A Small Estimated Euro-Area Model with Rational Expectations and Nominal Rigidities

Günter Coenen
European Central Bank

Volker Wieland *,†
Board of Governors of the Federal Reserve System and European Central Bank

January 2000

JEL Classification System: E31, E52, E58, E61

Keywords: European Monetary Union, Euro area, macroeconomic modelling, rational expectations, nominal rigidities, overlapping wage contracts, inflation persistence, monetary policy rules

---

* Correspondence: Coenen: Directorate General Research, European Central Bank, Frankfurt am Main, Germany, Tel.: (0)69 1344-7887, e-mail: gunter.coenen@ecb.int; Wieland: Federal Reserve Board, Washington, DC, 20009, U.S.A., Tel.: (202) 736-5620, e-mail: vwieland@frb.gov, Homepage: http://www.ifk-cfs.de/pages/makro/guests/wieland/index_e.htm.

† The opinions expressed are those of the authors and do not necessarily reflect views of the European Central Bank or the Board of Governors of the Federal Reserve System. Volker Wieland served as a consultant at the European Central Bank while preparing this paper. We are grateful for research assistance by Anna-Maria Agresti from the European Central Bank. Helpful comments by Jeffrey Fuhrer, Thomas Laubach, Athanasios Orphanides, Richard Porter, John Taylor and by seminar participants at the European Central Bank are greatly appreciated. Any remaining errors are of course the sole responsibility of the authors.

Abstract

With the formation of European Monetary Union (EMU) in 1999, the eleven countries that adopted the Euro began to conduct a single monetary policy oriented towards union-wide objectives. The objective of this paper is to construct a small model of the Euro area, which may serve as a laboratory for evaluating the performance of alternative monetary policy strategies for the Euro area in the vein of recent studies for the United States.

In estimating this model, we start with the relationship between output and inflation and investigate the nominal wage contracting model due to Taylor (1980) as well as three different versions of the relative real wage contracting model first proposed by Buiter and Jewitt (1981) and investigated empirically with U.S. data by Fuhrer and Moore (1995a). Contrary to Fuhrer and Moore, who reject the nominal contracting model and prefer a version of the relative contracting model which induces a higher degree of inflation persistence, we find that both types of contracting models fit the data for the Euro area. The best fitting specification, however, is a version of the relative contracting model, which is theoretically more plausible than the simplified version preferred by Fuhrer and Moore. These findings are relevant to the continuing debate on sticky- inflation versus sticky-price models (for example Roberts (1997), Sbordone (1999), Gali and Gertler (1999) and Taylor (1999)) and may have important implications for the short-run inflation-output tradeoff faced by the Eurosystem.

A drawback of the Euro area estimation is that the data are averages of the member economies, which experienced different monetary policy regimes prior to the formation of EMU. While Germany enjoyed stable inflation with fairly predictable monetary policy, countries such as France and Italy experienced a long-drawn out and probably imperfectly anticipated disinflation. To investigate the validity of our results, we also estimate the contracting models for France, Germany and Italy separately. We find that the relative contracting model dominates in countries which transitioned out of a high inflation regime such as France and Italy, while the nominal contracting model fits German data better. Thus, an optimist may conclude that the independent European Central Bank will face a similar environment in the future as the Bundesbank did in Germany and pick the nominal contracting specification, while a pessimist, who suspects that stabilizing Euro area inflation will require higher output losses, may want to pick the relative contracting specification. A robust monetary policy strategy, however, should perform reasonably well in both cases. We close the model by imposing a term-structure relationship and estimating an aggregate demand relationship. We then evaluate the performance of Taylor's rule as an example.

1 Introduction

With the formation of European Monetary Union (EMU) in 1999, the eleven countries that adopted the Euro began to conduct a single monetary policy oriented towards union-wide objectives.[1] As prescribed by the Maastricht Treaty the primary goal of this policy is to maintain price stability within the Euro area. The operational definition of this goal announced by the European Central Bank (ECB) is to aim for year-on-year increases in the Euro-area inflation rate below 2 percent.[2] To evaluate alternative strategies for achieving such a Euro-area-wide objective, it is essential to build empirical models that can be used to assess the area-wide impact of policy on key macroeconomic variables such as output and inflation. Thus, the objective of this paper is to construct a small model of the Euro area, which may serve as a laboratory for evaluating the performance of alternative monetary policy strategies in the vein of recent studies for the United States.[3]

One possible approach to building a model of the Euro area is to start from the bottom up by constructing separate models of the individual member economies and then link these models together in a multi-country model. The main alternative is to first aggregate the relevant macroeconomic time series across member economies, and then estimate a model for the Euro area as a whole. In this paper, we pursue the latter approach, the reason being that the objectives as well as the instruments of Eurosystem monetary policy are defined on the Euro-area level. Of course, a problem with this approach is that the data

[1] Austria, Belgium, Finland, France, Germany, Ireland, Italy, Luxembourg, the Netherlands, Portugal and Spain. Denmark, Sweden, Greece and the U.K., who have not adopted the Euro, are not part of the Euro area. Their central banks, however, are members of the European System of Central Banks but not of the Eurosystem, which comprises the central banks of the countries that adopted the Euro as well as the European Central Bank.

[2] As measured by the Harmonized Index of Consumer Prices (HICP). It was further clarified that this definition excludes decreases, thus implying a range from 0 to 2 %. A detailed discussion of these and other issues regarding the ECB's strategy can be found in Angeloni, Gaspar and Tristani (1999).

[3] The recent literature on evaluating monetary policy rules for the U.S. economy has used a variety of macro models, including small-scale backward-looking models such as Rudebusch and Svensson (1999), large-scale backward-looking models such as Fair and Howrey (1996), small-scale models with forward-looking rational expectations and nominal rigidities (cf. Fuhrer and Moore (1995a), (1995b), Fuhrer (1997a), Orphanides, Small, Wieland and Wilcox (1998)), large-scale models of this type such as Taylor (1993a) and the Federal Reserve Board's FRB/US model (cf. Brayton et al. (1997)) as well as small models with optimizing agents such as Rotemberg and Woodford (1997) and McCallum and Nelson (1999). Recent comparative studies of interest include Bryant, Hooper and Mann (1993) and Taylor (1999).

used in aggregation stems from periods prior to EMU, when the different member economies experienced different monetary regimes and policies. For this reason, we also estimate every model specification separately for the three largest member economies, France, Germany and Italy, which together comprise over 70% of economic activity in the Euro area. By comparing the estimates obtained with French, German and Italian data to the Euro area estimates, we can assess to what extent the choice of model specification for the Euro area is influenced by the aggregation itself. Furthermore, by comparing France and Italy, which experienced a convergence process prior to EMU, with Germany, which enjoyed stable inflation and interest rates, we can see whether the choice of specification is influenced by differences in the monetary regime prior to EMU.

