overline{\pi} = 1$, as different steady state values for the nominal interest rate $\overline{R}$ are associated with different steady state inflation rates $\overline{\pi}$.

According to (6), an increase in the central bank interest rate above its steady state value, $\widehat{R}_t > 0$, induces firms – ceteris paribus – to raise their prices, such that the current inflation rate exceeds its steady state value, $\widehat{\pi}_t > 0$. This is the quintessence of the cost channel of monetary transmission. The response of inflation to a monetary contraction, is nevertheless, jointly determined by the endogenous response of unit labor costs $\widehat{s}_t$, i.e., by adjustments of aggregate demand, and by the cost channel, rather than by the latter alone. This principle will be illustrated in Section 4, where we present inflation responses for a simple dynamic general equilibrium model, in which firms and financial intermediaries are embedded as described above. In the subsequent section, we assess the relevance of direct interest rate effects for the short-run aggregate supply behavior by providing estimates for the interest-rate-augmented Phillips curve (6).

3 Empirical analysis

Recently, several studies have found that a standard New Keynesian Phillips curve, which does not account for direct interest rate effects due to the cost channel, already serves as an useful description for inflation dynamics in the US, the Euro area, and its member countries (see Galí and Gertler, 1999; Galí et al., 2001; Benigno and López-Salido, 2001; Sbordone, 2002). In these studies, structural changes in inflation are due to New Keynesian demand effects, which depend on households' intertemporal substitution of consumption and leisure. Our empirical analysis builds on this evidence and aims at disclosing if there are – on top – significant interest rate effects on firms' costs in G7 countries, which contribute to the predictability of inflation rates. The empirical analysis can thus be viewed as an empirical assessment to a structural approach of the cost channel, for which Barth and Ramey (2001) already found significant industrial level evidence for the US by applying vector autoregressions.

3.1 The benchmark specification

According to the interest-rate-augmented Phillips curve (6), inflation is measured by percentage deviations from its mean, $\widehat{\pi}_t$, to account for different steady state levels of inflation and nominal interest rates across the G-7 countries. By (6), current inflation rises with lagged and expected future inflation, and with percentage deviations of real unit labor costs, $\widehat{s}_t$, and of the short-run nominal interest rate, $\widehat{R}_t$, from their means. Since we aim at assessing the magnitude and the significance of the coefficients on these determinants for current inflation, we do not apply a structural decomposition of the reduced form parameters:

$$\widehat{\pi}_t = \gamma_b \widehat{\pi}_{t-1} + \gamma_f E_t \widehat{\pi}_{t+1} + \gamma_s \widehat{s}_t + \gamma_R \hat{R}_t, \tag{7}$$

Estimations are carried out applying quarterly time series data from the OECD Business Sector database, the IMF's International Financial Statistics (IFS), and the Bureau of Labor Statistics. Phillips curve estimations are conducted for Canada, France, Germany, Italy, Japan, the UK, and the US. Inflation is measured by using the GDP deflator and, alternatively, by the consumer price index (CPI). As a measure for short-run nominal interest rates we use three-month treasury bill rates in our estimations. Real unit labor costs are constructed as the ratio of total compensation to GDP.[6]

The overall sample period spans the time interval 1980-1997. By choosing this interval, the time period of the two oil price crises during the 1970s and the recent past are excluded. Estimations are conducted using generalized methods of moments (GMM) and, hence, for the vector of instruments, $\mathbf{z}_t$, a set of orthogonality conditions hold. Consequently, equation (7) can be written as

$$E_t \left\{ \left(\widehat{\pi}_t - \gamma_b \widehat{\pi}_{t-1} - \gamma_f \widehat{\pi}_{t+1} - \gamma_s \widehat{s}_t - \gamma_R \widehat{R}_t \right) \mathbf{z}_t \right\} = 0. \tag{8}$$

Since not all current information are available to the public at the time they form expectations, contemporary variables are not used as instruments. In particular, the vector of instruments $\mathbf{z}_t$ includes four lags of inflation, real unit labor costs, and the T-bill rate, as well as up to four lags of real commodity prices.[7] To account for possible correlation in the moment conditions and to control for autocorrelation and heteroscedasticity in unknown form in the weighting matrix, we allow for Newey-West correction up to order eight.

Table 1 reports the estimates of the interest-rate-augmented Phillips curve as specified in (7). All estimated coefficients are positive in sign and are in general found to be statistically significant at the 5 percent level. Current inflation is always significantly affected by lagged and expected future inflation rates, while the impact of the latter component is found to be more pronounced in all countries except for Italy and the US. This pattern is consistent with the results in Galí et al. (2001), who find a higher degree of backward-lookingness in the US than in the Euro area. Real unit labor costs exhibit – except for Germany and

[6] More precisely, data on three-month treasury bill rates, CPI and the commodity price index are drawn from the IFS. All remaining data are taken from the OECD Business Sector database, except data on total compensation and total employees for the US. The latter are obtained from the Bureau of Labor Statistics.

[7] In particular, the lag length of commodity prices equals one for the UK, two for Canada, three for Germany and France and four for Italy and the US.

Japan – significant coefficients, lying between 0.01 for Italy and 0.099 for France. Our estimates further reveal significant direct interest rate effects in Canada, France, Italy, the UK, and in the US. In contrast, the estimated coefficient on the nominal interest rate $\hat{\gamma}_R$ is not significant for Germany and Japan (indicated by #). The smallest value for $\hat{\gamma}_R$ is observed for Italy (0.015) and the largest for the UK (0.076).

Table 1: Estimates of the Interest-rate-augmented Phillips-Curve (GDP-deflator)

	$\hat{\gamma}_f$	$\hat{\gamma}_b$	$\hat{\gamma}_s$	$\hat{\gamma}_R$	$\hat{\gamma}_R/\hat{\gamma}_s$	$J-Test$
Canada	0.71 (0.016)	0.27 (0.015)	0.015 (0.005)	0.017 (0.002)	1.1	0.56
France	0.54 (0.009)	0.33 (0.007)	0.099 (0.008)	0.024 (0.008)	0.2	0.53
Germany	0.53 (0.057)	0.47 (0.033)	$0.012^{\#}$ (0.008)	$0.005^{\#}$ (0.022)	–	0.40
Italy	0.48 (0.001)	0.50 (0.001)	0.010 (0.001)	0.015 (0.001)	1.5	0.62
Japan	0.77 (0.082)	0.18 (0.065)	$0.005^{\#}$ (0.010)	$0.024^{\#}$ (0.066)	–	0.74
UK	0.48 (0.041)	0.33 (0.059)	0.058 (0.024)	0.076 (0.022)	1.3	0.31
US	0.39 (0.016)	0.53 (0.013)	0.024 (0.012)	0.030 (0.009)	1.3	0.49

Notes: Figures in round brackets denote standard errors. The J-Test describes a test statistic for the null hypothesis that the overidentifying restrictions are satisfied. For the latter test p-values are reported; all estimated coefficients are statistically significant at the 5 percent level except those marked with #, which are not statistically significant at the 10 percent level; – implies that the estimated coefficient on real unit labour costs and on the T-bill rate is not statistically significant and therefore the ratio $\hat{\gamma}_R/\hat{\gamma}_s$ is not computed in these cases. Precise details of the estimation procedure are presented in the text.

