Identifying the Influences of Nominal and Real Rigidities in Aggregate Price-Setting Behavior

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Abstract

We formulate a generalized price-setting framework that incorporates staggered contracts of multiple durations and that enables us to directly identify the influences of nominal vs. real rigidities. Using German macroeconomic data over the period 1975Q1 through 1998Q4 to estimate this framework, we find that the data is well-characterized by a truncated Calvo-style distribution with an average duration of about two quarters. We also find that new contracts exhibit very low sensitivity to marginal cost, corresponding to a relatively high degree of real rigidity. Finally, our results indicate that backward-looking behavior is not needed to explain the aggregate data, at least in an environment with a stable monetary policy regime and a transparent and credible inflation objective.

1 Introduction

Micro-founded models of price-setting behavior are essential for understanding aggregate inflation dynamics and for evaluating the performance of alternative monetary policy regimes.[1] Both nominal and real rigidities play a crucial role in determining the particular implications of these models; thus, a large body of empirical research has been oriented towards gauging the frequency of price adjustment, the sensitivity of price revisions to demand and cost pressures, and the prevalence of indexation or rules of thumb.[2]

The recent empirical literature has mainly focused on estimating variants of the New Keynesian Phillips Curve (NKPC), which can be derived under the assumption that price contracts have random duration with a constant hazard rate.[3] Nevertheless, since the slope of the NKPC depends on the mean duration of price contracts as well as potential sources of real rigidity, the underlying structural parameters cannot be separately identified using this framework.[4] Furthermore, while most studies have obtained highly significant estimates of the coefficient on lagged inflation, no consensus has been reached about whether to interpret these results as reflecting backward-looking price-setting behavior or gradual learning about occasional shifts in the monetary policy regime.[5]

In this paper, we formulate a generalized price-setting framework that incorporates staggered contracts of multiple durations and that enables us to directly identify the influences of nominal vs. real rigidities. In analyzing contracts with random duration, we assume that every firm which resets its price faces the same *ex ante* probability distribution of contract duration, as in Calvo (1983), but we do not impose any restrictions on the shape of the hazard function. We also consider specifications in which each firm signs price contracts

[1] See Yun (1996), Goodfriend and King (1997), Rotemberg and Woodford (1997), Clarida, Galí, and Gertler (1999), and Woodford (2003).

[2] The importance of combining nominal and real rigidities has been emphasized by Ball and Romer (1990), Chari, Kehoe, and McGratten (2002), and Christiano, Eichenbaum, and Evans (2004).

[3] Following Galí and Gertler (1999) and Sbordone (2002), the literature has become too voluminous to be enumerated here; recent examples include Lindé (2001), Neiss and Nelson (2002), and Sondergaard (2003).

[4] See Galí, Gertler, and López-Salido (2001) and Eichenbaum and Fisher (2004).

[5] For example, Galí et al. (2001) considers a specification with rule-of-thumb price-setters, while Erceg and Levin (2003) show that the lagged inflation term in the hybrid Phillips curve can be generated by rational agents who use signal extraction to learn about shifts in the central bank's inflation objective.

with a fixed and known duration, as in Taylor (1980), but this duration is permitted to vary across different groups of firms. Finally, our framework encompasses two sources of real rigidity: firm-specific factors, and non-constant elasticity of demand.

Our empirical analysis utilizes German macroeconomic data over the period 1975Q1 through 1998Q4–a dataset that provides a virtually ideal setting for determining the structural characteristics of price-setting behavior in the context of a stable monetary policy regime. In particular, the Bundesbank maintained a transparent and, one may presume, reasonably credible medium-term inflation objective that declined gradually from 5 percent in 1975 to 2 percent in 1984, and remained essentially constant thereafter. Thus, our investigation proceeds by fitting the deviations of actual inflation from the Bundesbank's medium-term inflation objective.

Using simulation-based indirect inference methods to estimate the model, we find that price-setting behavior is well-characterized by a truncated Calvo-style distribution with an average duration of about two quarters. We also find that new price contracts exhibit very low sensitivity to marginal cost, corresponding to a relatively high degree of real rigidity involving *both* firm-specific inputs and strong curvature of the demand function. Finally, we confirm that the estimated model is not rejected by tests of overidentifying restrictions, and that the implied autocorrelations are virtually indistinguishable from those of an unrestricted vector autoregression.[6] Evidently, backward-looking behavior (due to informational constraints or rule-of-thumb price-setting) is not needed to explain the aggregate data, at least in the context of a stable policy regime with a transparent and credible inflation objective.

Our empirical findings regarding the frequency of price adjustments are broadly consistent with recent evidence from firm-level surveys and micro price records.[7] The microeco-

[6] Mash (2003) uses micro evidence to calibrate a similar price-setting framework with a generalized hazard function, and shows that the calibrated model can roughly match empirical autocorrelations.

[7] Survey evidence has been obtained by Blinder, Canetti, Lebow, and Rudd (1998), Hall, Walsh, and Yates (2000), Apel, Friberg, and Hallsten (2001), and Fabiani, Gattulli, and Sabbatini (2004). For recent evidence from micro price records, see Chevalier, Kashyap, and Rossi (2003), Bils and Klenow (2004), Aucremanne and Dhyne (2004), and Dias, Dias, and Neves (2004). Additional references and discussion may be found in Taylor (1999).

nomic evidence also provides some indirect support for our focus on time-dependent rather than state-dependent specifications of price-setting behavior.[8]

The remainder of this paper is organized as follows. Section 2 presents the generalized price-setting framework. Section 3 describes the data used in our analysis, while Section 4 reviews the estimation methodology. Section 5 reports the estimation results, confirms that these results are reasonably robust to alternative proxies for real marginal cost, and documents the importance of accounting for the evolution of the Bundesbank's medium-term inflation objective. Section 7 concludes.

2 The Generalized Price-Setting Framework

In this section we formulate a generalized price-setting framework that incorporates staggered nominal contracts of multiple durations. Within this framework, we allow price contracts to have either random duration á la Calvo (1983) or fixed duration á la Taylor (1980). In the former case, we assume that every firm which resets its price faces the same *ex ante* probability distribution of contract duration, without imposing any restrictions on the shape of the hazard function. In the latter case, every price contract has a fixed and known duration which varies across different groups of firms.

Our framework encompasses two sources of real rigidity. First, following Kimball (1995), each firm's demand may exhibit a high degree of curvature (approximating a "kinked demand curve") as a function of the firm's price deviation from the average price level. Thus, when a firm is resetting its price contract, its optimal price will be relatively less sensitive to changes in the firm's marginal cost. Second, the presence of fixed firm-specific inputs causes each firm's marginal cost to vary with its level of output and hence dampens the sensitivity of new contract prices to an aggregate shock. For example, in considering a price hike in response to a particular shock, the firm recognizes that lower demand will reduce

[8] Caplin and Leahy (1997) and Dotsey, King, and Wolman (1999) have developed models of state-dependent price-setting, while Klenow and Kryvstov (2004) provide recent evidence on its limited role in generating aggregate inflation variability.

its marginal cost, thereby partially offsetting the original rationale for raising its price.

Henceforth we will use the term "capital" to refer to the fixed factor in production, while the variable factor will be referred to as "labor." Nevertheless, it should be emphasized that the fixed factor could include land as well as any overhead labor that cannot easily be adjusted in the short run. Furthermore, while our analysis abstracts from the influence of endogenous capital accumulation, the results of Eichenbaum and Fisher (2004) indicate that the degree of real rigidity is quantitatively similar for specifications with fixed capital and for specifications with an empirically reasonable magnitude of adjustment costs for investment.[9]

2.1 The market structure

Consider a continuum of monopolistically competitive firms indexed by $f \in [0,1]$, each of which produces a differentiated good $Y_t(f)$ using the following production function:

$$Y_t(f) \quad = \quad A_t \bar{K}(f)^{\alpha} L_t(f)^{1-\alpha}. \tag{1}$$

Note that all firms have the same level of total factor productivity, A_t. To ensure symmetry in the deterministic steady state, we also assume that every firm owns an identical capital stock, $\bar{K}(f) = \bar{K}$.

A distinct set of perfectly competitive aggregators combine all of the differentiated products into a single final good, Y_t, using the following technology:

$$\int_0^1 G(Y_t(f)/Y_t)\, df = 1, \tag{2}$$

where the function $G(\cdot)$ is increasing and strictly concave with $G(1) = 1$. This formulation generalizes the Dixit-Stiglitz formulation, for which $G(x) = x^{\eta/(\eta-1)}$ for $\eta > 1$.

Under these assumptions, each firm f faces the following implicit demand curve for its output as a function of its price $P_t(f)$ relative to the price of the final good, P_t:

$$G'(Y_t(f)/Y_t) \quad = \quad \left(\frac{P_t(f)}{P_t}\right) \int_0^1 (Y_t(z)/Y_t) G'(Y_t(z)/Y_t)\, dz. \tag{3}$$

[9] Optimal price setting with firm-specific capital accumulation has recently been analyzed by Sveen and Weinke (2003) and Woodford (2004); see also Altig et al. (2004).