In building our small-scale Euro-area model we start with the relationship between inflation and output. In this respect we make three modelling assumptions, which are central to the key tradeoff that central banks are faced with, that is, the tradeoff between inflation and output variability. First, monetary policy has short-run real effects due to the existence of overlapping wage contracts. Second, expectations in labor, financial and goods markets are formed in a forward-looking, rational manner. Third, in estimating Euro-area inflation dynamics with pre-EMU data, we use the deviations of inflation from the downward trend, which was generated by the convergence of inflation in Italy, France, Spain and other member countries to German levels, rather than the inflation rate itself. Following this modelling approach, we obtain several new findings, which are relevant to the recent empirical literature on nominal rigidities, and may have important implications for the policy tradeoffs faced by the ECB.

As to overlapping wage contracts, we explore the empirical fit of the nominal wage contracting model due to Taylor (1980) as well as three different versions of the relative real wage contracting model first proposed by Buiter and Jewitt (1981) and investigated empirically by Fuhrer and Moore (1995a). The nominal contracting model belongs to the class of New-Keynesian sticky-price models which – as shown in the recent theoretical literature

– are consistent with intertemporal optimization by imperfectly competitive firms.[4] However, Fuhrer and Moore (1995a) have argued that the nominal contracting model cannot explain the degree of inflation persistence observed in U.S. data, while the relative real wage contracting model induces sufficient inflation stickiness. Of course, this difference has important policy implications. In the nominal contracting model, the central bank can commit to a policy, which achieves disinflation without recession, but not in the relative contracting model. Comparing these alternative specifications, we find that the relative wage contracting model fits Euro area inflation dynamics better than the nominal contracting model. Among the three different versions of the relative real wage contracting model, it is not the simplified specification preferred by Fuhrer and Moore, but a theoretically more plausible specification, which obtains the best fit. We also note that the nominal contracting model cannot be rejected by the data. Comparing the estimates for the individual countries, we find that the same relative wage contracting model fits Italian and French data quite well, but not the German data, which exhibits a substantially lower degree of inflation persistence. Only the nominal contracting model seems to have a shot at explaining inflation dynamics in Germany.

Fuhrer and Moore's empirical findings have generated a continuing debate on the sources of inflation stickiness. Roberts (1997) showed that a sticky-inflation model with rational expectations is observationally equivalent to a sticky-price model with expectations that are imperfectly rational and, using data on survey expectations in the U.S., found evidence of backward-looking behavior. More recently, Sbordone (1998) and Gali and Gertler (1999) have argued that the New-Keynesian sticky-price model is capable of explaining U.S. inflation dynamics, if one uses a measure of marginal cost rather than the output gap as the determinant of inflation.[5] Finally, Taylor (1999) has pointed out that expectations of inflation influence the pricing power of firms, and argued that inflation is more persistent

[4]See Goodfriend and King (1997) for a comprehensive survey.

[5]As the authors show, a model with price-stickiness is sufficient in this case, because marginal costs themselves exhibit persistence. An open question, which needs to be settled in order to construct a complete macro model, concerns the source of the observed persistence in marginal costs.

in a high inflation regime than in a low inflation regime with credible monetary policy.

Our comparative analysis contributes to this debate in two ways. The assumption of rational expectations implies that market expectations take into account the decision rule of the policymaker. This case serves as a useful benchmark for policy evaluation, because the alternative assumption of backward-looking expectations would imply that the central bank can exploit systematic expectational errors by market participants. However, as noted by Roberts (1997), a large degree of inflation persistence as in the Italian, French and Euro area data, may be evidence of adaptive expectations rather than structural rigidities. Such backward-looking behavior would seem plausible in those countries, since the downward trend in inflation may at best have been imperfectly anticipated by market participants. Thus, as far as the future of the EMU is concerned, the estimation based on historical Euro-area data may overstate the case for the relative real wage contracting model. This conclusion is supported by the better fit of the nominal contracting model in Germany, where inflation was rather stable and monetary policy fairly predictable. The better fit of this model with German data also provides empirical support to the thesis that the degree of inflation persistence is lower in a stable monetary policy regime with low average inflation as suggested by Taylor (1999).

The downward trend in Euro area inflation, which arose from the convergence process in countries such as Italy, Spain and France, should not be expected to persist nor to be reversed in the future, if the ECB achieves its policy objective. The short-run variations around this trend, however, to the extent that they were due to structural rigidities, may still help predicting the inflation dynamics after the formation of EMU. In principle, a complete model of the European inflation process prior to EMU would need to account for both, the short-run variations as well as the long-run convergence. However, describing the latter process appropriately, would require taking into account the varying degree of credibility of exchange rate pegs, the possibility of crises and realignments and learning by market participants about the long-run inflation objectives of European policymakers. Such an analysis would be beyond the objective of this paper. Instead, we take a shortcut and simply detrend the Euro-area, French and Italian inflation rates. We then use the resulting inflation gap series together with series for the output gap to estimate the structural overlapping contracts models. This assumption would have been correct, if the disinflation had been fully anticipated, but if not, it could lead us to underestimate the degree of inflation persistence and understate the case for the relative real wage specification. Again, the estimation with German data provides a useful benchmark for comparison, because it is the only case for which the inflation series exhibits no strong trend. In addition, we conduct a sensitivity study to assess how our Euro-area estimates would change if market participants had been consistently surprised by the downward trend.

In terms of evaluating alternative monetary policy strategies for the Euro area, an analyst who is pessimistic about the output losses associated with stabilizing inflation might prefer to use the best fitting version of the relative wage contracting model. An optimist might prefer to use the nominal wage contracting model with the coefficients estimated with German data. A robust monetary policy strategy however should perform reasonably well under both specifications. We provide an illustrative example for the case of Taylor's rule.