The values for $\hat{\gamma}_R$ are more informative when they are compared to the point estimates for the coefficient on unit labor costs, which summarizes the strength of the demand channel. We therefore present the ratio $\hat{\gamma}_R/\hat{\gamma}_s$ in Table 1, which serves as a measure for the relative importance of both cost components.[8] Notably, this ratio always exceeds one, except for France. Now recall that the ratio $\hat{\gamma}_R/\hat{\gamma}_s$ equals $1+\psi_R$ according to the theoretical model, where the coefficient ψ_R governs the impact of the monetary policy rate on the lending costs of firms. Since the estimates imply a positive value for ψ_R for Canada (0.1), Italy (0.5), the UK (0.3) and for the US (0.3), the impact of a rise in the monetary policy rate on firms' costs seems to be accelerated in these countries. Notably, our result for the US ($\hat{\gamma}_R/\hat{\gamma}_s = 1.3$) closely relates to the findings in Ravenna and Walsh (2003), who estimate a purely forward-looking version of the interest-rate-augmented Phillips curve for the US.

Table 2: Estimates of the Interest-rate-augmented Phillips-Curve (CPI-based)

	$\hat{\gamma}_f^{cpi}$	$\hat{\gamma}_b^{cpi}$	$\hat{\gamma}_s^{cpi}$	$\hat{\gamma}_R^{cpi}$	$\hat{\gamma}_R^{cpi}/\hat{\gamma}_s^{cpi}$	$J-Test$
Canada	0.48 (0.049)	0.46 (0.039)	0.029# (0.021)	0.049* (0.029)	1.7	0.13
France	0.46 (0.002)	0.50 (0.010)	0.027 (0.002)	0.013* (0.007)	0.5	0.63
Germany	0.31 (0.001)	0.68 (0.002)	0.010 (0.001)	0.004 (0.001)	0.4	0.27
Italy	0.41 (0.026)	0.56 (0.029)	0.006# (0.016)	0.031# (0.032)	–	0.62
Japan	0.65 (0.036)	0.40 (0.027)	0.005# (0.008)	0.003# (0.019)	–	0.80
UK	0.44 (0.004)	0.48 (0.006)	0.048 (0.014)	0.077 (0.003)	1.6	0.61
US	0.31 (0.009)	0.62 (0.008)	0.036 (0.005)	0.030 (0.005)	0.8	0.69

Notes: Coefficients that are only significant at the 10 percent level are marked with *. See notes to Table 1.

To examine whether our findings are sensitive with regard to the choice of the price level, we further carry out Phillips curve estimations using CPI instead of the GDP deflator. The estimations are conducted as before, where the set of instruments only differs with regard to the inflation measure. Overall, the results, which are reported in Table 2, confirm our

[8]Since real unit labor costs and short-term interest rates are not found to significantly contribute to inflation dynamics in Germany and Japan, we abstain from reporting the ratio $\hat{\gamma}_R/\hat{\gamma}_s$ for both countries.

earlier findings, in particular, on the existence of direct interest rate effects. In five of the G7 countries the nominal interest rate is found to have a statistically significant impact on CPI-based inflation.[9] When inflation is measured by using CPI, Germany exhibits significant coefficients on real unit labor costs and on the nominal interest rate, while both coefficients are insignificant for Italy and Japan. Overall, the cross-country evidence suggests that the interest-rate-augmented Phillips curve is less appropriate for the description of short-run inflation dynamics, when inflation is measured by CPI than by the GDP-deflator.

Nevertheless, our results reported in Table 1 and Table 2 show that the coefficient on the interest rate in the Phillips curve is significant under both inflation measures for Canada, France, the UK, and the US. For France and the UK, for example, the relative strength of the direct interest rate effect, as measured by $\hat{\gamma}_R/\hat{\gamma}_s$, is more pronounced for the CPI-based estimates. Put differently, the GDP-deflator-based estimates appear to indicate a weaker cost channel in these countries, where the ratio $\hat{\gamma}_R/\hat{\gamma}_s$ rises from 0.2 (GDP-deflator) to 0.5 (CPI) and from 1.3 to 1.6, respectively. A reason for this result might be that cyclical components of production costs are more pronounced for consumption goods than for investment goods, such that the marginal costs of production of the latter are less prone to changes in the short-run nominal interest rate at business cycle frequency.

To summarize, our estimates for the benchmark specification (7) reveal that there exists a significant impact of short-term interest rates on inflation dynamics in the majority of G7 countries. Direct interest rate effects on current inflation are further found to be larger in Canada, Italy, the UK and in the US than in France and in Germany (for CPI), while we could not provide any evidence at all on the cost channel in Japan.

3.2 Alternative specifications

In the previous section, we presented evidence on direct interest rate effects on current inflation for the majority of G7-countries. To provide further evidence on the robustness of our main result and to facilitate comparisons with related studies, we continue by applying alternative specifications. We first consider a version of the aggregate supply constraint, where costs of working capital are neglected and the interest rate coefficient is assumed to be zero ($\gamma_R = 0$):

$$E_t \left\{ \left(\widehat{\pi}_t - \gamma_b \widehat{\pi}_{t-1} - \gamma_f \widehat{\pi}_{t+1} - \gamma_s \widehat{s}_t \right) \mathbf{z}_t \right\} = 0. \tag{9}$$

In what follows, we refer to equation (9) as the *standard* (hybrid) New Keynesian Phillips curve. Such a specification of the aggregate supply relation has already been shown to serve as an useful description of short-run inflation dynamics for a smaller set of countries and for different time intervals (see, e.g. Galí and Gertler, 1999; Galí et al., 2001; Sbordone, 2002; Benigno and López-Salido, 2002). The purpose of this exercise is twofold. Firstly, we want to assess whether real unit labor costs alone serve as an useful proxy for real marginal costs for the G7 countries in the investigated time period. Secondly, a comparison between the estimates of the coefficients in (8) and (9) allows to assess the impact of the inclusion of interest rates for the pricing decision of firms. In Table 3 we report GMM estimates

[9] For France and Germany, interest rate coefficients are found to be only significant at the 10 percent level (indicated by *).

for (9), in which inflation is measured using the GDP deflator, as in the specification that underlies the estimates presented in Table 1.[10]

Table 3: Estimates of the standard (hybrid) New Keynesian Phillips-Curve (GDP-deflator)

	$\hat{\gamma}_f^{nk}$	$\hat{\gamma}_b^{nk}$	$\hat{\gamma}_s^{nk}$	$J-Test$
Canada	0.54 (0.021)	0.45 (0.016)	0.017* (0.009)	0.21
France	0.63 (0.029)	0.47 (0.014)	0.021 (0.004)	0.45
Germany	0.67 (0.090)	0.35 (0.062)	0.075 (0.023)	0.25
Italy	0.37 (0.023)	0.58 (0.013)	0.018 (0.003)	0.54
Japan	0.64 (0.003)	0.36 (0.004)	0.012 (0.001)	0.22
UK	0.59 (0.044)	0.35 (0.051)	0.031# (0.022)	0.31
US	0.38 (0.018)	0.54 (0.016)	0.068 (0.005)	0.54

Notes: See notes to Table 1 and 2.

Consistent with the theoretical model, all coefficients are found to have a positive and – except for the UK – statistically significant impact on inflation. Our estimates are generally of similar size to the ones reported in Benigno and López-Salido (2002), and Galí et al. (2001, 2003). The results further confirm the finding in Galí et al. (2001, 2003), that the forward-looking inflation component is in general more pronounced in European countries than in the US. A closer look at the estimated coefficients on the forward-looking inflation component in Table 1 and Table 3 shows that they are larger in France and in the UK when direct interest rate effects are disregarded, while they are smaller in Canada, Italy and the US. For the former set of countries one might suspect that a significant coefficient on the

[10] The set of instruments only differs with regard to the nominal interest rate, which is now omitted.

interest rate measures information on future inflation instead of working capital costs. To illustrate this, suppose that the central bank adjusts its instrument in response to changes in the (expected future) inflation rate, as for example suggested by Clarida et al. (2000), such that the monetary policy rate contains some information about (future) inflation. If these information are not contained in the other regressors in the benchmark specification (7), then it might be possible to find positive interest rate effects on inflation, even if there is no cost channel at work. However, the conditional expectation of the future inflation rate is already considered as an explanatory variable in (7), which implies that the current monetary policy rate cannot contain any additional information about future realizations of this particular inflation rate under rational expectations.