The concavity of $G(\cdot)$ ensures that the demand curve is downward-sloping; that is, $dY_t(f)/dP_t(f) < 0$.

Finally, the firm's real marginal cost function $MC_t(f)$ is given as follows:

$$MC_t(f) \quad = \quad \frac{W_t}{(1-\alpha)P_t A_t \bar{K}(f)^{\alpha} L_t(f)^{-\alpha}}, \tag{4}$$

where W_t denotes the nominal wage rate.

2.2 The duration of price contracts

We assume that the prices for the differentiated goods, $P_t(f)$, are determined by staggered nominal contracts with a maximum duration of J periods. For $j = 1, \ldots, J$, let ω_j denote the fraction of price contracts that have a duration of j periods, where $\omega_j \geq 0$ and $\sum_{j=1}^{J} \omega_j = 1$.

In the case of random contract durations, every firm has the same hazard function, which determines the probability that the firm is permitted to reset its price. The specification here generalizes that of Calvo (1983), because the probability of a price revision can depend on the number of periods that the existing contract has been in effect. Specifically, a firm f whose contract has been in effect for k periods faces the probability $\sum_{j=1}^{k} \omega_j$ of receiving permission to reset its contract in the current period, where, as noted above, $\omega_j \geq 0$ and $\sum_{j=1}^{J} \omega_j = 1$. With probability $\sum_{j=k+1}^{J} \omega_j$, this firm is not permitted to reset its contract in period t, and its price remains unchanged; that is, $P_t(f) = P_{t-1}(f)$.

In the case of fixed contract durations, ω_j denotes the fraction of firms that sign price contracts with a duration of j periods ($j = 1, \ldots, J$), where again $\omega_j \geq 0$ and $\sum_{j=1}^{J} \omega_j = 1$. For each contract length j, an equi-proportionate fraction ω_j/j of firms reset their contracts in any given period t. For each firm which does not reset its contract in period t, its price remains unchanged, that is $P_t(f) = P_{t-1}(f)$.

In formal terms, the distribution of fixed-duration contracts can be represented as follows. Let $\{\,\Omega_1, \ldots, \Omega_J\,\}$ denote a partition of the continuum of monopolistically competitive firms $\Omega = [\,0, 1\,]$ into subintervals with $\Omega_j = [\,s_{j-1}, s_j\,)$ for $j = 1, \ldots, J-1$ together with

$\Omega_J = [\, s_{J-1}, s_J \,]$, satisfying $0 = s_0 \leq s_1 \leq \cdots < s_J = 1$; and let $\omega_j = \mu(\Omega_j)$ denote the measure of the subinterval Ω_j. The individual firms in Ω_j may be indexed so that every firm with index $f \in [\, s_{j-1}, s_{j-1}+(s_j - s_{j-1})/j \,)$ resets its contract price whenever the period t is evenly divisible by j; similarly firms with index $f \in [\, s_{j-1} + (s_j - s_{j-1})/j, s_{j-1} + 2(s_j - s_{j-1})/j \,)$ reset prices during periods in which modulus$(t, j) = 1$, and so forth.

2.3 The optimal price-setting decision

In period t, each firm resetting its contract chooses its new price $P_t(f)$ to maximise the firm's expected discounted profits over the life of the contract,

$$\mathrm{E}_t \left[\sum_{j=0}^{J-1} \chi_j \lambda_{t,t+j} \left(P_t(f) Y_{t+j}(f) - W_{t+j} L_{t+j}(f) \right) \right], \tag{5}$$

subject to the production function (1) and the implicit demand curve (3), where the stochastic discount factor $\lambda_{t,t+j}$ can be obtained from the consumption Euler equation of the representative household.

If the price contract has random duration, then the coefficient χ_j indicates the probability that the price contract will still be in effect after j periods; that is, $\chi_j = \sum_{k=j+1}^{J} \omega_k$ for $j = 0, \ldots, J-1$. In the special case of Calvo-style contracts, the firm faces a constant probability ξ of not revising its contract in any given period; thus, $\chi_j = \xi^j$, and the maximum duration $J \to \infty$.

If the price contract has a fixed duration, then the coefficient χ_j is simply an indicator function. In particular, when the contract has a duration of i periods (for $i = 1, ..., J$), then $\chi_j = 1$ for $j = 0, \ldots, i-1$ and 0 otherwise.

2.4 The log-linearization with random contract duration

We now proceed to log-linearize the pricing equation and the aggregate price identity around the deterministic steady state with zero inflation.[10] We use π_t to denote the aggregate

[10] For analysis of the log-linearization around a non-zero steady state, see Ascari (2003) for the case of random-duration contracts, and Erceg and Levin (2003) for the case of fixed-duration contracts.

inflation rate, while mc_t denotes the average real marginal cost across all firms in the economy (expressed as a logarithmic deviation from its steady-state value), and y_t denotes the logarithmic deviation of aggregate output from steady state.

In the case of random contract duration, all firms signing new contracts at date t set the same price. Thus, using x_t to denote the logarithmic deviation of the new contract price from the aggregate price level, we obtain the following expression for the log-linearised optimal price-setting equation:

$$x_t \;=\; \mathrm{E}_t \left[\sum_{j=1}^{J-1} \Phi_j \, \pi_{t+j} + \gamma \sum_{j=0}^{J-1} \phi_j \, mc_{t+j} \right], \tag{6}$$

where β denotes the household's discount factor, and the weights satisfy $\phi_j = \beta^j \sum_{k=j+1}^{J} \omega_k / \left(\sum_{i=1}^{J} \sum_{k=0}^{i-1} \beta^k \omega_i \right)$ and $\Phi_j = \sum_{k=j}^{J-1} \phi_k$.

The coefficient γ in equation (6) determines the sensitivity of new price contracts to aggregate real marginal cost. In particular, as shown by Eichenbaum and Fisher (2004), this coefficient can be expressed as the product of two components; that is, $\gamma = \gamma_d \, \gamma_{mc}$, where:

$$\gamma_d \;=\; \frac{1 + G''(1)/G'(1)}{2 + G'''(1)/G''(1)} \tag{7}$$

$$\gamma_{mc} \;=\; \frac{1}{1 + \frac{\alpha}{1-\alpha} \eta \gamma_d}, \tag{8}$$

and the steady-state elasticity of demand $\eta = G'(1)/G''(1)$. The coefficient γ_d is solely related to the curvature of the demand curve; in the special case with constant elasticity of demand, $\gamma_d = 1$. The coefficient γ_{mc} reflects the degree to which the firm's relative price influences its marginal cost; thus, in the special case with no fixed factors, $\alpha = 0$ and $\gamma_{mc} = 1$.

The log-linearized aggregate price identity can be expressed as follows:

$$\sum_{j=0}^{J-1} \psi_j \, x_{t-j} \;=\; \sum_{j=0}^{J-2} \Psi_{j+1} \, \pi_{t-j}, \tag{9}$$

where the weights satisfy $\psi_j = \sum_{k=j+1}^{J} \omega_k / \left(\sum_{i=1}^{J} \sum_{k=0}^{i-1} \omega_i \right)$ and $\Psi_j = \sum_{k=j}^{J-1} \psi_k$.

2.5 The log-linearization with fixed contract duration

For the case of fixed-duration contracts, let x_t^i indicate the logarithmic deviation of the new contract price of duration i from the aggregate price level. Then the log-linearised price-setting equation can be expressed as follows:

$$x_{it} = \mathrm{E}_t \left[\sum_{j=1}^{i-1} \tilde{\Phi}_{ij}\, \pi_{t+j} + \gamma \sum_{j=0}^{i-1} \tilde{\phi}_{ij}\, mc_{t+j} \right], \tag{10}$$

where $\tilde{\phi}_{ij} = \beta^j / \sum_{k=0}^{i-1} \beta^k$ and $\tilde{\Phi}_{ij} = \sum_{k=j}^{i-1} \phi_{ik}$. The aggregate price level depends on all of the price contracts in effect at date t; thus, we obtain the following expression for the aggregate price identity:

$$\sum_{j=0}^{J-1} \sum_{i=j+1}^{J} \frac{\omega_i}{i}\, x_{i,t-j} = \sum_{j=0}^{J-2} \sum_{i=j+2}^{J} \frac{\omega_i}{i}\, \pi_{t-j}. \tag{11}$$

The empirical results reported below for fixed-duration contracts are based on analysis of equations (10) and (11). In addition, for purposes of comparison with some of the earlier literature on fixed-duration contracts, the Appendix provides results using the simplifying assumption of negligible variation across new price contracts.

3 The Data

In estimating a structural price-setting framework, it is essential to avoid spurious influences due to shifts in the monetary policy regime. In cases where the shift is not transparent or credible, price-setting behavior may appear to be backward-looking when in fact private agents are using optimal filtering to determine the policy regime (cf. Erceg and Levin 2003). Even a transparent and credible change in the central bank's inflation objective tends to raise the measured degree of inflation persistence unless the shift is explicitly taken into account in the estimation procedure (cf. Levin and Piger 2004).