The paper proceeds as follows. Section 2 reviews the overlapping contracts specifications. The data is discussed in section 3. Section 4 summarizes inflation and output dynamics in form of unconstrained VAR models, while the structural estimates of the contracting specifications by means of simulation-based indirect inference methods are reported in section 5. In section 6 we close the model with an aggregate demand equation, a term structure equation and a policy rule. Impulse responses and disinflation scenarios under alternative specifications are compared in section 7. Section 8 concludes and the appendix provides the details of the indirect estimation methodology.

2 Modelling Inflation Dynamics with Overlapping Contracts

We estimate four different specifications of overlapping wage contracts, the nominal wage contracting model of Taylor (1980) and three variants of the relative real wage contracting model estimated by Fuhrer and Moore (1995a) (FM in the following) for the United

States. While these models are motivated by the existence of long-term wage contracts, the implications for price and wage dynamics are essentially the same if prices are related to wages by a fixed markup. Thus, we follow FM in using price data instead of wage data in estimation, and from here on we use the terms "contract price" and "contract wage" interchangeably.[6]

A common feature of the four specifications is that the log aggregate price index in the current quarter, p_t, is a weighted average of the log contract wages, x_{t-i} $(i = 0, 1, \ldots)$, which were negotiated in the current and the preceding quarters and are still in effect. The sticky price index can be observed directly, while the flexible contract wage is an unobserved variable. As a benchmark we consider the case of a one-year weighted average:

$$p_t = f_0\, x_t + f_1\, x_{t-1} + f_2\, x_{t-2} + f_3\, x_{t-3}. \tag{1}$$

The weights f_i $(i = 0, 1, 2, 3)$ on contract wages from different periods are assumed to be time-invariant, $f_i \geq 0$ and $\sum_i f_i = 1$. As shown in Taylor (1980), these weights would be equal to .25, if 25 percent of all workers sign contracts each quarter and if each contract lasts one year. Taylor (1993a) provides an interpretation for the more general case with unequal weights in terms of the distribution of workers by lengths of contracts. He shows that the weights f_i are time-invariant, if the distribution of workers by contract length is time-invariant and if the variation of average contract wages over contracts of different length is negligible.[7] Restricting the number of lags in (1) to three is consistent with a maximum contract length of four quarters.[8] Rather than estimating each of the weights f_i separately, we follow FM and assume that the weights are a downward-sloping linear function of contract length, s.t. $f_i = .25 + (1.5 - i)\, s$, where $s \in (\,0, 1/6\,]$. This distribution depends on a single parameter, the slope s.

The determination of the nominal contract wage x_t for the different specifications is

[6] For recent studies considering both, wage and price stickiness, see Taylor (1993a), Erceg, Henderson and Levin (1999) and Amato and Laubach (1999).

[7] For the derivation see Taylor (1993a), pp. 35-38.

[8] Fuhrer and Moore (1995a) found this lag length sufficient to explain the degree of persistence in U.S. inflation data. Similarly, Taylor (1993a) estimated the nominal contracting model for all G-7 countries with such a lag length.

best explained starting with Taylor's nominal wage contracting model (the NW model in the following). In this case, x_t is negotiated with reference to the price level that is expected to prevail over the life of the contract, as well as the expected degree of excess demand over the life of the contract, which is measured in terms of the deviations of output from its potential, q_t:

$$x_t^{NW} = \mathrm{E}_t \left[\sum_{i=0}^{3} f_i \, p_{t+i} + \gamma \sum_{i=0}^{3} f_i \, q_{t+i} \right] + \sigma_{\epsilon_x} \, \epsilon_{x,t}. \tag{2}$$

The structural shock term, $\epsilon_{x,t}$, is scaled by the parameter σ_{ϵ_x}, which denotes its standard deviation. Since the price indices p_{t+i} are functions of contemporaneous and preceding contract wages, equation (2) implies that in negotiating the current contract wage, agents look at an average of the nominal contract wages that were negotiated in the recent past as well as those that are expected to be negotiated in the near future. In other words, they take into account nominal wages that apply to overlapping contracts. In addition, wage setters take into account expected demand conditions. For example, when they expect demand to exceed potential, $q_{t+i} > 0$, the current contract wage is adjusted upwards relative to contracts negotiated recently or expected to be negotiated in the near future. The parameter γ measures the sensitivity of contract wages to the future excess demand term.

Next, we turn to the relative real wage contracting specification (the RW specification in the following). In this case, wage setters compare the average real wage over the life of their contract with the real wages negotiated on overlapping contracts in the recent past and near future. While this comparison is carried out in real terms, it is still the nominal wage that is negotiated. It remains to define the two elements of this comparison. The average real contract wage is defined using the weighted average of current and future price indices prevailing over the life of the contracts, denoted by $\bar{p}_t$:

$$x_t^{RW} - \mathrm{E}_t \left[\bar{p}_t \right] = x_t^{RW} - \mathrm{E}_t \left[\sum_{i=0}^{3} f_i \, p_{t+i} \right]. \tag{3}$$

To summarize real wages on nearby contracts it is helpful to define an index of real contract wages negotiated on the contracts that are currently in effect:

$$v_t = \sum_{i=0}^{3} f_i (x_{t-i}^{RW} - \bar{p}_{t-i}). \tag{4}$$

The current nominal contract wage under the RW specification is then determined by:

$$x_t^{RW} - \mathrm{E}_t\left[\bar{p}_t\right] = \mathrm{E}_t\left[\sum_{i=0}^{3} f_i\, v_{t+i} + \gamma \sum_{i=0}^{3} f_i\, q_{t+i}\right] + \sigma_{\epsilon_x}\, \epsilon_{x,t}. \tag{5}$$

In this case, agents negotiate the real wage under contracts signed in the current period with reference to the average real contract wage index expected to prevail over the current and the next three quarters. Thus, in negotiating current contracts agents compare the current real contract wage to an average of the real contract wages that were negotiated in the recent past and those expected to be negotiated in the near future. Again, agents also adjust for expected demand conditions and push for a higher real contract wage when they expect output above potential.