In fact, larger values for the forward-looking component $\hat{\gamma}_f^{nk} > \hat{\gamma}_f$ can easily be rationalized by direct cost channel effects: Consider again that the central bank raises the nominal interest rate with (expected future) inflation. Under the hypothesis that the cost channel is effective, the exclusion of the nominal interest rate from the Phillips curve should lead to an overestimation of the forward-looking inflation component, as it contains information about the current nominal interest rate.[11] While this argument is consistent with our results regarding the forward-looking component for France and for the UK, it can, evidently, not account for the inverse shifts observed in the forward-looking inflation components for Canada, Italy, and the US.

According to the cost channel view, direct interest rate effects on current inflation should still prevail if a monetary policy reaction function is explicitly considered. In particular, we expect to find significant interest rate coefficients in the aggregate supply relation, even if it is jointly estimated with an interest rate feedback rule, which links the nominal interest rate to current or future inflation rates. These expectations are in fact confirmed by the results, which are obtained by estimating an interest-rate augmented Phillips curve as specified in (8) with a simple interest rate feedback rule using simultaneous GMM.[12] The results are provided in Table 4 in Appendix B. For each country we conduct two sets of estimations, of which the first set corresponds to an interest rate rule featuring current inflation $\widehat{R}_t = \rho_\pi \widehat{\pi}_t + \varepsilon_t$, and the second to a forward-looking rule, $\widehat{R}_t = \rho_\pi E_t \widehat{\pi}_{t+1} + \varepsilon_t$, where ε_t denotes innovations. Though, these rules are evidently too simple to summarize the conduct of monetary policy of real world central banks, they suffice to account for the alternative hypothesis about positive interest rate coefficients as laid out before. The results show that all examined countries exhibit positive and significant interest rate coefficients $\hat{\gamma}_R^{sim}$ under both specifications of the interest rate rule. The single exception is France, where

[11] A similar argument can be applied for the coefficient on real unit labor costs, which is positively related to the output gap that is found to serve as an indicator for interest rate policy (see, e.g., Clarida et al., 2000). In accordance with this view, significant coefficients on real unit labor costs presented in Table 1 and Table 3 show that their impact on inflation is – except for France – more pronounced when the nominal interest rate is excluded.

[12] Hereby, the set of instruments includes two lags of inflation, real unit labor costs and interest rates. We additionally allow for two lags of real commodity prices in all countries except for the US and France, where the lag length equals one and three, respectively. Estimations were not carried out for Germany and Japan, since in both countries the single equation estimations based on (8) did not produce significant interest rate coefficients.

the applied specification does not lead to reasonable results for the aggregate supply relation. Overall, this exercise disproves the hypothesis that positive interest rate coefficients in the aggregate supply relation just reflect a positive feedback from inflation to the central bank instrument.

Another argument that might be relevant for interest rate effects on current inflation refers to the role of commodity prices as an indicator variable for a central bank. Commodity prices are often included in the central bank's information set in monetary VARs (see Christiano et al., 1999), which allows for a more precise isolation of exogenous policy actions (shocks) from endogenous adjustments of the interest rate target, i.e., the federal funds rate. As stressed in Sims (1992), by omitting commodity prices an interest rate hike can cause prices to rise, when the central bank has access to additional information about a nascent inflation, which is included in commodity prices. The positive price response to interest rate shocks is also known as the "price puzzle", where the notion indicates that prices are usually expected to decline in response to contractionary monetary actions, which are independent from endogenous interest rate adjustments.

Similarly, it might be possible that significant interest rate effects, as reported in Table 1 and 2, are due to the fact that the current monetary policy rate contains information about commodity prices, which exert an independent upward pressure on final goods prices (either measured by the GDP deflator or CPI). To assess the validity of this alternative hypothesis, we consider commodity prices as a determinant for current inflation, and include percentage deviations of real commodity prices from their mean, $\widehat{q}_t$, as an additional variable in the aggregate supply relation. In order to be consistent with the theoretical framework presented in Section 2, percentage deviations of the real wage rate from its steady state value, $\widehat{w}_t$, are further considered as an explanatory variable,

$$E_t \left\{ \left(\widehat{\pi}_t - \gamma_b \widehat{\pi}_{t-1} - \gamma_f \widehat{\pi}_{t+1} - \gamma_w \widehat{w}_t - \gamma_R \hat{R}_t - \gamma_q \widehat{q}_t \right) \mathbf{z}_t \right\} = 0, \tag{10}$$

where real wages are constructed as the ratio of total real compensation to total employees. The specification (10) is derived from the first-order conditions of firms (see Section 2), where changes in total factor productivity are neglected, $\widehat{a}_t = 0$, such that real marginal costs can be expressed as a function of real factor prices, $\widehat{mc}_t = \widehat{R}_t^l + \alpha \widehat{q}_t + (1-\alpha)\,\widehat{w}_t$.

The coefficients in (10) are estimated as in Section 3.1. The set of instruments includes four lags of inflation, the real wage rate, the nominal interest rate, and real commodity prices for all countries. The results, which are provided in Table 5 in Appendix B, show that the real-factor-price based specification (10) fails to summarize the inflation dynamics in the countries under consideration. To be more precise, we only find significant coefficients on all components of real marginal costs for Canada and Italy. For Germany and Japan, we are unable to disclose any significant cost component. Nevertheless, we can not find evidence in any of the seven countries for the hypothesis that interest rate effects are due to omitted real commodity prices: Whenever real commodity prices significantly contribute to current inflation rates, like in Canada, Italy or in the UK, we find significant coefficients on the nominal interest rate. Overall, specification (10) seems less useful to adequately describe inflation dynamics than the former specifications. This might be due to the fact that a real-factor-price based specification – in contrast to a specification with unit labor costs – does not account for changes in the total factor productivity and for alternative technologies featuring additional production inputs, as for example, physical capital (see footnote 4).

3.3 Financial structure and interest rate pass-through

The cost channel view suggests that the impact of nominal interest rate changes on firms' marginal costs and, hence, on their price setting behavior originates in their holdings of working capital. The opportunity costs of the latter generally rise with short-run nominal interest rates and can, therefore, be affected by monetary policy measures. The extent to which changes in the monetary policy rate alter firms' marginal costs, however, depends on the pass-through of official interest rates to market rates or (bank) lending rates. While the pass-through to lending rates is likely to depend on the regulation and competition in the financial intermediary sector (see Hannan and Berger, 1991), short-run market rates should in general immediately adjust to changes in the monetary policy rate. Thus, the financial structure should be relevant for the strength of the cost channel of monetary transmission. Accordingly, a stronger reliance of firms on bank loans and a lower degree of interest rate pass-through, should lead to a less pronounced impact of monetary policy rate changes on firms' marginal costs.