Thus, German macroeconomic data for 1975-1998 provides a virtually ideal setting for determining the structural characteristics of price-setting, because the Bundesbank maintained a reasonably transparent and credible medium-term inflation objective over this period. In particular, in the process of deriving money growth targets (starting in 1975), the Bundesbank regularly stated its assumptions regarding the level of inflation over the medium run, set in the broader context of the ultimate goal of price stability. During the late 1970s and early 1980s, the medium-term assumption was referred to as the "unavoidable" level of inflation, reflecting the Bundesbank's willingness to attain price stability over a longer horizon rather than inducing a sudden sharp contraction in real economic activity.[11] After reaching the neighborhood of price stability in the mid-1980s, the Bundesbank referred to its inflation assumption as the "medium-term price norm".[12]

The upper-left panel of **Figure 1** depicts the evolution of actual inflation and the Bundesbank's medium-term inflation objective over the period 1974-1998. At the beginning of the sample period, GDP price inflation was at a transitory peak of about 8 percent in the wake of the collapse of the Bretton Woods regime and the first OPEC oil price shock. Inflation subsequently stabilized around the Bundesbank's medium-term inflation objective

[11] As shown by Erceg and Levin (2003), even a transparent and credible disinflation causes a transitory recession in a model with four-quarter Taylor-style wage and price contracts. In contrast, disinflations can be costless in models with Calvo-style contracts.

[12] Further details regarding the Bundesbank's monetary policy strategy may be found in Schmid (1999) and Gerberding, Seitz, and Worms (2004).

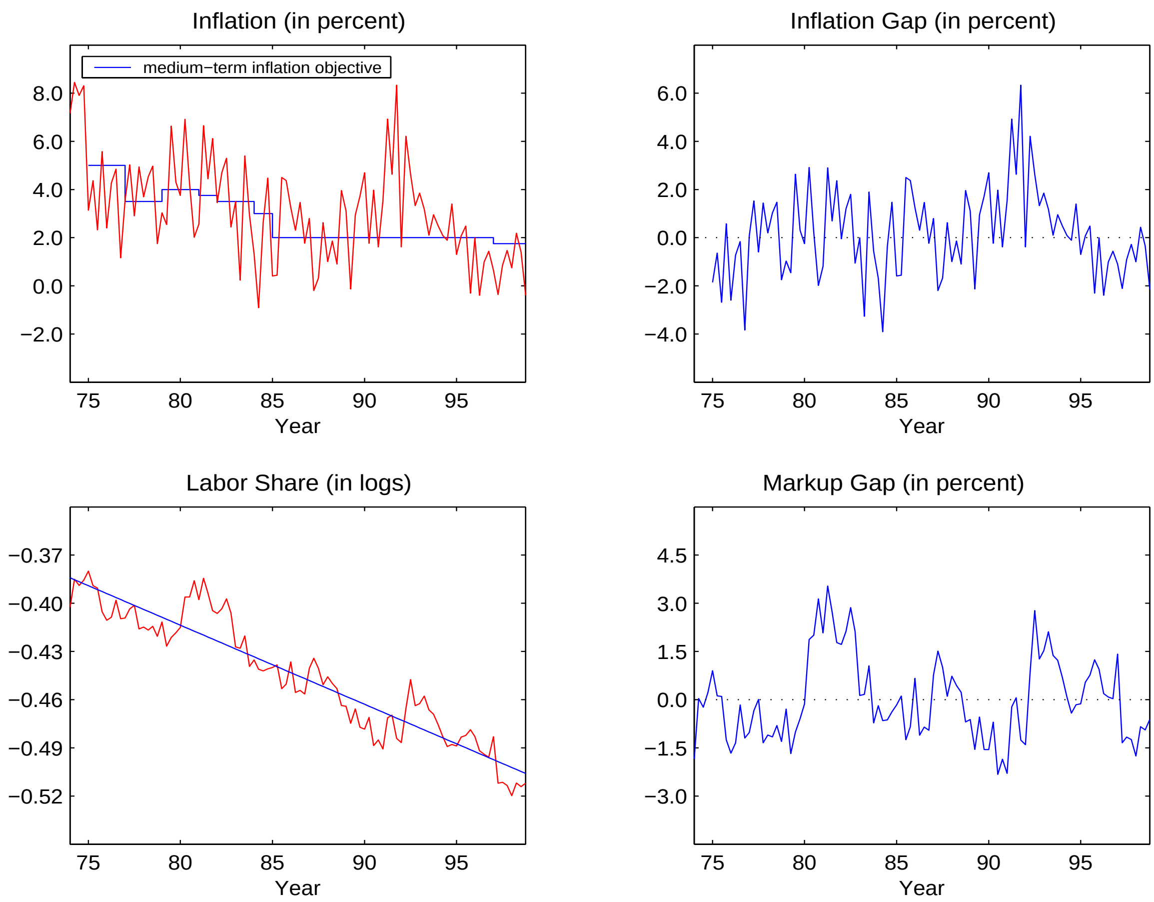
of about 5 percent, and then declined fairly gradually through the late 1970s and early 1980s, roughly in parallel with reductions in the Bundesbank's medium-term objective. From about 1985 through the advent of the European Monetary Union, the inflation objective remained essentially constant at 2 percent; actual inflation exhibited an average level fairly close to this objective, with only one large deviation in the early 1990s during the process of German reunification. Our empirical investigation proceeds by fitting the deviations of actual inflation from the Bundesbank's medium-term inflation objective; this "inflation gap" is shown in the upper-right panel of Figure 1.

The labor share serves as our benchmark proxy for real marginal cost. In measuring the labor share, it is important to account for the significant role of self-employed workers in the German economy. In the absence of direct measures of labor compensation for self-employed workers, we follow the fairly standard approach of computing the labor share by taking the compensation of employees (which does not include self-employed workers), multiplying this figure by the ratio of total employment (including self-employed workers) to the number of employees, and then dividing by nominal GDP. In effect, this procedure uses the average compensation rate of employees to impute the labor compensation of self-employed workers.

The lower-left panel of Figure 1 depicts the evolution of the German labor share. This series exhibits a clear downward trend over the sample period, presumably reflecting gradual structural changes in the German economy. Since our analytical framework follows the standard New Keynesian view that prices adjust in response to deviations of the actual markup from a desired level, we interpret the low-frequency movement of the labor share as a deterministic trend in the desired markup. Thus, our price-setting framework is estimated using the detrended labor share–henceforth referred to as the markup gap–as depicted in the lower-right panel of Figure 1.

In performing sensitivity analysis, we consider several alternative proxies for real marginal cost, each of which is depicted in **Figure 2**. The upper-right panel shows two measures of the output gap, which have been constructed from real GDP (shown in the upper-left panel) using linear detrending and Hodrick-Prescott filtering, respectively. The

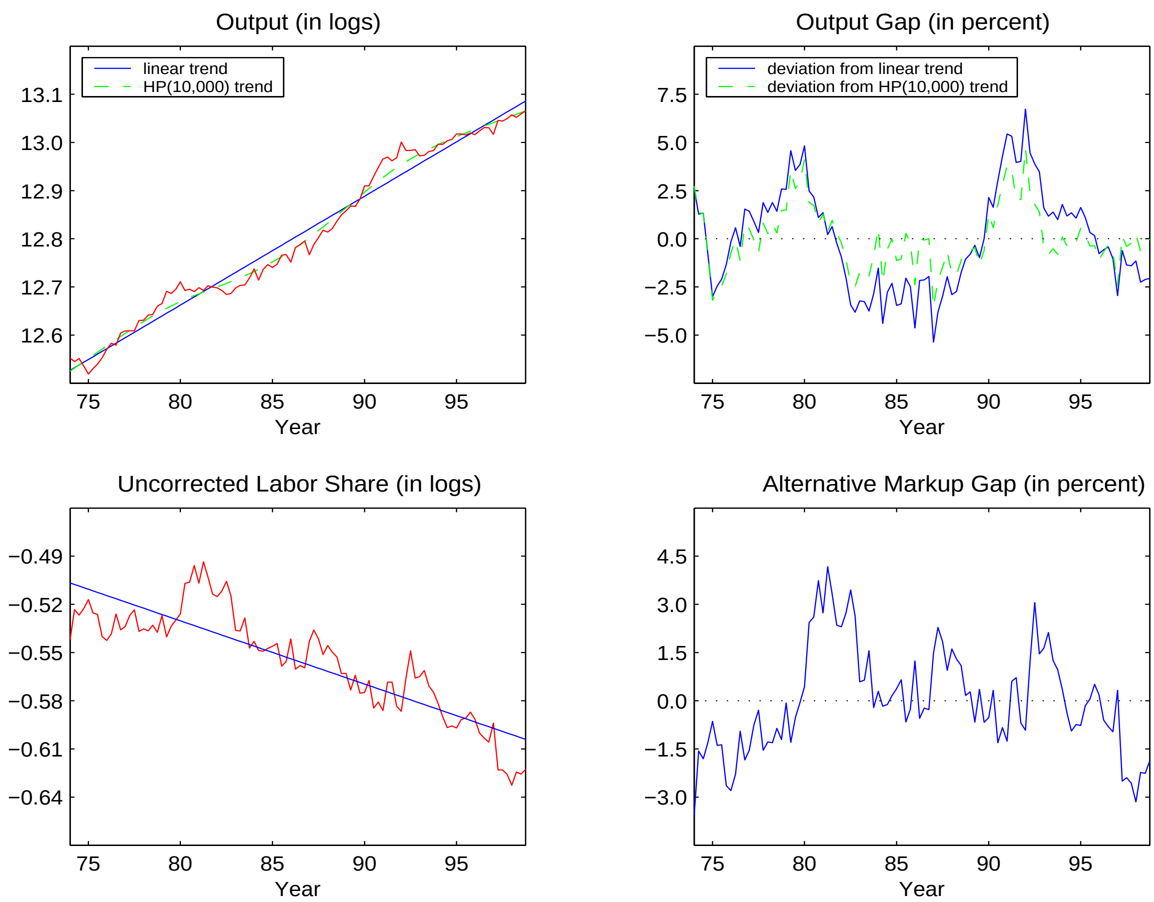
Figure 1: German Inflation and Markup Gaps, 1974-1988