For the RW specification a subtle but important question arises with respect to the timing of the price expectations in the real contract wage indices v_{t+i}. For example, the current contract wage x_t depends on the index of real contract wages currently in effect, v_t, which in turn is a function of the real contract wages from periods $t-1$, $t-2$ and $t-3$. The question is whether as in equation (5) the relevant reference points for the determination of the current contract wage are the *ex-post realized real contract wages* from these periods, which are now known to wage setters, or the *ex-ante expected real contract wages*, which formed the basis of the negotiations in earlier periods. To give an example, the average real contract wage from period $t-1$, which enters the index v_t in (4) conditional on period t information, would then be defined as $x_{t-1}^{RW} - (f_0\, p_{t-1} + f_1\, p_t + f_2\, \mathrm{E}_t[p_{t+1}] + f_3\, \mathrm{E}_t[p_{t+2}])$. In period $t-1$, however, the real wage considered in the negotiations was conditioned on period $t-1$ information, $x_{t-1}^{RW} - (f_0\, p_{t-1} + f_1\, \mathrm{E}_{t-1}[p_t] + f_2\, \mathrm{E}_{t-1}[p_{t+1}] + f_3\, \mathrm{E}_{t-1}[p_{t+2}])$.

Since both definitions seem plausible, we will consider both in estimation. We refer to the relative contracting specification with price expectations conditioned on historically available information as the RW-C specification and redefine the equations (4) and (5) accordingly.

Fuhrer and Moore (1995a) discuss the RW and RW-C specification only in the appendix of their paper. Their preferred specification for U.S. data, which is the main focus of their paper, was instead a simplified version of the RW model, which they chose based on a specification test. The simplification concerns the definition of the average real contract wage. Instead of using the average price level expected to prevail over the life of the contracts, $\mathrm{E}_t[\bar{p}_t] = \mathrm{E}_t[\sum_{i=0}^{3} f_i\, p_{t+i}]$, they simply use the current price level, p_t. Thus, the current real wage simplifies to $x_t^{RW} - p_t$ and the index of real contract wages that are in effect, v_t, simplifies to $\sum_{i=0}^{3} f_i (x_{t-i}^{RW} - p_{t-i})$. We refer to this case as the RW-S specification. In this case the index v_t no longer uses price expectations. Consequently, the point regarding the timing of these expectations discussed above is mute.

Before turning to the data used in estimation, we note that although the above specifications are written in terms of the price level, they can be rewritten in terms of the quarterly inflation rate. Thus, either price levels or inflation rates can be used in estimation. Furthermore, we note that the contracting specifications only pin down the steady-state real contract wage, but not the steady-state inflation rate. Steady-state inflation will eventually be determined by the central bank's inflation target, once we close the model in section 6.

3 The Data

The data we use are quarterly series of inflation, the output gap and the short-term nominal interest rate. As noted previously, using price data instead of wage data in estimating staggered contracting specifications is usually justified by linking prices to wages with a simple markup equation. The measures we use for output and prices are real GDP and the GDP deflator. The interest rate is the three-month money market rate. To obtain measures for the Euro area we aggregate over data for ten of the eleven member countries (excluding Luxembourg) using fixed 1993 GDP weights at PPP rates.[9]

The historical path of these Euro-area aggregates between 1974:1 and 1998:4 is shown

[9] This data is drawn from the ECB area-wide model database (see Fagan et al. (1999)) The weights for the 10 Euro-area countries are as follows: Austria (.0306), Belgium (.0392), Finland (.0157), France (.2166), Germany (.3016), Ireland (.0099), Italy (.2043), the Netherlands (.0541), Portugal (.0234) and Spain (.1045).

in **Figure 1**. As shown in the top left panel average inflation in the Euro area steadily declined over the last 25 years. Similarly, the average short-term nominal interest rate depicted in the top right panel tended to decline from the mid 1980s onwards except for the EMS crises in the early 1990s. This downward trend is a unique feature of Euro-area data and complicates the empirical investigation of European inflation dynamics relative to similar analyses for the United States. We will return to this issue below.

The contracting model in section 2 relates the short-run dynamics of inflation to the output gap. While a measure for actual real GDP in the Euro area is available and shown in the bottom left panel of **Figure 1** (solid line), we need to estimate the unobservable potential output. Constructing a structural estimate of potential for the Euro area prior to EMU goes beyond the objective of this paper. Even for the individual member countries this would be rather difficult. A common alternative estimate of potential used in the macroeconomic modelling literature is the log-linear trend (see for example Fuhrer and Moore (1995a) and Taylor (1993a) among many others), which is shown as the dashed line in the bottom left panel. The bottom right panel compares the output gap implied by the log-linear trend to the OECD's (1999) estimate of the Euro-area output gap (dotted line). Since these estimates are surprisingly similar, except for a small difference in the 1990s, we will follow Taylor and Fuhrer and Moore in using output gaps based on a log-linear trend for estimating the overlapping contracts model.[10]

The source of the downward trend in Euro-area inflation noted previously is directly apparent from **Figure 2**. As shown in the top left panel inflation rates in the early 1970s were much higher in countries such as France and Italy than in Germany due to oil price shocks and accommodative monetary policy. It took 10 to 15 years, respectively, for French and Italian inflation rates to decline to German levels. Convergence in inflation rates was accompanied by convergence in nominal interest rates in the late 1990s as can be seen from

[10] Other alternatives include estimates based on the HP filter or unobserved components methods. We have conducted some sensitivity studies in this respect. We stick with the linear trend as our benchmark for comparability with the results obtained by Fuhrer and Moore and Taylor for the U.S. and because of the similarity to the OECD estimate of the Euro-area output gap.

the upper right panel of **Figure 2**. Over time, as economic convergence and the future formation of a monetary union became more widely expected the inflation premium incorporated in Italian and French short-term nominal interest rates relative to German rates eventually disappeared. This convergence process and the role of the European Monetary System (EMS) in its context have been widely debated and analyzed in the academic and policy literature of the last decade. There is little doubt that the decline of inflation has largely been due to the growing commitment on the part of monetary policy makers in the Euro area to achieve and maintain low inflation. The credibility of this commitment, however, likely varied over time, probably being rather low in the early stages of the EMS in the early and mid 1980s and higher during the "hard" EMS period in the late 1980s up to the EMS crises in 1992 and 1993. Following these crises credibility regarding the central banks' commitment to achieve low inflation likely increased again with the progress of preparations for EMU. To the extent that disinflation during these periods was credible and expected by wage and price setters, the associated output losses should have been rather low. In fact, a casual comparison of the extent of disinflation in Italy and France relative to Germany and the output gap estimates for these three economies based on the log-linear trends that are shown in the bottom right panel of **Figure 2**, suggests that the disinflations in Italy and France did not require large and protracted recessions and thus may have been partly anticipated.