One would, therefore, expect a lower interest rate pass-through in "bank-based systems", such as in continental European countries (see Allen and Gale, 2004), and herein in Germany more than, for example, in France, as the former has been known to exhibit a highly regulated banking sector (see Mayer, 1990; Mojon, 2000). This view is supported by Borio and Fritz (1995) and Mojon (2000), according to which the pass-through is most incomplete for Germany, while it is somewhat higher for France and Italy.[13] Moreover, tight relations between banks and firms in Germany are likely to lower the pass-through (Mojon, 2000). A similar argument also applies for Japan where bank loans are particularly important (see Corbett and Jenkinson, 1997). On the contrary, Canada, the UK and the US are referred to as "market-based systems" and are characterized by a high degree of financial market liberalization and securitization (see, e.g., Engert et al., 1999; Allen and Gale, 2000, 2004). These characteristics are consistent with empirical evidence by Cottarelli and Kourelis (1994) and Sellon (2002), suggesting that the interest rate pass-through is rather instantaneously and is regarded as more complete than in continental European countries. In view of these arguments, the effect of the cost channel should be more pronounced in Canada, the UK, and the US than in continental European countries and in Japan.

The results for our benchmark specification, which are presented in Table 1, in fact confirm these expectations. The point estimates for the coefficients on the nominal interest rate $\hat{\gamma}_R$ and on unit labor costs $\hat{\gamma}_s$, and, in particular their ratio $\hat{\gamma}_R/\hat{\gamma}_s$ reveal the existence of a strong cost channel in Canada, Italy, the UK, and the US, whereas France, Germany, and Japan exhibit a small or even no significant cost channel. Since the ratio $\hat{\gamma}_R/\hat{\gamma}_s$

[13] Hofmann (2003) also provides evidence on heterogeneous interest rate pass-through in European countries.

corresponds to the elasticity $1+\psi_R$ in the theoretical model, we can further read off values for the coefficient ψ_R, which should be equal to zero when changes in the monetary policy rate are passed through by one for one. Evidently, ψ_R is positive for Canada (0.1), Italy (0.5), UK (0.3) and for the US (0.3) (see Table 1), indicating an amplification of monetary policy rate effects on firms' costs for working capital. In accordance with the (broad) credit channel view, our results thus indicate the existence of substantial financial frictions, which are responsible for the acceleration of monetary policy shocks, consistent, for example, with the findings in Oliner and Rudebusch (1996), Bernanke et al. (1999), or Dedola and Lippi (2003). Notably, Italy exhibits the highest value for ψ_R, such that interest rate changes have a strong impact on firms' costs, which might, as argued by Cecchetti (2000), be due to a "less healthy" banking system in the investigated time period. In the case of France (and Germany), however, ψ_R is negative, indicating that interest rate effects are dampened by an incomplete interest rate pass-through, consistent with the empirical evidence cited above. Thus, the variations found in our estimates for the interest rate coefficients in the aggregate supply relation and, therefore, on the strength of direct interest rate effects correspond to the evidence on differences in the financial structure and interest rate pass-through.

4 Monetary policy and the Cost Channel

In this section we examine the implications of direct interest rate effects for the transmission of shocks to the monetary policy rate. For this, we embed firms and financial intermediaries, as characterized in Section 2, in a simple monetary business cycle model. The model further features households, which supply labor as well as commodities to firms and deposit funds at the financial intermediaries at the monetary policy rate. The details of the model can be found in Appendix A. Log-linearizing the model at the deterministic steady state and reducing the set of endogenous variables, we end up with the following conditions, which describe a rational expectations equilibrium in inflation $\widehat{\pi}_t$, output $\widehat{y}_t$, and the short-run nominal interest rate, $\widehat{R}_t$:

$$\widehat{\pi}_t = \gamma_f E_t\widehat{\pi}_{t+1} + \gamma_b\widehat{\pi}_{t-1} + \gamma_R\widehat{R}_t + \gamma_y\widehat{y}_t, \tag{11}$$

$$\sigma\widehat{y}_t = \sigma E_t\widehat{y}_{t+1} - \widehat{R}_t + E_t\widehat{\pi}_{t+1}, \tag{12}$$

$$\widehat{R}_t = \rho_\pi\widehat{\pi}_t + \varepsilon_t, \tag{13}$$

where ε_t denotes the monetary policy shock, which is i.i.d. with a zero mean. The coefficients γ_R and γ_y are defined as $\gamma_R \equiv \chi\,(1+\psi_R) > 0$ and $\gamma_y \equiv \chi[\sigma + (\sigma_l + \alpha)\,/\,(1-\alpha)] > 0$, where σ (σ_l) denotes the strictly positive inverse of the intertemporal substitution elasticity of consumption (labor). Equation (12) is also known as the forward-looking IS curve and stems from the households' consumption Euler equation. It should be noted that the interest rate feedback rule (13) does not feature responses to the output gap, interest rate smoothing, or forward-looking elements, which are, for example, found to contribute to the description of US monetary policy (see, e.g., Clarida et al., 2000). We apply a simple feedback rule in order to facilitate a straightforward identification of the cost channel effects, and to avoid instabilities stemming from interest rate responses to future inflation (see, e.g., Woodford, 2003).

All results in this section are based on the model's fundamental solution, i.e., the minimum state variable solution. Before presenting simulated impulse responses, we analytically derive the main qualitative properties of a simplified version, where prices are set in a non-backward-looking way, $\omega = 0$.[14] The following proposition summarizes the main results for interest rate coefficients $\gamma_R < \widetilde{\gamma_R}$, where $\widetilde{\gamma_R} \equiv \gamma_y \sigma^{-1} + 1/\rho_\pi$, which is clearly satisfied by all point estimates $\hat{\gamma}_R$ for any reasonable inflation elasticity ρ_π of the interest rate rule (13).

Proposition 1 *Suppose that price setters are entirely forward-looking* $\omega = 0$ *and that the aggregate supply constraint satisfies* $\gamma_R \in (0, \widetilde{\gamma_R})$. *Then, a contractionary monetary policy shock leads to a decline in inflation if* $\gamma_R < \gamma_y/\sigma$, *and to a rise in inflation if* $\gamma_R > \gamma_y/\sigma$, *and to a decline in output. Higher values for the interest rate coefficient* γ_R *decrease the impact response of output and increase the impact response of inflation.*

Proof. Since the model with $\omega = 0$ exhibits no endogenous state, the fundamental solution takes the generic form, $\widehat{\pi}_t = \delta_\pi \varepsilon_t$ and $\widehat{y}_t = \delta_y \varepsilon_t$. Applying the method of undetermined coefficients for the model (11)-(13) with $\omega = 0$, leads to the conditions $\sigma \delta_y = -(\rho_\pi \delta_\pi + 1)$ and $(1 - \gamma_R \rho_\pi) \delta_\pi = \gamma_y \delta_y + \gamma_R$, and thus to the following solutions for δ_π and δ_y : $\delta_\pi = -(\gamma_y \sigma^{-1} - \gamma_R)[1 + (\gamma_y \sigma^{-1} - \gamma_R)\rho_\pi]^{-1}$ and $\delta_y = -\sigma^{-1}[1 + (\gamma_y \sigma^{-1} - \gamma_R)\rho_\pi]^{-1}$. For $\gamma_R < \widetilde{\gamma_R}$, the output response δ_y is negative and decreases with γ_R, while the inflation response δ_π is negative if $\gamma_R < \gamma_y \sigma^{-1}$ and positive if $\gamma_R > \gamma_y \sigma^{-1}$, and rises with γ_R. ■

According to the result presented in Proposition 1, a contractionary monetary policy shock leads to a more pronounced decline in output and a mitigated decline in inflation for higher values of the interest rate coefficient γ_R. Since the latter is defined by $\gamma_R \equiv \chi(1 + \psi_R)$, the impact of the cost channel is more pronounced for higher values of the coefficient ψ_R, where $\psi_R > 0$ indicates an acceleration of interest rate effects, and $\psi_R < 0$ implies an incomplete pass-through from the monetary policy rate R_t to the lending rate R_t^l. When interest rate effects are strongly accelerated, $\gamma_R > \gamma_y \sigma^{-1} \Leftrightarrow \psi_R > (\sigma_l + \alpha)/[\sigma(1-\alpha)]$, a positive interest rate innovation can then even lead to a rise in inflation. Evidently, the occurrence for such an inverse inflation response further depends on households' preferences that affect the strength of the demand channel, by which inflation tends to decline in response to a monetary contraction. In particular, the likelihood of an inverse inflation response rises with σ and declines with σ_l, which will subsequently be discussed in detail.