Note: Inflation is measured as the annualized quarter-on-quarter change in the logarithm of the GDP price deflator. The inflation gap is defined as the deviation of inflation from the Bundesbank's medium-term inflation objective. The labor share is constructed as the ratio of total compensation (including imputed labor income of self-employed workers) to nominal GDP. The markup gap is defined as the deviation of the labor share from a linear trend.

lower-left panel depicts the ratio of employee compensation to nominal GDP. This measure–henceforth referred to as the uncorrected labor share–implicitly attributes all of the income of self-employed workers as compensation to capital rather than labor. The behavior of the linearly-detrended series (shown in the lower-right panel) is broadly similar to that of the benchmark series, but the deviation from trend is much larger in the mid-1970s; given that this deviation is not accompanied by substantial movement in inflation, we shall see below that the uncorrected labor share implies an even higher degree of real rigidity than the benchmark series.

Figure 2: Alternative Proxies for the Markup Gap



Note: Output is measured as the logarithm of real GDP. The output gap is constructed by detrending output using either a linear trend or a Hodrick-Prescott filter with a smoothing parameter of 10,000. The uncorrected labor share is the ratio of employee compensation to nominal GDP, and does not incorporate the imputed labor income of self-employed workers. The corresponding markup gap is obtained by linearly detrending the uncorrected labor share.

4 Estimation Methodology

Our empirical analysis essentially follows the approach of Coenen and Wieland (2004). In the first stage, we estimate an unconstrained VAR model that provides an empirical description of the dynamics of the inflation gap, the markup gap, and the output gap. In the second stage, we employ simulation-based indirect inference methods to estimate the structural price-setting equations, using the unconstrained VAR as the auxiliary model. In effect, this method determines the parameters of the structural model by matching its reduced form–which constitutes a constrained VAR–as closely as possible with the unconstrained VAR.[13]

In the remainder of this section, we compare our procedure with alternative approaches that have been employed in the literature, and then describe the estimation methodology in further detail.

4.1 Comparison with Alternative Approaches

Unlike most of the literature on estimating NKPCs, standard method-of-moments procedures cannot be applied to our generalized price-setting framework due to the presence of unobserved variables (namely, the new contracts signed each period). Furthermore, since each contract price depends on expected future markup gaps, we need to specify how these gaps are determined. To avoid imposing any additional restrictions, we simply take the markup gap and output gap equations from the unconstrained VAR and combine these with the structural price-setting equations; we refer to the combined set of equations as the "structural model" even though only part of the model is truly structural.[14]

Our estimation methodology has some appealing features compared with several other commonly-employed procedures. For example, one alternative approach is to specify a complete structural model and estimate its parameters by matching some of the implied

[13]The method of indirect inference was proposed by Smith (1993) and Gouriéroux, Monfort and Renault (1993); see also Gouriéroux and Monfort (1996). For a summary of the asymptotic properties of this procedure, see the appendix of the working paper version of Coenen and Wieland (2004).

[14]This limited-information approach follows Taylor (1993) and Fuhrer and Moore (1995), and is similar in spirit to the approach of Sbordone (2002).

impulse response functions (IRFs) to those of an identified VAR model.[15] In contrast, our procedure matches the implications of the structural model to those of an unconstrained VAR, thereby avoiding the need to impose potentially controversial identifying assumptions on the auxiliary model. Furthermore, our procedure essentially matches all of the sample autocorrelations and cross-correlations rather than a limited set of characteristics of the data.

Another alternative approach involves the use of full-information methods to estimate a complete structural model.[16] Nevertheless, one potential pitfall of that approach is that the price-setting parameter estimates could be sensitive to misspecifications in other aspects of the model–a particular important issue in this case due to the lack of consensus about which labor market rigidities are relevant in determining the behavior of the markup gap.

4.2 Details of the Estimation Procedure

We begin by estimating an unconstrained VAR for the inflation gap, the markup gap, and the output gap; we use T to denote the number of sample periods. Note that the vector of parameter estimates $\hat{\zeta}_T$ includes the VAR coefficients as well as the variances and contemporaneous correlations of the residuals; we use $\hat{\Sigma}_T(\hat{\zeta}_T)$ to denote the estimated covariance matrix of $\hat{\zeta}_T$. The estimated model has a lag order of 3, because this corresponds to the reduced-form VAR representation of the structural model when price contracts have a maximum duration of 4 quarters; furthermore, with this choice of lag order, the Ljung-Box Q(12) statistic indicates that the residuals of the unconstrained VAR are serially uncorrelated.

As noted above, the structural model consists of the generalized price-setting framework combined with the markup gap and output gap equations taken from the unconstrained VAR. The vector of estimated structural parameters, θ, includes the distribution of contract durations (ω_i for $i = 1, ..., 4$), the sensitivity of new contracts to aggregate real marginal cost (γ), and the standard deviation of the disturbance to the optimal price-setting equation

[15] Recent examples of this approach include Rotemberg and Woodford (1997), Christiano et al. (2001), and Altig et al. (2004).

[16] For recent examples of full-information estimation, see Smets and Wouters (2003), Schorfheide (2003), and Onatski and Williams (2004).

(σ_ϵ). The distribution of contract durations is estimated subject to the constraint that these parameters are non-negative and sum to unity. Finally, rather than estimating the discount factor, we simply calibrate $\beta = 0.9925$, corresponding to an annualized steady-state real interest rate of about 3 percent.

For any particular vector of the structural parameters, θ, we confirm that the model has a unique linear rational expectations solution and then obtain its reduced-form VAR representation using the AIM algorithm of Anderson and Moore (1985). Using this reduced-form model, we generate "artificial" time series of length S for the endogenous variables, namely, the relative contract prices, the inflation gap, the marginal cost gap, and the output gap.[17] We then fit the latter three randomly-generated series with an unconstrained VAR model that is isomorphic to the one applied to the observed data. The vector of fitted VAR parameters is denoted by $\hat{\zeta}_S(\theta)$ because these VAR parameters depend on the particular values of the structural parameters as well as the restrictions of the structural model and the sample size S of the simulated data.

We then use a numerical optimization algorithm to determine the set of structural parameters that maximizes the fit between the simulation-based VAR parameters and those of the unconstrained VAR of the observed data. In particular, the estimated value of θ minimizes the following criterion function:

$$Q_{S,T}(\theta) \quad = \quad \left(\hat{\zeta}_T - \hat{\zeta}_S(\theta)\right)' \delta \left[\delta' \hat{\Sigma}_T(\hat{\zeta}_T) \delta\right]^{-1} \delta' \left(\hat{\zeta}_T - \hat{\zeta}_S(\theta)\right), \tag{12}$$

where δ is the matrix of zeros and ones that selects the elements of $\hat{\zeta}_T$ that correspond to the inflation equation of the unconstrained VAR.

Because this criterion function employs the optimal weighting matrix, the resulting estimator of θ is asymptotically efficient. In particular, under certain regularity conditions (including the assumption that the sample size ratio S/T converges to a constant q as

[17]To simulate the model, we employ a Gaussian random-number generator for the disturbances, and we use steady-state values as initial conditions for the endogenous variables; the first few years of simulated data are excluded from the sample used for indirect inference to ensure that the results are not influenced by these particular initial conditions.

$T \to \infty$), this estimator is consistent and has the following asymptotic normal distribution:

$$\sqrt{T}\,(\hat{\theta}_{S,T} - \theta_0) \xrightarrow{d} \mathrm{N}[0, (1 + q^{-1})(\mathcal{Z}'\ [\delta\Sigma_\zeta\delta\prime]^{-1}\ \mathcal{Z})^{-1}], \tag{13}$$

where θ_0 is the probability limit of $\hat{\theta}_{S,T}$; ζ_0 is the plim of $\hat{\zeta}_T$; $\Sigma_\zeta(\zeta_0)$ is the plim of $\hat{\Sigma}_T(\hat{\zeta}_T)$; $z(\theta_0)$ is the plim of $\hat{\zeta}_S(\theta_0)$ as $S \to \infty$; and $\mathcal{Z} = (\partial z(\theta_0)/\partial\theta')$.