Conceivably, one could attempt to model and estimate the processes of expectations and credibility regarding policy targets and exchange rate pegs throughout the 1980s and 1990s explicitly in the context of a complete macroeconomic model. This would go beyond the objective of our paper. Instead we simply approximate the implicit time-varying inflation objective with a linear trend, and then estimate the overlapping contracts models using inflation deviations from this trend. We detrend the average inflation rate for the Euro area as well as the French and Italian inflation rates in this manner. Similar approaches have been used by other researchers with regard to European inflation data, notably Gerlach and

Svensson (1999) and Cecchetti, McConnell and Quiros (1999).[11] This approach would be appropriate, if the source of the disinflation had been a fully expected and credible reduction in the policymakers' inflation target having been gradually phased in. However, given this was not the case, this approach introduces an error that may bias our estimation results. In particular, it could lead us to underestimate the degree of inflation persistence and possibly understate the case for the relative real wage specification. We return to this question later on.

4 Empirical Inflation and Output Dynamics

Our empirical analysis proceeds in two stages. In the first stage, we estimate an unconstrained bivariate VAR model of Euro-area output and inflation. In the second stage, we use this unconstrained VAR as an auxiliary model in estimating the structural overlapping wage contracting specifications by indirect inference methods. These methods are a simulation-based procedure for calibrating the parameters of the structural model by matching its reduced form, which corresponds to a constrained VAR, as closely as possible with the estimated unconstrained VAR model.

The unconstrained VAR provides an empirical summary description of Euro-area inflation and output dynamics.[12] We estimate the short-run dynamics jointly with a deterministic linear trend for inflation and the logarithm of output over the sample period. Following Fuhrer and Moore (1995a) we then compute the autocorrelation functions implied by the VAR including the associated asymptotic confidence bands.[13]

These autocorrelation functions indicate that the lead-lag relationship between inflation

[11] Gerlach and Svensson use an exponential trend for the Euro-area inflation rate in estimating a P-star model of inflation dynamics á la Hallman, Porter and Small (1991) for the Euro area. Cecchetti et al. construct inflation and output deviations from a 12-month moving average of actual values and estimate inflation-output tradeoffs based on this data for a number of Euro-area economies.

[12] Although interest rates are important determinants of output and inflation, we restrict attention to bivariate VARs without including an interest rate, primarily because it is unclear what would be an appropriate interest rate for the Euro area. We return to this problem later on in section 6 when estimating an aggregate demand equation that closes the small macroeconomic model.

[13] For a detailed discussion of the methodology and the derivation of the asymptotic confidence bands for the estimated autocorrelation functions the reader is referred to Coenen (2000).

and output is consistent with a short-run tradeoff, that is, with a short-run Phillips curve. Furthermore, the estimated autocorrelation functions constitute a benchmark against which we can evaluate the ability of the alternative overlapping contracts specifications to explain the dynamics of inflation in Euro-area data. Such an approach has also been recommended by McCallum (1999), who argued that autocovariance and autocorrelation functions are a more appropriate device for confronting macroeconomic models with the data than impulse response functions because of their purely descriptive nature.

The empirical model for output and inflation, written in terms of the level of inflation, Π_t and the logarithm of output, Q_t, corresponds to

$$\begin{bmatrix} \Pi_t \\ Q_t \end{bmatrix} = \begin{bmatrix} a_{0,\Pi} \\ a_{0,Q} \end{bmatrix} + \begin{bmatrix} a_{1,\Pi} \\ a_{1,Q} \end{bmatrix} t + \begin{bmatrix} \pi_t \\ q_t \end{bmatrix}, \tag{6}$$

where π_t and q_t refer to the inflation and the output gap respectively, which are determined by an unconstrained VAR of lag order 3:

$$\begin{bmatrix} \pi_t \\ q_t \end{bmatrix} = A_1 \begin{bmatrix} \pi_{t-1} \\ q_{t-1} \end{bmatrix} + A_2 \begin{bmatrix} \pi_{t-2} \\ q_{t-2} \end{bmatrix} + A_3 \begin{bmatrix} \pi_{t-3} \\ q_{t-3} \end{bmatrix} + \begin{bmatrix} u_{\pi,t} \\ u_{q,t} \end{bmatrix}. \tag{7}$$

The A_i matrices ($i = 1, 2, 3$) contain the coefficients on the first three lags of the inflation and the output gap.[14] The error terms $u_{\pi,t}$ and $u_{q,t}$ are assumed to be serially uncorrelated with mean zero and positive definite covariance matrix Σ_u.

We fit this model to the aggregated output and inflation data for the Euro area as a whole for the period from 1974:Q1 to 1998:Q4. Since we are merely interested in the parameters of the VAR model (7), we proceed in two steps. First, we detrend the data by a simple projection technique and then we estimate the parameters of the V AR model, that is the coefficient matrices A_i and the covariance matrix Σ_u by Quasi-Maximum-Likelihood (QML) methods.[15] The estimates of the unconstrained VAR model are shown in **Table 1**. Standard lag selection procedures based on the HQ and SC criteria suggest that a lag order of 2 would be sufficient to capture the empirical inflation and output dynamics. The Ljung-Box Q(12) statistic indicates serially uncorrelated residuals with a marginal probability

[14] Here, we use a maximum lag order of 3, simply because this corresponds to the reduced-form VAR representation of the overlapping contract models of section 2 with a contract length of 4 quarters.

[15] For some more detail refer to the appendix, section A.2.

value of 42.8%. The estimates of the parameters of the VAR(2) model are shown in panel A of **Table 1**. Our point estimates imply that the smallest root of the characteristic equation $\det(I_2 - A_1 z - A_2 z^2) = 0$ is 1.2835, thereby suggesting that the inflation and output gap are stationary. Our findings are supported by the results of standard univariate Dickey-Fuller tests for the presence of unit roots.

We also estimate an unconstrained VAR(3) model. This is of interest, because all contracting specifications discussed in section 2 have a reduced form which is a constrained VAR of order 3 if the maximum contract length is one year. To assess the sensitivity of our results to the lag length, we will use the VAR(2) and VAR(3) models in parallel in the estimation of the contracting specifications in the following section. On a statistical basis, the third lag would not be absolutely necessary, as can be seen from panel B in **Table 1**, which shows that the A_3 coefficients are insignificant.