Next, we want to disclose the contribution of the cost channel of monetary transmission for the inflation response. We compute impulse responses of inflation to interest rate shocks by applying the coefficients of the aggregate supply relation (6), estimated for the benchmark specification, i.e., by setting $\gamma_f = \hat{\gamma}_f$, $\gamma_b = \hat{\gamma}_b$, $\gamma_y = \hat{\gamma}_s(\sigma + \frac{\sigma_l + \alpha}{1-\alpha})$ and $\gamma_R = \hat{\gamma}_R$. To isolate the effects stemming from the cost channel of monetary transmission, we further compute responses for the case where the cost channel is assumed to be non-existent. For this, we set the coefficient γ_R equal to zero, while we leave the values for the remaining coefficients in the aggregate supply relation (11) unchanged.

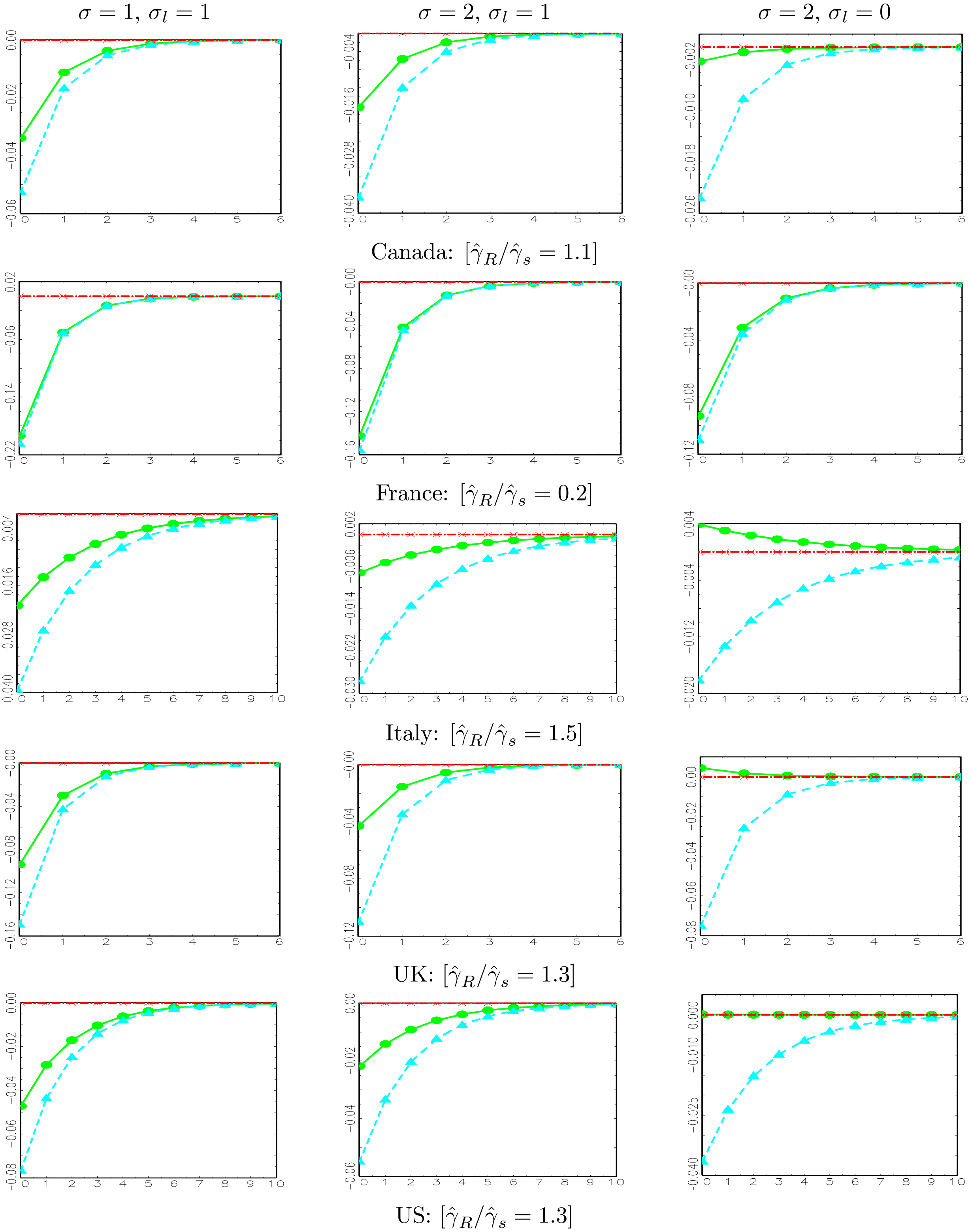
[14] The conditions for uniqueness of the fundamental solution for this model can be found in Brueckner and Schabert (2003).

The impulse responses to an interest rate shock are derived by applying the method of undetermined coefficients of the model (11)-(13) (see Appendix A). For this, we calibrate the four parameters σ, σ_l, α and ρ_π and apply the point estimates for $\hat{\gamma}_b$, $\hat{\gamma}_f$, $\hat{\gamma}_s$ and $\hat{\gamma}_R$, which are reported in Table 1 for the GDP deflator based inflation measure. To facilitate comparisons, the parameters for the aggregate demand constraint (12) and the policy rule (13) are held constant in all cases (countries) and are set equal to values that can often be found in the literature. In particular, we set the labor income share $1-\alpha$ equal to $2/3$ and the inflation elasticity ρ_π equal to 1.5. For the benchmark case, we assume that the inverse of the intertemporal substitution elasticity of consumption and of labor equal one, $\sigma = \sigma_l = 1$, implying that utility increases with log consumption and decreases with the square of labor. We further consider cases where $\sigma = 2$ and $\sigma_l = 0$, to disclose the impact of changes in consumption and labor supply on the New Keynesian demand channel and, thus, on the inflation response.

Figure 1 presents the simulated impulse response of inflation, measured in percentage deviations from its steady state value, to an one percent innovation to the interest rate, $\varepsilon_t > 0$ (see equation 13). We present results for Canada, France, Italy, the UK and the US, where direct interest rate effects are found to be statistically significant for our benchmark estimation (see Table 1). The solid lines (with circles) in Figure 1 show the inflation response to a temporary monetary contraction for the case of an existing cost channel, $\hat{\gamma}_R > 0$, and the dotted lines (with triangles) display the inflation response when the cost channel is assumed to be non-existent, $\hat{\gamma}_R = 0$. Given that $\rho_\pi > 1$, the real interest rate rises and induces households to increase their savings and to reduce consumption, such that output declines. The associated decline in employment and, thus, in desired real wages decreases real marginal costs, such that firms tend to lower their prices.

Overall, the impulse responses in Figure 1 show that direct interest rate effects on firms' marginal costs dampen the demand induced decline of inflation, consistent with the prediction in proposition 1. A closer look at the impulse responses reveals that the impact of the cost channel exhibits substantial differences between the countries. For our benchmark parameterization ($\sigma = 1$ and $\sigma_l = 1$), the initial inflation response is reduced by 38 percent (37 percent) for the US (UK) when direct interest rate effects are present, while the difference in the inflation response only amounts to 5 percent for France. Higher direct interest rate effects further raise the persistence of inflation responses, as revealed by the half-life of the impact effect: Under a cost channel, the inflation deviation from its steady state value equals half of the initial impact after 2 (1.5) periods for Italy (US), while the half-life equals 1.7 (1.2) periods when the cost channel is assumed to be non-existent. In contrast, the half-life roughly equals 0.8 periods for France regardless whether the cost channel is present or not. These effects accord to the relative size of the estimated coefficients on the interest rate and on real unit labor costs as summarized by the ratio $\hat{\gamma}_R/\hat{\gamma}_s$, which is repeated in Figure 1.