5 The Estimated Price-Setting Framework

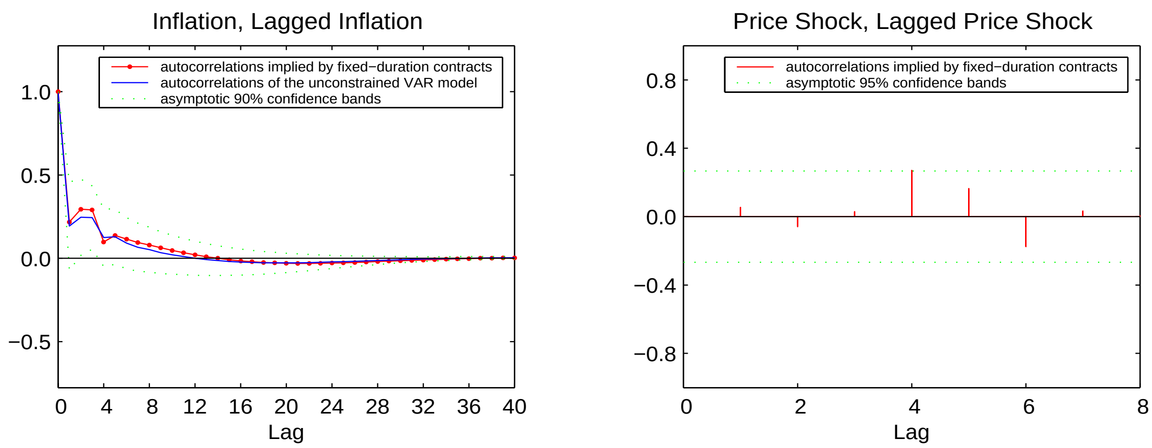
In this section, we present the results obtained from estimating the generalized price-setting framework using the methodology described above. We start by evaluating the extent to which the model is consistent with the data, and then examine the parameter estimates to determine the implied degree of nominal and real rigidities in price-setting behavior. Finally, we document the importance of accounting for the evolution of the Bundesbank's medium-term inflation objective.

5.1 Consistency with the Data

As discussed earlier, our estimation procedure is aimed at matching the reduced-form implications of the structural model to those of an unconstrained VAR. Thus, following McCallum (1999), a natural starting point for evaluating the goodness-of-fit of the structural model is to compare its implied autocorrelations with the sample autocorrelations of the observed time series. We also check the residuals of the reduced-form VAR implied by the structural model to check whether the autocorrelogram of these residuals is consistent with the maintained assumption of white-noise innovations.

According to both metrics, the generalized price-setting framework performs very well in fitting the characteristics of the German macroeconomic data. Complete results are given in the Appendix; here we simply illustrate the general pattern using one particular specification, namely, the fixed-duration contract model estimated using the benchmark markup gap as the measure of real marginal cost. As shown in the left panel of **Figure 3**, the

Figure 3: Inflation Dynamics and Correlations of Price Shocks



Notes: Solid lines with bold dots: Autocorrelation function of inflation implied by the estimated fixed-duration staggered-contracts specification. Solid lines: Autocorrelation functions implied by the trivariate VAR(3) model of the inflation gap, markup gap, and output gap. Solid bars: Autocorrelation function of price shocks implied by the estimated fixed-duration staggered-contracts specification. Dotted lines: Asymptotic confidence bands.

autocorrelogram of inflation implied by the structural model is virtually indistinguishable from that of the observed data and lies well within the asymptotic confidence bounds.[18] Furthermore, as depicted in the right panel, the contract price shocks exhibit negligible autocorrelation–a finding which is confirmed by portmanteau tests for serial correlation.

A more formal means of evaluating the structural model is to test whether the overidentifying restrictions of the model are consistent with the data. The degrees of freedom of the overidentification test depends on the number of free parameters in the structural model compared with the unconstrained VAR. When the structural model is estimated using one of the markup gap series as a proxy for real marginal cost, the model is matched to an trivariate VAR involving the markup gap, inflation gap, and output gap; in this case, the test of overidentifying restrictions has seven degrees of freedom. When the structural model is estimated using the output gap as the proxy for real marginal cost, then the correspond-

[18]See Coenen (2004) for a detailed discussion of the methodology used in computing the asymptotic confidence bands for the estimated autocorrelation functions.

Table 1: Testing the Overidentifying Restrictions

	Random-Duration Contracts	Fixed-Duration Contracts
Benchmark Markup Gap	0.39	0.41
Alternative Markup Gap	0.03**	0.03**
Output Gap (Linear Trend)	0.14	0.25
Output Gap (HP Trend)	0.06*	0.19

Note: This table indicates the probability that the overidentifying restrictions are consistent with each specification of the generalized price-setting framework for each of the four different proxies for real marginal cost. The test statistic has 7 degrees of freedom for each of the markup gap measures, and 3 degrees of freedom for each of the output gap measures. The inclusion of an asterisk indicates rejection at the 90% confidence level, while the inclusion of two asterisks denotes rejection at the 95% confidence level.

ing unconstrained model is a bivariate VAR involving the inflation gap and the output gap, and the overidentification test has three degrees of freedom.

As shown in **Table 1**, when the model is estimated using either the benchmark markup gap or the linearly-detrended output gap, the overidentifying restrictions are not rejected at the 95 percent confidence level for either the random-duration or fixed-duration versions of the model. Evidently, these results are not simply due to lack of statistical power: the overidentifying restrictions are rejected at a confidence level exceeding 95 percent when the uncorrected markup gap is used as the proxy for real marginal cost, and these restrictions are rejected at nearly the 95 percent confidence level for the random-duration model when the HP-detrended output gap is used as the proxy variable.

5.2 The Estimated Degree of Nominal and Real Rigidities

Our estimates for the nominal and real rigidity parameters of the generalized price-setting framework are reported in **Table 2**; the estimated standard deviation of the contract price shock may be found in Appendix Table A1.

The estimated distribution of price contract durations indicates a relatively moderate

Table 2: Estimates of Nominal and Real Rigidities

	Distribution of Contract Durations ω_1	ω_2	ω_3	ω_4	Mean Duration	Real Rigidity (γ)
A. Random-Duration Contracts						
Benchmark	0.55	0.17	0.06	0.22	1.95	0.0267
Markup Gap	(0.12)	(0.07)	(0.04)	(0.10)	(0.24)	0.0065
Alternative	0.54	0.20	0.08	0.18	1.90	0.0078
Markup Gap	(0.14)	(0.07)	(0.06)	(0.09)	(0.21)	(0.0028)
Output Gap	0.49	0.17	0.14	0.21	2.07	0.0064
(Linear Trend)	(0.11)	(0.08)	(0.07)	(0.09)	(0.23)	(0.0016)
Output Gap	0.44	0.16	0.14	0.27	2.25	0.0280
(HP Trend)	(0.11)	(0.08)	(0.07)	(0.10)	(0.26)	(0.0035)
B. Fixed-Duration Contracts						
Benchmark	0.33	0.21	0.12	0.34	2.46	0.0139
Markup Gap	(0.08)	(0.07)	(0.07)	(0.10)	(0.29)	(0.0022)
Alternative	0.34	0.24	0.14	0.27	2.35	0.0039
Markup Gap	(0.09)	(0.08)	(0.08)	(0.10)	(0.27)	(0.0013)
Output Gap	0.33	0.21	0.21	0.24	2.36	0.0032
(Linear Trend)	(0.09)	(0.07)	(0.09)	(0.09)	(0.26)	(0.0008)
Output Gap	0.33	0.23	0.19	0.25	2.37	0.0145
(HP Trend)	(0.08)	(0.07)	(0.09)	(0.09)	(0.26)	(0.0023)

Note: This table reports the distribution of contract durations and the real rigidity parameter for each specification of the generalized price-setting framework (that is, random or fixed durations) using each of the four different proxies for real marginal cost. Estimated standard errors are given in parentheses.

degree of nominal rigidity, broadly consistent with the evidence from surveys and micro price records regarding the frequency of price adjustment. In particular, for the specification with random-duration contracts, the mean duration is estimated to be about 2 quarters, regardless of the particular measure of real marginal cost. The mean duration is estimated to be a bit longer (namely, about 2-1/2 quarters) under the assumption of fixed-duration contracts, but the difference from the random-duration estimate is not statistically significant.

In evaluating the degree of real rigidity, the model with no firm-specific inputs and a constant elasticity of demand provides a natural benchmark, because in this case $\gamma = \gamma_d = \gamma_m c = 1$; that is, a one percent increase in real marginal cost causes a one percent rise in the level of new price contracts. In contrast, Table 2 indicates that new price contracts exhibit much lower sensitivity to real marginal cost, with γ only about one-twentieth as large as in the benchmark case. Furthermore, equation (7) shows that both firm-specific inputs and strong curvature of the demand function are needed to generate the estimated degree of real rigidity.

5.3 Ignoring Time-Variation in the Inflation Objective

The generalized price-setting framework is oriented towards explaining short-run inflation dynamics in response to shifts in real marginal cost, treating the central bank's objective as fixed and known. For this reason, our discussion thus far has focused on the estimation results obtained using the "inflation gap," that is, the deviation of actual inflation from the Bundesbank's medium-term inflation objective. Now we briefly turn to the implications of ignoring time-variation in the inflation objective–an approach which characterizes much of the empirical NKPC literature.