The autocorrelations functions associated with the unconstrained VAR(3) model of the Euro area are depicted in **Figure 3**. The diagonal elements show the autocorrelations of the detrended inflation rate and the output gap, the off-diagonal elements the lagged cross correlations, The solid lines represent the point estimates, while the dotted lines indicate 95% confidence bands. Both inflation and output are quite persistent with positive autocorrelations out to lags of about 5 and 8 quarters which are highly significant. The cross correlations in the off-diagonal panels confirm much of conventional wisdom about inflation and output dynamics. For example, in the second panel of the top row, a high level of output is followed by a high level of inflation a year later and again these cross correlations are statistically significant. In the first panel of the bottom row a high level of inflation is followed by a low level of output a year later. These lead-lag interactions are highly indicative of the existence of a conventional short-run tradeoff between output and inflation. All in all these correlations are stylized facts which any structural model of output and inflation dynamics ought to be able to explain.

The results for the bivariate VARs of order 3 for France, Germany and Italy are summarized in **Table 2**.[16] Again, the VAR model seems fairly successful at summarizing the observed inflation and output dynamics. Note that for Germany the estimates are obtained without a trend in inflation, ($a_{1,\Pi} = 0$). The estimated autocorrelation functions for output and inflation in France and Italy display qualitative similar characteristics than for the Euro area as a whole, in particular regarding the persistence in inflation and output variations. The cross correlations, however, are somewhat smaller. As to Germany, the degree of persistence in inflation is substantially lower and the correlations between current output and lagged inflation have the opposite sign, albeit statistically insignificant. We return to these issues in the next section, when we use the empirical autocorrelation functions as a benchmark to evaluate the fit of alternative structural overlapping contracts specifications.

5 Estimating the Overlapping Contracts Specifications

In the following we use the unconstrained VAR models as approximating probability models in estimating the coefficients of the different overlapping contracts specifications discussed in section 2. As can be seen from equations (1) through (5) there are only three structural parameters to estimate for each specification: (i) the slope of the contracting distribution s that determines the series of contract weights f_i; (ii) the sensitivity of the contract wage to expected future aggegrate demand over the life of the contract γ; and (iii) the standard deviation of the contract wage shock σ_{ϵ_x}.

Of course, the overlapping contracts specifications discussed in section 2 do not represent a complete model of inflation determination. Since the contract wage equations (2) and (5) contain expected future output gaps, we need to specify how the output gap is determined in order to solve for the reduced-form representation of inflation and output dynamics under each of the contracting specifications. A full-information estimation approach would require a complete macroeconomic model and estimate all the models's structural parameters jointly. A simple version of such a model would include an aggregate demand equation,

[16] To save space, we do not report separate figures for the estimated autocorrelation functions. Instead, see the dotted lines in **Figures 5** to **7**, which depict the accompanying asymptotic confidence bands.

which relates output gaps to ex-ante long-term real rates, as well as a Fisher equation, a term structure relationship and a monetary policy rule. While our ultimate objective in this paper is to build precisely such a model, we take a less ambitious approach in estimating the contracting parameters. We simply use the output gap equation from the unconstrained VAR models, which corresponds to the second row in (7), for output determination. This limited-information approach is close to the estimation approaches used by Taylor (1993a) and Fuhrer and Moore (1995a) and is likely to be more robust than a full-information approach. We estimate the aggregate demand equation later on by single-equation methods and discuss those results in the next section.

Using the output equation from the unconstrained VAR together with the wage-price block from section 2, we can solve for the reduced-form inflation and output dynamics under each of the four different contracting specifications (RW, RW-C, RW-S and NW).[17] For this purpose it is convenient to rewrite the wage-price block, which was originally defined in levels of nominal contract wages and prices, in terms of the real contract wage $xp_t = x_t - p_t$ and the annualized quarterly inflation rate π_t. The reduced-form solution of this rational expectations model is a trivariate *constrained* VAR. While the quarterly inflation rate π_t and the output gap q_t are observable variables, the real contract wage xp_t is unobservable. For a contracting specification with a one-year maximum contract length this constrained VAR can be written as

$$\begin{bmatrix} xp_t \\ \pi_t \\ q_t \end{bmatrix} = B_1 \begin{bmatrix} xp_{t-1} \\ \pi_{t-1} \\ q_{t-1} \end{bmatrix} + B_2 \begin{bmatrix} xp_{t-2} \\ \pi_{t-2} \\ q_{t-2} \end{bmatrix} + B_3 \begin{bmatrix} xp_{t-3} \\ \pi_{t-3} \\ q_{t-3} \end{bmatrix} + B_0\, \epsilon_t, \tag{8}$$

where the B_i matrices $(i = 0, 1, 2, 3)$ contain the coefficients of the constrained VAR and ϵ_t is a vector of serially uncorrelated error terms with mean zero and positive (semi-) definite covariance matrix which is assumed to be diagonal with its non-zero elements normalized to unity.

The coefficients in the bottom row of the B_i matrices coincide exactly with coefficients

[17] We assume that ouput and wage expectations in the contract wage equations are formed rationally, and use the AIM algorithm of Anderson and Moore (1985) for linear rational expectations models to solve for the reduced form dynamics.

of the output equation of the unconstrained VAR, with the B_0 cofficients obtained by means of a Choleski decomposition of the covariance matrix Σ_u. The reduced-form coefficients in the upper two rows of the B_i matrices, which are associated with the determination of the real contract wage and inflation, are functions of the structural parameters $(s, \gamma, \sigma_{\epsilon_x})$ as well as the coefficients of the output equation of the unconstrained VAR.

We estimate the structural parameters of the overlapping contracts specifications s, γ and σ_{ϵ_x} using the indirect inference methods proposed by Smith (1993) and Gouriéroux, Monfort and Renault (1993) and developed further in Gouriéroux and Monfort (1995, 1996). The estimation procedure, including its asymptotic properties, is discussed in detail in the appendix of this paper. In the appendix we also compare this procedure to the Maximum-Likelihood methods used by Taylor (1993a) and Fuhrer and Moore (1995a).