Figure 1: Simulated Inflation Responses to Interest Rate Shocks



Notes: Impulse responses for an existing cost channel (solid line with circles) are computed for $\alpha = 1/3$ and $\rho_\pi = 1.5$, $\gamma_f = \hat{\gamma}_f$, $\gamma_b = \hat{\gamma}_b$, $\gamma_y = \hat{\gamma}_s[\sigma + (\sigma_l + \alpha)/(1-\alpha)]$ and $\gamma_R = \hat{\gamma}_R$. Dotted lines with triangles show the responses without a cost channel, $\gamma_R = 0$. All impulse responses are expressed as percentage deviations from the steady state (marked by a dotted line with stars) to a one percent interest rate innovation.

To unveil the impact of the demand channel on the inflation response, we further apply variations of the parameters σ and σ_l. The impulse responses displayed in the second column in Figure 1 refer to the case where the degree of households' risk aversion is raised to $\sigma = 2$. Evidently, the difference between the solid lines and the dotted lines increases, indicating that the impact of direct interest rate effects rises in all countries, even though the coefficient on the interest rate in the aggregate supply relation (11) is unchanged. For the US (UK) the initial inflation response is mitigated by 60 percent (61 percent) by the cost channel, while the difference in the impact response equals 8 percent for France. The reason for this enhanced effect is that the households' willingness to substitute current for future consumption in response to a rise in the real interest rate is less pronounced for a higher degree of risk aversion σ. Consequently, aggregate demand declines to a smaller extend, reducing the effect on unit labor costs, such that the deflationary impact of a monetary tightening is mitigated.[15]

The third column of Figure 1 presents impulse responses for the case where the inverse of the intertemporal substitution elasticity of labor is set to zero, $\sigma_l = 0$. The latter value implies an infinite labor supply elasticity, which additionally weakens the cost alleviating effect of the New Keynesian demand channel. In this case the (demand induced) decrease in employment is not associated with a change in real wages demanded by households. Consequently, the cost pressure of direct interest rate effects can dominate the demand channel, such that inflation might even rise in response to a monetary tightening, as for the estimates for Italy and the UK. Thus, a positive interest rate innovation, i.e., a contractionary monetary policy measure, can lead to a rise in inflation when the cost channel is associated with a weak demand channel.

These inverse inflation responses, which seem to be at odds with conventional expectations about monetary policy effects, relate to the price puzzle, i.e. increases in the price level in response to a monetary contraction, which is often found in monetary VARs. The price puzzle is commonly viewed as the consequence of inappropriate VAR identification schemes, which do not fully account for the information set of the monetary authority (see Sims, 1992). Recent studies, however, show that the inverse price response does not completely vanish, even after accounting for the latter deficiency (see Christiano et al., 1999; Hanson, 2004). Based on these findings and on industry level evidence, Barth and Ramey (2001) suggest the cost channel of monetary transmission as a solution for the price puzzle. In support of this view, Christiano et al. (2004) are able to reproduce the inverse price response found in the US by embedding costs of working capital in a general equilibrium model, which is more complex than the one applied in this paper. Thus, positive price responses to interest rate hikes seem to be a rather reasonable implication of pronounced direct interest rate effects on firms' costs than a puzzling feature.

According to our simulation exercise, inflation rises in response to an interest rate shock in Italy and the UK. Though, our model is too simple to capture the entire transmission

[15] For the parameterization $\sigma = 1$ and $\sigma_l = 1$, the cost channel amplifies the initial output decline by 3.2% for Canada, 1.5% for France, 5.5% for Italy, 12% for the UK and by 5.9% for the US. For a higher degree of risk aversion ($\sigma = 2$ and $\sigma_l = 1$), this effect is more pronounced and amounts 4.3% for Canada, 2.7% for France, 6.5% for Italy, 15% for the UK and 6.8% for the US.

mechanism, which is usually characterized by a decline in the price level after a couple of periods, the inverse inflation responses relate to empirical evidence. Sims (1992), for example finds positive price responses to contractionary monetary policy shocks for France and for the UK, which decrease but do not disappear when commodity prices are included in the VARs. Similarly, temporary increases in the price level in response to an unanticipated rise in nominal interest rates are also found in recent studies for France, Italy, the UK and the US (see, e.g., Peersman and Smets, 2001; Bean et al., 2002; Dedola and Lippi, 2003). Notably, our simulated inflation response for the US almost equals zero ($\delta_{\pi e} = 0.0001$) for the parameter values $\sigma = 2$ and $\sigma_l = 0$, though the US exhibits the same ratio $\hat{\gamma}_R/\hat{\gamma}_s$ as the UK. This finding, on the one hand, evidently shows that the occurrence of the inverse inflation responses not only relies on this ratio, but also on the degree of forward-lookingness and price stickiness. On the other hand, it is consistent with the results in Barth and Ramey (2001) and Hanson (2004), who report that the price puzzle, in fact, only emerges in the pre-1980 time period, while it vanishes for the post-1980 period, for which we estimated the interest-rate-augmented Phillips curve.

The empirical evidence cited above indicates that our simulation exercise cannot fully account for cross-country differences in inverse inflation responses. This, however, is hardly surprising, given that the model is set up in a simple way to disclose the main mechanisms and that the households' behavior and policy are held constant. Nevertheless, the simulations carried out in this section indicate that the strength of the cost channel, as found in the data, is able to affect inflation responses to interest rate shocks in a substantial way. Moreover, the impact of estimated direct interest rate effects on the inflation response to monetary policy shocks can strongly vary and can even lead to opposite inflation reactions across G7 countries.

5 Conclusion

When firms hold working capital, interest rate changes can affect their costs and, therefore, their pricing decisions. This supply side mechanism, which gives rise to the cost channel of monetary transmission, implies that the deflationary impact of an interest rate innovation is weakened, whereas the negative output response is strengthened. While VAR-based studies already seem to indicate the relevance of the cost channel, this paper applies a structural approach, i.e., a marginal cost based Phillips curve, to assess the impact of the cost channel on short-run movements in inflation. Thereby, we found that estimated direct cost effects of short-run nominal interest rates significantly contribute to inflation dynamics in the majority of G7 countries. Our findings indicate that the cost channel is more pronounced in the UK and the US and less so in continental European countries and in Japan, which corresponds to evidence on differences in the financial structure and on the degree of interest rate pass-through. The estimates of the aggregate supply relation are further applied to assess the impact of direct interest rate effects for inflation responses to interest rate shocks in a simple dynamic general equilibrium model. Simulations reveal that the cost channel can substantially dampen inflation responses, and is even able to account for inverse inflation reactions, which relate to the price puzzle often found in monetary

VARs. Thus, our analysis points at substantial direct interest rate effects on short-run inflation dynamics, indicating that the cost channel is non-negligible for the assessment of monetary transmission.