Table 3 reports the contract distribution and real rigidity parameters obtained when the price-setting framework is estimated using the level of inflation rather than the inflation gap. Evidently, the mean duration of price contracts is noticeably longer–close to three quarters for the random-duration model, and a bit longer for the fixed-duration model. Furthermore, the restrictions implied by a truncated Calvo distribution can be clearly rejected in either case, because the estimated distribution involves a relatively large number of one and four quarter contract durations with relatively few 2-3 quarter contracts.

Nevertheless, our model diagnostics indicate that this specification falls short of a satisfactory match with the observed data. In particular, while the left panel of Figure 4 shows that the inflation autocorrelations implied by the model match those of the data, the right panel reveals that the autocorrelogram of the price contract shocks looks unreasonable in this case, with a highly significant level of serial correlation of order 4. Based on

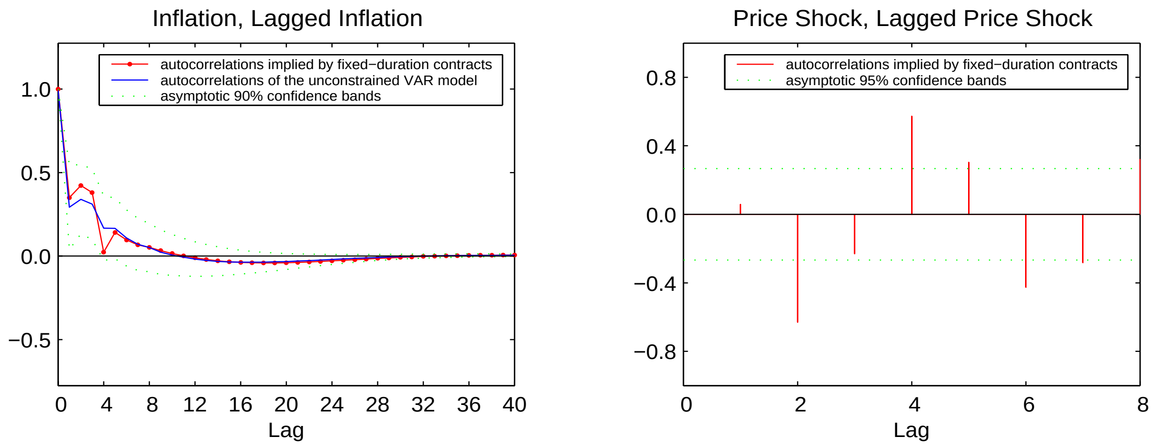
Table 3: Implications of Ignoring Time-Variation in the Inflation Objective

	Distribution of Contract Durations ω_1	ω_2	ω_3	ω_4	Mean Duration	Real Rigidity (γ)
A. Random-Duration Contracts						
Benchmark	0.42	0.02	0.06	0.51	2.66	0.0413
Markup Gap	(0.10)	(0.03)	(0.05)	(0.13)	(0.40)	(0.0022)
Output Gap	0.38	0	0.11	0.51	2.76	0.0113
(Linear Trend)	(0.06)	—	(0.06)	(0.147)	(0.45)	0.0027
B. Fixed-Duration Contracts						
Benchmark	0.18	0.05	0.16	0.61	3.19	0.0082
Markup Gap	(0.05)	(0.05)	(0.08)	(0.16)	(0.49)	(0.0019)
Output Gap	0.17	0	0.08	0.75	3.40	0.0319
(Linear Trend)	(0.05)	—	(0.06)	(0.17)	(0.56)	0.007

Note: This table reports the parameter estimates for each specification of the generalized price-setting framework (that is, random or fixed contract duration) when the model is estimated using the level of inflation and the specified proxy for real marginal cost over the period 1974-1998. These results correspond to the assumption of a constant medium-term inflation objective over the entire sample period. Estimated standard errors are given in parentheses.

this evidence, we conclude that accounting for time-variation in the Bundesbank's inflation objective during the period 1975-84 is important in obtaining accurate estimation results.

Figure 4: Model Diagnostics from Ignoring Time-Variation in the Inflation Objective



Notes: Solid lines with bold dots: Autocorrelation function of inflation implied by the estimated fixed-duration staggered-contracts specification. Solid lines: Autocorrelation functions implied by the trivariate VAR(3) model of inflation, the markup gap, and the output gap. Solid bars: Autocorrelation function of price shocks implied by the estimated fixed-duration staggered-contracts specification. Dotted lines: Asymptotic confidence bands.

6 Conclusion

In this paper, we have formulated a generalized price-setting framework that incorporates staggered contracts of multiple durations and that directly identifies the influences of nominal vs. real rigidities. Using German macroeconomic data over the period 1975Q1 through 1998Q4 to estimate this framework, we find that the data is well-characterized by a truncated Calvo-style distribution with an average duration of about two quarters and by a relatively high degree of real rigidity. Finally, our results indicate that backward-looking behavior is not needed to explain the aggregate data, at least in an environment with a stable monetary policy regime and a transparent and credible inflation objective.

This paper has proceeded under the assumption that all firms face the same output elasticity of marginal cost. In subsequent work, it will be interesting to explore whether this parameter varies systematically across groups of firms with different contract durations; that is, whether the aggregate data imply a cross-sectional relationship between nominal and real rigidities. Furthermore, the approach used here can easily be applied to other economies, especially for sample periods over which the inflation objective has been reasonably stable or has evolved gradually in a transparent way. Finally, our approach can be extended to consider the joint determination of aggregate wages and prices, in a framework that allows for multiple-period durations of both types of contracts.

Appendix A

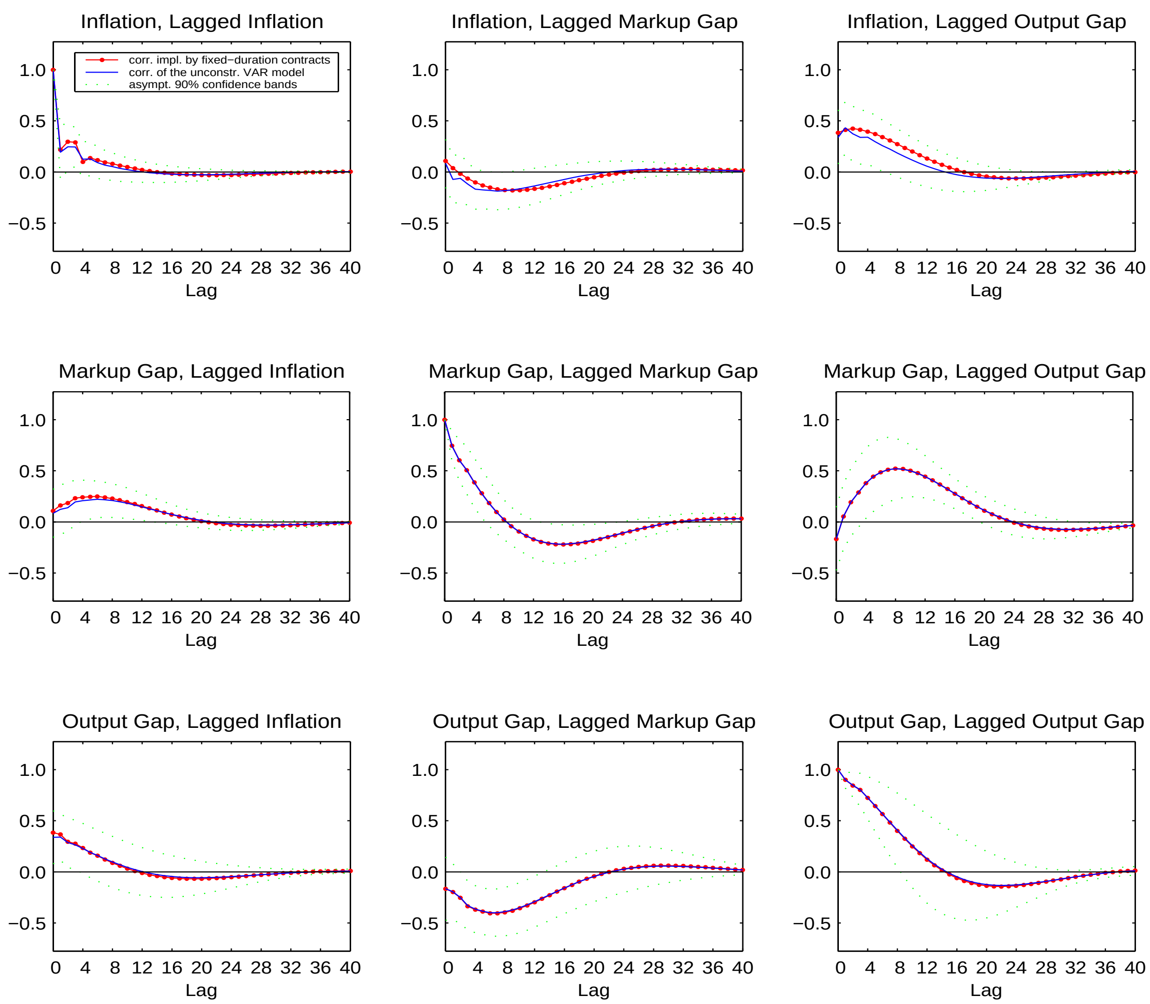
This appendix provides further details regarding the estimation results for the generalized price-setting framework.