Indirect inference is a simulation-based procedure for calibrating a structural model with the objective of finding parameter values such that its dynamic characteristics match the dynamic properties of the observed data as summarised by an approximating probability model. The latter should fit the empirical dynamics reasonably well, but need not necessarily nest the structural model. In the case at hand, the unconstrained VAR models discussed in section 4 are natural candidates for such an approximating probability model.

For given values of the structural parameters $(s, \gamma, \sigma_{\epsilon_x})$ and the parameters of the output equation from the unconstrained VAR model (7), we simulate the reduced form of the structural model, that is the constrained VAR model (8), to generate "artificial" series for the real contract wage, the inflation rate and the output gap. All that is needed for simulation are three initial values for each of these variables and a sequence of random shocks.[18] Subsequently we fit the unconstrained VAR model to the artificial series of inflation and the output gap and match the simulation-based estimates of the inflation equation as closely as possible with the empirical estimates by searching over the feasible space of the structural

[18]In estimation we use steady-state values as initial conditions and are careful to only use simulation data for later periods that are essentially unaffected by this choice of initial conditions. This issue is discussed in more detail in the appendix, section A.3.

parameters.[19]

Euro-area estimation results for the baseline version of the relative real wage contracting model (RW), the version with price expectations conditioned on historically available information (RW-C), the simplified version preferred by Fuhrer and Moore (RW-S) and the nominal wage contracting model (NW) are reported in **Table 3**. As a sensitivity check we consider both the VAR(2) and VAR(3) models estimated in section 4 as approximating probability models.[20] The estimation results indicate that all four contracting models fit the Euro-area inflation dynamics reasonably well, in particular when we allow for a maximum contract length of one year and thus three lags in the VAR. As can be seen from the standard errors given in parentheses, the estimates of the structural parameters are in almost all cases statistically significant, with the appropriate sign and economically significant magnitude.

We also compute the probability (P-) values of the test for the over-identifying restrictions, which were imposed when estimating the structural parameters. According to this test, none of the four contracting specifications is rejected by the data, when we use the VAR(3) as approximating probability model and allow for a one-year maximum contract length. When we use the VAR(2) model and constrain the maximum contract length to three quarters, both, the RW-C and the RW-S specification can be rejected at convenient confidence levels, but not the RW or the NW specifications. Though the estimates of the real wage contracting specifications are not directly comparable, since these imply structures with different degrees of forward-looking behaviour, it is worthwile to note that the RW-S specification reveals stronger rigidities than the RW and the RW-C specifications as reflected by the smaller estimates of the slope parameter s of the contracting distributions. While neither the RW nor the NW specification can be rejected, we use the associated

[19] We do not need data for the unobservable real contract wage since the unconstrained VAR is only fitted to the observable data for inflation and the output gap.

[20] In the case of the VAR(2) model, we restrict the maximum contract length in the structural contracting specification to three quarters instead of one year, such that its lag order corresponds to that of the structural model's reduced-form solution. In this case, the slope parameter s is restricted to lie in the interval $(0, 1/3]$. Note, because of the difference in its domain the magnitude of the slope parameter will not be directly comparable across the specifications with three-quarter and one-year maximum contract length, respectively.

minimum value of the criterion function to discriminate between these two specifications. In the case of our preferred setup with one-year maximum contract length and, thus, the VAR(3) model, the RW specification implies a higher P-value than the NW specification. For the estimation based on the VAR(2) model, however, the NW specification entails a higher P-value.

In sum, our findings for the Euro area differ quite a bit from Fuhrer and Moore (1995a), who reject the nominal wage contracting model for U.S. data and find that the RW-S specification of the relative wage contracting model fits the U.S. inflation dynamics better than the theoretically more plausible RW specification.

To provide further insight regarding our estimation results, we compare the autocorrelation functions implied by the constrained VAR(3) representation of each of the four contracting models with the autocorrelation functions from the unconstrained VAR. As shown in **Figure 4**, the autocorrelation functions for all four models tend to remain inside the 95% confidence bands (dotted lines) associated with the autocorrelation functions of the unconstrained VAR. The three relative real wage contracting specifications (RW: solid line with bold dots, RW-C: dash-dotted line, RW-S: solid line) are rather similar. They exhibit substantial inflation peristence and quite pronounced cross correlations that are indicative of a short-run Phillips curve tradeoff. The upper right panel indicates that high levels of output are followed by high inflation, while the lower left panel shows that high levels of inflation are followed by low levels of output. The only noticeable difference from the unconstrained VAR, is that the latter set of cross correlations are somewhat larger in absolute magnitude for the constrained VAR. The autocorrelations for the nominal contracting model (NW: dashed line) indicate a lower degree of inflation persistence and less pronounced cross correlations than for the different relative wage contracting models.

As noted in the introduction to this paper, the estimation results with Euro-area data may be questioned for a number of reasons. First of all, the data are artificial in the sense that they are only averages of the data from the member economies prior to the formation of EMU. Furthermore, the member economies experienced different monetary policy regimes. While Germany enjoyed stable inflation with fairly predictable monetary policy, other countries such as France and Italy experienced a long-drawn out convergence process, which was not fully anticipated by market participants. As a result, Euro-area inflation data exhibits a long decline which we removed from the data by subtracting a linear trend.

Thus, to investigate the validity of our results, we also estimate the different contracting models for France, Germany and Italy separately. The results are summarized in **Table 4**. Here we only focus on the case of the VAR(3) model. For France we reject the RW-C and the RW-S specifications, but not the RW and the NW specification. The NW model exhibits the highest P-value. However, in this case the parameter measuring the sensitivity to aggregate demand, γ, is statistically insignificant. The parameter estimates for the RW specification are significant and relatively close to the values obtained for the Euro area. For Italy, which experienced the most dramatic transition process, the estimation of the NW model did not converge. Instead, the RW and the RW-C model seem to fit Italian inflation data reasonably well and imply statistically significant parameter estimates. For Germany, where inflation exhibited no long-run trend, we find that all three relative real wage contracting models are strongly rejected by the data. While the nominal contracting model is also rejected, it does fit better in the sense of implying a higher P-value. The parameter estimates for the NW model with German data are surprisingly close to the NW estimates obtained with Euro area data.