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Appendix A: The general equilibrium model

Households and the public sector In this appendix, we present the problem of households and the public sector, which – together with the supply side of the model presented in Section 2 – complete the description of a dynamic general equilibrium model, which is employed to analyze the impact of working capital costs on monetary transmission. We assume that there is a continuum of identical and infinitely lived households of measure one. The objective of a representative household is given by:

$$E_0 \sum_{t=0}^{\infty} \beta^t \left[(1-\sigma)^{-1} c_t^{1-\sigma} - (1+\sigma_l)^{-1} l_t^{1+\sigma_l} \right], \quad \text{with } \sigma,\, \sigma_l > 0, \tag{14}$$

where c_t denotes consumption, l_t working time, $\beta \in (0,1)$ the subjective discount factor, and E_0 the expectation operator conditional on the information in period 0. At the beginning of period t, households are endowed with cash M_t and government bond holdings B_t. Households are further endowed with raw materials x_t, which they supply inelastically to firms. Their endowment with x_t is assumed to be exogenously given, such that variations in x_t, which are disregarded in this paper, might be induced by endowment shocks. Before households enter the goods market, they deposit funds D_t at financial intermediaries. Consumption expenditures are restricted by the following liquidity constraint:

$$P_t c_t \leq M_t - D_t + P_t w_t l_t + P_t q_t x_t + P_t \tau_t, \tag{15}$$

where q_t denotes the real commodity price, w_t the real wage rate, P_t the aggregate price level and τ_t a lump-sum government transfer. As in Christiano et al. (1997), cash holdings net of deposits and factor earnings can be used as a means of payment. The representative household owns the firms and the intermediaries and receives the respective profits Δ^f and Δ^b. Its budget constraint is given by

$$B_{t+1} + M_{t+1} + D_t + P_t c_t \leq (1+i_t) B_t + M_t + (1+i_t^d) D_t + P_t w_t l_t + P_t q_t x_t + P_t \tau_t + \Delta_t^f + \Delta_t^b, \tag{16}$$

where i (i^d) denotes the nominal interest rate on bonds (deposits). Maximizing the objective (14) subject to the cash constraint (15), the budget constraint (16), and a no-Ponzi-game condition for given initial values M_0 and B_0, leads to the following first order conditions:

$$\begin{aligned} & c_t^{-\sigma} = \lambda_t + \lambda_t^c, \quad (\lambda_t + \lambda_t^c)\, w_t = l_t^{\sigma_l}, \\ & \frac{1}{\beta}\lambda_t = E_t \left[\frac{1}{\pi_{t+1}} \left(\lambda_{t+1} + \lambda_{t+1}^c \right) \right], \quad \frac{1}{\beta}\lambda_t = E_t \left[\frac{1+i_{t+1}}{\pi_{t+1}} \lambda_{t+1} \right], \quad i_t^d \lambda_t = \lambda_t^c, \end{aligned} \tag{17}$$

and $0 = \eta_t(m_t - d_t + w_t l_t + \tau_t - c_t)$ with $\eta_t \geq 0$ and $m_t - d_t + w_t l_t + \tau_t - c_t \geq 0$, the budget constraint (16) holding with equality and the transversality condition $\lim_{i\to\infty} E_t \lambda_{t+i} \beta^{t+i} (B_{t+i} + M_{t+i})/P_{t+i} = 0$, where λ denotes the shadow price of wealth and λ^c the multiplier on the cash constraint (15). Hence, households relate a unit of consumption in period $t+s$ to a unit of consumption in period t by the stochastic discount factor $\eta_{t,t+s} = \beta^s (\lambda_{t+s} + \lambda_{t+s}^c)(\lambda_t + \lambda_t^c)^{-1} \pi_{t+s}^{-1}$.

The public sector consists of a central bank and a fiscal authority. The central bank is assumed to set the nominal interest rate $R_t = 1 + i_t$ according to the feedback rule $R(\pi_t, \varepsilon_t) = \mathcal{R}\pi_t^{\rho_\pi} \exp(\varepsilon_t)$, where $\rho_\pi > 0$ and ε_t, the innovations, are assumed to have an expected value of zero and to be serially uncorrelated. The support of ε_t is further restricted to be small enough, such that the central bank can always choose $\mathcal{R}$ to ensure that $R_t \geq 1$. Log-linearizing this feedback rule, gives the following simple interest rate feedback rule:

$$\widehat{R}_t = \rho_\pi \widehat{\pi}_t + \varepsilon_t. \tag{18}$$

The flow budget constraint of the public sector is given by $M_{t+1} - M_t + B_{t+1} - R_t B_t = P_t \tau_t$. Fiscal policy is assumed to ensure public sector solvency, i.e., to satisfy $\lim_{i\to\infty}(B_{t+i} + M_{t+i})\Pi_{v=1}^{i} (1 + i_{t+v})^{-1} = 0$. In particular, we assume that the net supply of government bonds equals zero.

The equilibrium is characterized by the set of conditions (17), the cash-constraint, the optimal pricing condition approximated by (3), (4), and aggregate production, the banks' first order condition, the policy rule, and the aggregate resource constraint. For the local analysis of the model, we apply a log-linear approximation of the equilibrium condition at the steady state with an inflation rate satisfying $\overline{\pi} = \beta\overline{R}$. A rational expectations equilibrium of the linear approximation to the model at the steady state is then a set of sequences $\{\widehat{\pi}_t,\ \widehat{y}_t,\ \widehat{R}_t\}_{t=0}^{\infty}$ satisfying $\widehat{\pi}_t = \gamma_f E_t \widehat{\pi}_{t+1} + \gamma_b \widehat{\pi}_{t-1} + \gamma_R \widehat{R}_t + \gamma_y \widehat{y}_t$, and $\sigma \widehat{y}_t = \sigma E_t \widehat{y}_{t+1} - \widehat{R}_t + E_t \widehat{\pi}_{t+1}$, where $\gamma_R \equiv \chi(1 + \psi_R)$ and $\gamma_y \equiv \chi(\sigma + \frac{\sigma_l + \alpha}{1-\alpha})$, and (18).

Solution In order to derive the solution for the model in (11)-(13), it is reduced by eliminating the nominal interest rate with the policy rule (18) leading to the following conditions in inflation and output:

$$\sigma \widehat{y}_t = \sigma E_t \widehat{y}_{t+1} - \rho_\pi \widehat{\pi}_t - \varepsilon_t + E_t \widehat{\pi}_{t+1} \tag{19}$$

$$\widehat{\pi}_t (1 - \gamma_R \rho_\pi) = \gamma_f E_t \widehat{\pi}_{t+1} + \gamma_b \widehat{\pi}_{t-1} + \gamma_R \varepsilon_t + \gamma_y \widehat{y}_t. \tag{20}$$

Note that the model (19)-(20) is not entirely forward-looking, as the lagged inflation rate enters the aggregate supply constraint. For the equilibrium sequences of the model to be stable, there must be one stable eigenvalue that can be assigned to this relevant predetermined variable, $\widehat{\pi}_{t-1}$. It is straightforward to show that an active interest rate rule, $\rho_\pi > 1$, is sufficient to ensure the existence of a stable eigenvalue, which lies between zero and one.[16] Thus, for $\rho_\pi > 1$ the fundamental (minimum state) solution of the model, which takes the generic form

$$\widehat{\pi}_t = \delta_\pi \widehat{\pi}_{t-1} + \delta_{\pi e} \varepsilon_t, \quad \widehat{y}_t = \delta_{y\pi} \widehat{\pi}_{t-1} + \delta_{ye} \varepsilon_t, \tag{21}$$

is stable and non-oscillatory, $\delta_\pi \in (0, 1)$, if $\rho_\pi > 1$. Eliminating the endogenous variables with the solution form (21) in (19)-(20) and using that $E_t \widehat{\pi}_{t+1} = \delta_\pi \widehat{\pi}_t$ and $E_t \widehat{y}_{t+1} = \delta_\pi \widehat{y}_t$,

[16] The proof can be made available upon request by the authors.