Appendix Table A1: The Standard Deviation of the Contract Price Shock

	Random-Duration Contracts	Fixed-Duration Contracts
A. Estimated using Inflation Gap		
Benchmark	0.0035	0.0030
Markup Gap	(0.0006)	(0.0004)
Alternative	0.0032	0.0027
Markup Gap	(0.0007)	(0.0005)
Output Gap	0.0038	0.0029
(Linear Trend)	(0.0005)	(0.004)
Output Gap)	0.0043	0.0029
(HP Trend	(0.0004)	(0.0004)
B. Estimated using Level of Inflation		
Benchmark	0.0053	0.0049
Markup Gap	(0.0003)	(0.0004)
Output Gap	0.0055	0.0051
(Linear Trend)	(0.0003)	(0.0003)

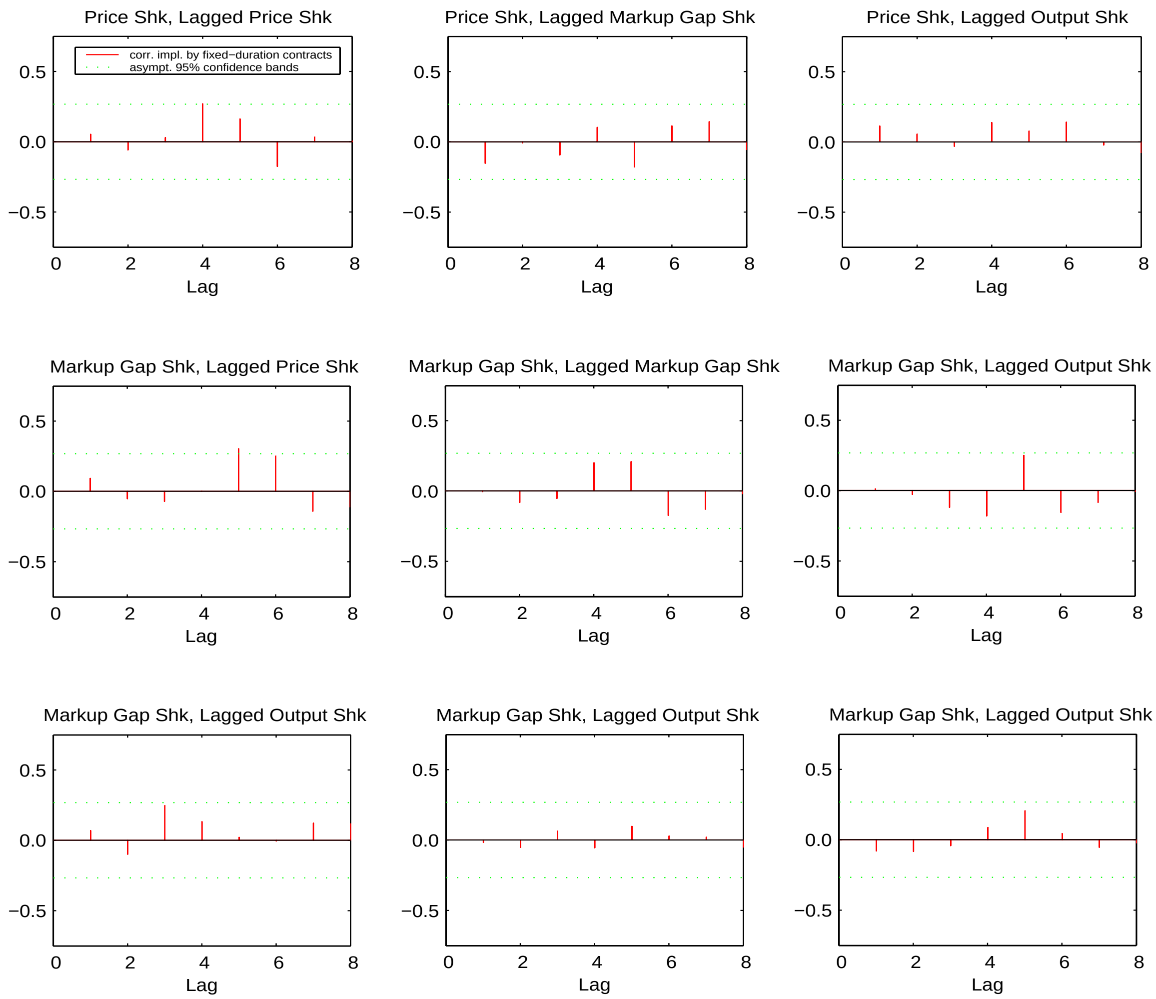
Note: This table indicates the estimated standard deviation of the contract price shock for each specification of the generalized price-setting framework for each of the different proxies for real marginal cost. Panel A shows results when the model is estimated using the inflation gap (which incorporates the Bundesbank's medium-term inflation objective), while Panel B provides results when the model is estimated using inflation in levels (corresponding to the assumption of a time-invariant inflation objective). The standard error of each estimate is enclosed in parentheses.

Appendix Figure A1: Comparison of Autocorrelations



Note: This figure depicts results for the fixed-duration contract model estimated using the benchmark markup gap. Solid lines with bold dots denote the autocorrelation functions implied by the estimated model. Plain solid lines denote the autocorrelation functions implied by the unconstrained trivariate VAR(3) model, and dotted lines denote the asymptotic 90 percent confidence bands for each autocorrelation function.

Appendix Table A2: Autocorrelations of the Implied VAR Residuals



Note: This figure depicts results for the fixed-duration contract model estimated using the benchmark markup gap. Solid bars denote the autocorrelation functions of the residuals of the reduced-form VAR implied by the model, and the dotted lines denote the asymptotic 95 percent confidence bands.

Appendix B

This appendix considers a simplified version of the generalized price-setting framework for purposes of comparison with some earlier literature. In particular, in the case of fixed-duration contracts, the average new contract price $\bar{x}_t$ is given by

$$\bar{x}_t \quad = \quad \mathrm{E}_t \left[\sum_{j=1}^{J-1} \bar{\Phi}_j \, \pi_{t+j} + \gamma \sum_{j=0}^{J-1} \bar{\phi}_j \, mc_{t+j} \right]$$

with $\bar{\phi}_j = \beta^j \sum_{i=j+1}^{J} \left(\omega_i / \sum_{k=0}^{i-1} \beta^k \right)$ and $\bar{\Phi}_j = \sum_{k=j}^{J-1} \bar{\phi}_k$.

As in Taylor (1993) and Guerrieri (2002), we consider the approximation obtained by assuming negligible variation across new price contracts of different durations; that is, $x_{jt} \approx \bar{x}_t$. In this case, log-linearization around the zero steady-state inflation rate yields the following expression for the aggregate price identity:

$$\sum_{j=0}^{J-1} \bar{\psi}_j \, \bar{x}_{t-j} \quad = \quad \sum_{j=0}^{J-2} \bar{\Psi}_{j+1} \, \pi_{t-j}.$$

where $\bar{\psi}_j = \sum_{k=j+1}^{J} (\omega_k / k)$ and $\bar{\Phi}_j = \sum_{k=i}^{J-1} \bar{\phi}_k$.

Thus, using this approximation, the weights in the aggregate identity are identical to those in the price-setting equation, just as in the case of random-duration contracts. Furthermore, in the special case of no discounting ($\beta = 1$), the simplified fixed-duration contract specification is observationally equivalent to the random-duration specification.

Evidently, the simplified fixed-duration contract specification implies somewhat longer average duration compared with the generalized fixed-duration or random-duration specifications considered above.

Appendix Table A2: The Simplified Fixed-Duration Specification

	ω_1	ω_2	ω_3	ω_4	γ	σ_ϵ	Mean Duration	p-value
A. Inflation Gap								
Benchmark Markup Gap	0.280 (0.057)	0.175 (0.063)	0.091 (0.104)	0.454 (0.314)	0.0263 (0.0037)	0.0035 (0.0008)	2.718 (0.761)	0.393 [7]
Alternative Markup Gap	0.202 (0.079)	0.125 (0.029)	0.132 (0.019)	0.541 (0.089)	0.0058 (0.0019)	0.0042 (0.0006)	3.014 (0.331)	0.008 [7]
Output Gap (Linear Trend)	0.235 (0.049)	0.162 (0.276)	0.204 (0.080)	0.399 (0.503)	0.0062 (0.0114)	0.0038 (0.0018)	2.766 (0.342)	0.140 [3]
Output Gap (HP Trend)	0.194 (0.041)	0.140 (0.079)	0.184 (0.090)	0.482 (0.148)	0.0276 (0.0050)	0.0043 (0.0001)	2.955 (0.437)	0.058 [3]
B. Inflation in Levels								
Benchmark Markup Gap	0.157 (0.036)	0.014 (0.024)	0.062 (0.050)	0.767 (0.186)	0.0410 (0.0042)	0.0053 (0.0005)	3.439 (0.654)	0.461 [7]
Output Gap (Linear Trend)	0.136 (0.050)	0	0.123 (0.049)	0.741 (0.247)	0.0112 (0.0023)	0.0055 (0.0003)	3.469 (0.841)	0.294 [4]

Notes: For each measure of real marginal cost, this table reports the estimated parameters of the price-setting framework with fixed-duration contracts, using the simplifying assumption of negligible variation across new price contracts of different durations. Estimated standard errors are given in parentheses. The column labelled "p-value" indicates the probability that the overidentifying restrictions can be rejected, where the number of overidentifying restrictions is enclosed in brackets.