leads to the following conditions for the solution coefficients $\delta_{y\pi}$, δ_{ye}, $\delta_{\pi e}$ and δ_{π} :

$$\delta_{y\pi} = \frac{(\delta_{\pi} - \rho_{\pi})\,\delta_{\pi}}{\sigma\,(1-\delta_{\pi})}, \qquad \delta_{ye} = \frac{\delta_{\pi e}\,(\sigma\delta_{y\pi} - \rho_{\pi} + \delta_{\pi}) - 1}{\sigma}, \tag{22}$$

$$\delta_{\pi} = \frac{\gamma_b + \gamma_y\delta_{y\pi} + \gamma_f\delta_{\pi}^2}{1-\gamma_R\rho_{\pi}}, \qquad \delta_{\pi e} = -\frac{\gamma_R + \gamma_y\delta_{ye}}{\gamma_f\delta_{\pi} - (1-\gamma_R\rho_{\pi})}. \tag{23}$$

Eliminating $\delta_{y\pi}$ in the first condition in (23) gives the cubic characteristic equation in δ_{π} : $0 = \gamma_b + \gamma_y \frac{(\delta_{\pi}-\rho_{\pi})\delta_{\pi}}{\sigma(1-\delta_{\pi})} + \gamma_f\delta_{\pi}^2 - (1-\gamma_R\rho_{\pi})\,\delta_{\pi}$. For given values for the parameters σ, σ_l, ρ_{π}, α, and $\gamma_y = \chi(\sigma + \frac{\sigma_l+\alpha}{1-\alpha})$, and $\chi = \hat{\gamma}_s$, $\gamma_b = \hat{\gamma}_b$, $\gamma_f = \hat{\gamma}_b$, and $\gamma_R = \hat{\gamma}_R$, the cubic equation can be solved numerically. For our parameterizations, there is always exactly one stable root of this equation that lies between zero and one, which is assigned to δ_{π}. Given this value for δ_{π}, one can easily determine the remaining coefficient $\delta_{y\pi}$, δ_{ye} and $\delta_{\pi e}$ from (22) and (23). With these values for the coefficients $\delta_{y\pi}$, δ_{ye}, $\delta_{\pi e}$ and δ_{π}, impulse response functions are then computed with the fundamental solution (21).

Appendix B: Results for the alternative specifications

The tables provided in this appendix present the results for the alternative empirical specification discussed in Section 3.2. They are included in the paper for the convenience of the referees and not (necessarily) intended for publication. In any case, they can be made available upon request from the authors.

Table 4: Simultaneous Estimation Results (GDP-deflator)

		$\hat{\gamma}_f^{sim}$	$\hat{\gamma}_b^{sim}$	$\hat{\gamma}_w^{sim}$	$\hat{\gamma}_R^{sim}$	$\hat{\rho}_\pi$	$J-Test$
Canada	IR1	0.53 (0.013)	0.43 (0.014)	0.009 (0.004)	0.026 (0.006)	0.815 (0.053)	0.14
	IR2	0.54 (0.009)	0.44 (0.011)	0.008 (0.003)	0.019 (0.006)	0.887 (0.041)	0.19
France	IR1	0.34 (0.065)	0.53 (0.044)	0.077* (0.042)	0.074 (0.032)	0.518 (0.089)	0.33
	IR2	0.62 (0.038)	0.39 (0.038)	$-0.015^{\#}$ (0.023)	$-0.017^{\#}$ (0.017)	0.435 (0.065)	0.21
Italy	IR1	0.47 (0.034)	0.50 (0.035)	$-0.017^{\#}$ (0.011)	0.059 (0.030)	0.685 (0.060)	0.14
	IR2	0.44 (0.039)	0.49 (0.041)	$-0.016^{\#}$ (0.012)	0.092 (0.038)	0.714 (0.062)	0.10
UK	IR1	0.54 (0.046)	0.41 (0.053)	$0.021^{\#}$ (0.029)	0.030 (0.014)	0.600 (0.16)	0.26
	IR2	0.54 (0.040)	0.43 (0.045)	$0.009^{\#}$ (0.024)	0.026 (0.013)	0.575 (0.090)	0.36
US	IR1	0.53 (0.066)	0.33 (0.091)	$0.047^{\#}$ (0.032)	0.079* (0.045)	1.064 (0.113)	0.48
	IR2	0.54 (0.071)	0.29 (0.108)	0.063* (0.038)	0.098* (0.052)	1.104 (0.183)	0.29

Notes: Table 4 provides two sets of estimates using simultaneous GMM estimations, where the aggregate supply constraint (7) is jointly estimated with a simple interest rate feedback rule. The set of instruments includes two lags of inflation, real unit labour costs and interest rates. We additionally allow for two lags of real commodity prices in all countries except for the US and France, where the lag length equals one and three, respectively. Estimations were not carried out for Germany and Japan, since in both countries the single equation estimations based on (8) did not produce significant interest rate coefficients. The first set corresponds to an interest rate rule featuring current inflation (IR1) $\widehat{R}_t = \rho_\pi \widehat{\pi}_t + \varepsilon_t$, and the second to a forward-looking rule (IR2), $\widehat{R}_t = \rho_\pi E_t \widehat{\pi}_{t+1} + \varepsilon_t$. Figures in round brackets denote standard errors. The J-Test describes a test statistic for the null hypothesis that the overidentifying restrictions are satisfied. For the latter test p-values are reported; all estimated coefficients are statistically significant at the 5 percent level except those marked with #, which are not statistically significant at the 10 percent level. Coefficients that are only significant at the 10 percent level are marked with *.

Table 5: Estimates of the interest-rate-augmented Phillips-Curve with real factor prices (GDP-based)

	$\hat{\gamma}_f^{com}$	$\hat{\gamma}_b^{com}$	$\hat{\gamma}_w^{com}$	$\hat{\gamma}_R^{com}$	$\hat{\gamma}_q^{com}$	$J-Test$
Canada	0.42 (0.055)	0.53 (0.049)	0.004 (0.002)	0.060 (0.022)	0.002* (0.001)	0.63
France	0.43 (0.049)	0.50 (0.036)	$0.006^{\#}$ (0.006)	0.101 (0.031)	$0.001^{\#}$ (0.003)	0.32
Germany	0.45 (0.076)	0.56 (0.036)	$0.001^{\#}$ (0.002)	$0.022^{\#}$ (0.032)	$0.001^{\#}$ (0.001)	0.65
Italy	0.37 (0.046)	0.50 (0.038)	0.004 (0.002)	0.134 (0.054)	0.007 (0.003)	0.64
Japan	0.68 (0.115)	0.30 (0.080)	$0.001^{\#}$ (0.005)	$0.02^{\#}$ (0.099)	$-0.001^{\#}$ (0.001)	0.46
UK	0.48 (0.032)	0.49 (0.022)	$0.001^{\#}$ (0.001)	0.046 (0.016)	0.004 (0.002)	0.62
US	0.39 (0.055)	0.61 (0.058)	0.030 (0.007)	0.049 (0.016)	$0.001^{\#}$ (0.001)	0.76

Notes: Table 5 presents GMM estimation results for the specification (10), where real factor prices are considered as explanatory variables for current inflation. The set of instruments includes four lags of inflation, the real wage rate, the nominal interest rate, and real commodity prices for all countries. Figures in round brackets denote standard errors. The J-Test describes a test statistic for the null hypothesis that the overidentifying restrictions are satisfied. For the latter test p-values are reported; all estimated coefficients are statistically significant at the 5 percent level except those marked with #, which are not statistically significant at the 10 percent level. Coefficients that are only significant at the 10 percent level are marked with *.

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2007

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2006

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2005

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2004

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