References

Altig, D., Christiano, L., Eichenbaum, M., Linde, J. 2004. An estimated model of the U.S. business cycle. Manuscript, Northwestern University.

Anderson, G., Moore, G. 1985. A linear algebraic procedure for solving linear perfect foresight models. *Economics Letters* 17, pp.247-252.

Angeloni, I., Ehrmann, M. 2004. Euro area inflation differentials. Manuscript, European Central Bank.

Ascari, G., 2003, Staggered prices and trend inflation: Some nuisances, Bank of Finland Discussion Papers 27/2003.

Aucremanne, L., Dhyne, E. 2004. How frequently do price changes? First results on the micro data underlying the Belgian CPI. European Central Bank Working Paper No. 332.

Ball, L., Romer, D. 1990. Real Rigidities and the Non-Neutrality of Money. *Review of Economic Studies*, Vol. 57, pp. 183-203.

Baudry, L., LeBihan, H., Sevestre, P., Tarrieu, S. 2004. Price Rigidity: Evidence from French CPI Microdata. Forthcoming, European Central Bank Working Paper.

Bernanke, B., Woodford, M. 1997. Inflation Forecasts and Monetary Policy. *Journal of Money, Credit, and Banking* 29(4), pp. 65-84.

Bils, M., Klenow, P. 2004. Some Evidence on the Importance of Sticky Prices. Forthcoming, *Journal of Political Economy.*

Apel, M., Friberg, R., Hallsten, K. 2001. Micro Foundations of Price Adjustment: Survey Evidence from Swedish Firms. Sveriges Riksbank Working Paper No. 128.

Blinder, A., Canetti, E., Lebow, D., Rudd, J. 1998. *Asking about Prices: A New Approach to Understand Price Stickiness.* New York: Russell Sage Foundation.

Calvo, G. 1983. Staggered Prices in a Utility-Maximizing Framework. *Journal of Monetary Economics* 12, pp.383-398.

Caplin, A., Leahy, J. 1997. Aggregation and Optimization with State-Dependent Pricing. *Econometrica* 65(3), pp.601-625.

Chari, V., Kehoe, P., McGrattan, E. 2000. Sticky Price Models of the Business Cycle: Can the Contract Multiplier Solve the Persistence Problem? *Econometrica*, pp.1151-80.

Chevalier, J., Kashyap, A., Rossi, P. 2003. Why Don't Prices Rise during Periods of Peak Demand? Evidence from Scanner Data. *American Economic Review* 93(1), pp.15-37.

Christiano, L., Eichenbaum, M., Evans, C. 2001. Nominal Rigidities and the Dynamic Effects of a Shock to Monetary Policy. Forthcoming, *Journal of Political Economy.*

Clarida, R., Gali, J., Gertler, M. 1999. The Science of Monetary Policy: A New Keynesian Perspective." *Journal of Economic Literature*, December 1999, 37(4), pp. 1661-1707.

Coenen, G. 2004. Asymptotic Confidence Bands for the Estimated Autocovariance and Autocorrelation Functions of Vector Autoregressive Models. Forthcoming, *Empirical Economics*.

Coenen, G., Wieland, V. 2004. A Small Estimated Euro Area Model with Rational Expectations and Nominal Rigidities. European Central Bank Working Paper No. 30. Forthcoming, *European Economic Review*.

Dias, M., Dias, D., Neves, P. 2004. Stylised feature of price setting behaviour in Portugal: 1992-2001. European Central Bank Working Paper No. 331.

Dotsey, M., King, R., Wolman, A. 1999. State-Dependent Pricing and the General Equilibrium Dynamics of Money and Output. *Quarterly Journal of Economics* 114(2), pp.655-690.

Eichenbaum, M., Fisher, J. 2004. Evaluating the Calvo model of sticky prices. Manuscript, Northwestern University.

Erceg, C., Levin, A. 2003. Imperfect Credibility and Inflation Persistence. *Journal of Monetary Economics* 50, pp.915-944.

Fabiani, S., Gattulli, A., Sabbatini, R. 2004. The pricing behaviour of Italian firms: new survey evidence on price stickiness. European Central Bank Working Paper No. 331.

Fuhrer, J., Moore, G. 1995. Inflation Persistence. *Quarterly Journal of Economics* 110, pp.127-159.

Galì, J., Gertler, M. 1999. Inflation Dynamics: A Structural Econometric Analysis. *Journal of Monetary Economics* 44, pp.195-222.

Galì, J., Gertler, M., Lpez-Salido, J.D. 2001, European Inflation Dynamics. *European Economic Review* 43, pp.1237-1270.

Gerberding, C., Seitz, F., Worms, A. 2004. How the Bundesbank really conducted monetary policy: An analysis based on real-time data, Manuscript.

Golosov, M., Lucas, R. 2003. Menu Costs and Phillips Curves. NBER Working Paper.

Goodfriend, M., King, R. 1997. The New Neoclassic Synthesis and the Role of Monetary Policy. In Bernanke, B., Rotemberg, J., eds., *NBER Macroeconomics Annual 1997*. Cambridge, MA: MIT Press.

Guerrieri, L. 2002. Persistent issues in inflation persistence. Manuscript, Board of Governors of the Federal Reserve System.

Hall, S., Walsh, M., Yates, T. 2000. Are UK Companies Prices Sticky? *Oxford Economic Papers* 52(3), pp. 425-446.

Kimball, M. 1995. The Quantitative Analytics of the Basic Neomonetarist Model. *Journal of Money, Credit and Banking* 27, pp.1241-1277.

Klenow, P., Kryvtsov, O. 2004. State-dependent or time-dependent pricing: Does it matter for recent U.S. inflation? Manuscript, Stanford University.

Levin, A., Piger, J. 2004. Is Inflation Persistence Intrinsic in Industrial Economies? European Central Bank Working Paper.

Lindé, J. 2001. Estimating New-Keynesian Phillips Curves: A full information maximum likelihood approach. Sveriges Riksbank Working Paper Series No. 129.

Mash, R. 2003. Optimising Microfoundations for Inflation Persistence. Oxford University Department of Economics Discussion Paper No. 174.

McCallum, B. 1999. Analysis of the monetary transmission mechanism: methodological issues. Manuscript, Carnegie-Mellon University.

Neiss, K., Nelson, E. 2002. Inflation dynamics, marginal cost, and the output gap: Evidence from three countries, Manuscript.

Onatski, A., Williams, N. 2004. Empirical and Policy Performance of a Forward-Looking Monetary Model. Manuscript, Princeton University.

Rotemberg, J., Woodford, M. 1997. An Optimization-Based Econometric Framework for the Evaluation of Monetary Policy. In Bernanke, B., Rotemberg, J., eds., *NBER Macroeconomics Annual 1997*. Cambridge, MA: MIT Press.

Sbordone, A. 2002. Prices and Unit Labor Costs: A New Test of Price Stickiness. *Journal of Monetary Economics* 49, pp.265-292.

Schorfheide, F. 2003. Manuscript, University of Pennsylvania.

Schmid, P. 1999. Monetary Targeting in Practice: The German Experience. Center for Financial Studies Working Paper No. 1999-03.

Smets, F., Wouters, R. 2003. An estimated dynamic stochastic general equilibrium model of the Euro area. *Journal of the European Economic Association* 1, pp.1123-1175.

Sondergaard, L. 2003. Inflation dynamics in the traded sectors of Frances, Italy and Spain. Manuscript, European Central Bank.

Sveen, T., Weinke, L. 2003. Inflation and output dynamics with staggered price setting and capital accumulation. Manuscript.

Taylor, J. 1980. Aggregate dynamics and staggered contracts. *Journal of Political Economy* 88, pp.1-22.

Taylor, J. 1993. *Macroeconomic policy in the world economy: From econometric design to practical operation.* New York: W.W. Norton.

Taylor J. 1999. Staggered price and wage setting in macroeconomics. In *Handbook of Macroeconomics, volume 1*, Taylor, J., Woodford, M., eds. Elsevier Press.

Woodford, M. 2003. *Interest and Prices: Foundations of a Theory of Monetary Policy.* Princeton, NJ: Princeton University Press.

Woodford, M., 2004, Inflation and output dynamics with firm-specific investment. Manuscript, Princeton University.

Yun, T. 1996. Nominal Price Rigidity, Money Supply Endogeneity, and Business Cycles. *Journal of Monetary Economics* 37(2), pp.345-370.