Again, a promising alternative approach for evaluating the fit of the RW and NW specifications is to compare the autocorrelation functions of the constrained and unconstrained VAR models. As shown in **Figure 5** for France, the RW specification does better than the NW specification in terms of fitting the inflation persistence in the top left panel, but worse in terms of fitting the cross-correlations in the diagonal panels. In the case of Italy the RW model comes very close to matching the empirical autocorrelations of inflation in the top left panel of **Figure 7**, and also does reasonably well with regard to the cross correlations. The results for Germany in **Figure 6** are, as expected, quite different. The autocorrelation functions of the unconstrained VAR (solid line with dotted confidence interval) indicated a much smaller degree of inflation persistence and a counterintuitive, albeit statistically insignificant, positive correlation between output and lagged inflation. Consequently, the nominal contracting model has a better chance at fitting German inflation dynamics than the relative contracting models.

We conclude from these results that, both the RW and the NW specifications are plausible alternatives for the Euro area. On the one hand, the estimation with aggregated Euro-area data indicates a slight preference for the relative wage contracting model. On the other hand, the comparison between France, Germany and Italy suggests that this preference may partly be due to the high-inflation regime in countries such as France and Italy and the fact that the subsequent long-run decline in inflation was not fully anticipated. Thus, an optimist would conclude that the independent European Central Bank will likely face a similar environment in the future as the Bundesbank did in Germany. In this case, the inflation-output relationship is best characterized by the nominal contracting specification with the parameter estimates obtained with German data. A pessimist, who suspects that stabilizing Euro-area inflation will require higher output losses, would instead prefer to use the RW specification with parameter estimates based on Euro-area data. A robust monetary policy strategy, however, should perform reasonably under both specifications. In the remainder of the paper we close the model by estimating an aggregate demand equation and a term structure relationship and then evaluate the performance of Taylor's under these two alternative wage contracting specifications.

6 Closing the Model: Output Gaps and Interest Rates

We model aggregate demand with a simple reduced-form IS equation, which relates the current output gap, q_t, to two lags of itself and to the lagged long-term ex-ante real interest rate, r^l_{t-1}:

$$q_t = \delta_0 + \delta_1\, q_{t-1} + \delta_2\, q_{t-2} + \delta_3\, r^l_{t-1} + \sigma_{\epsilon_d}\, \epsilon_{d,t}. \tag{9}$$

$\epsilon_{d,t}$ denotes an unexpected demand shock re-scaled with the parameter σ_{ϵ_d}, which measures the standard deviation of the demand shock. The rationale for including lags of output is to account for habit persistence in consumption as well as adjustment costs and accelerator effects in investment. We use the lagged instead of the contemperaneous value of the real interest rate to allow for a transmission lag of monetary policy. For now we neglect the possibility of effects of the real exchange rate and expected future income on aggregate demand.[21] Fuhrer and Moore (1995b) found that a similiar aggregate demand specification fits U.S. output dynamics quite well.

Next we turn to the financial sector and relate the long-term ex-ante real interest rate, which affects aggregate demand, to the short-term nominal interest rate, which is the principal instrument of monetary policy. Three equations determine the various interest rates in the model. The short-term nominal interest rate, i_t^s, is set according to a Taylor-type interest rate rule (see Taylor (1993b)). This rule incorporates policy responses to inflation deviations from target and output deviations from potential output, and allows for some degree of partial adjustment:

$$i_t^s = \alpha_r \, i_{t-1}^s + (1 - \alpha_r)(r^* + \pi_t^{(4)}) + \alpha_\pi (\pi_t^{(4)} - \pi^*) + \alpha_q \, q_t. \tag{10}$$

r^* denotes the long-run equilibrium real rate, while π^* refers to the policymaker's target for inflation. The inflation measure is the four-quarter moving average of the annualized quarterly inflation rate, that is $\pi_t^{(4)} = \frac{1}{4} \sum_{j=0}^{3} \pi_{t-j} = p_t - p_{t-4}$, and the interest rate is annualized.

As to the term structure, we rely on the accumulated forecasts of the short rate over two years which, under the expectations hypothesis, will coincide with the long rate forecast for this horizon. The term premium is assumed to be constant and equal to zero:

$$i_t^l = \mathrm{E}_t \left[\frac{1}{8} \sum_{j=0}^{7} i_{t+j}^s \right]. \tag{11}$$

[21] In the future, we plan to investigate how important these two factors are for the determination of Euro area aggregate demand. Since the Euro area is a large, relatively closed economy, the exchange rate is likely to play a less important role than it did in the individual member economies prior to EMU.

By subtracting inflation expectations over the following 8 quarters, we then obtain the long-term ex-ante real interest rate:

$$r_t^l = i_t^l - \mathrm{E}_t\left[\frac{1}{2}(p_{t+8} - p_t)\right]. \tag{12}$$

To estimate the parameters of the aggregate demand equation (9) we first construct the ex-post real long-term rate as defined by equations (11) and (12) but replacing expected future with realized values. We then proceed to estimate the parameters by means of Generalized Method of Moments (GMM) using lagged values of output, inflation and interest rates as instruments. The estimation results are reported in **Table 6**. The sample period for this estimation is 1974:4 to 1998:4.

Panel A refers to the estimates for the Euro area. In the first row, the output gap and interest rate data are area-wide GDP-PPP-weighted averages. The coefficients on the two lags of the output gap are significant and exhibit an accelerator pattern. The interest rate sensitivity of aggregate demand has the expected negative sign, however the parameter estimate is only borderline significant and rather small. It is not clear however, what is the appropriate real interest rate measure for the Euro area. For example, instead of GDP weights it may be a more appropriate to use the relative weights in debt financing. Or, one could make that the argument that the relevant rate for the Euro area was the German interest rate. After all, movements in German interest rates presumably had to be mirrored eventually by the other countries to the extent that they intended to maintain exchange rate parities within the European Monetary System. For this reason we also use the German real interest rate to estimate the interest rate sensitivity of Euro-area aggregate demand. In this case, as shown in the second row, we find similar coefficients on the lags of the output gap, but the estimate of the interest rate sensitivity is highly significant an d three times as large.

We have subjected this specification of aggregate demand to a battery of sensitivity tests. For example, we have investigated alternative specifications of potential output, alternative horizons on the term structure equation, including the use of average long-term rates instead of a term structure based on short-term rates and we have varied the length of the sample period. At least qualitatively the estimation results remain the same.

For comparison, we have also estimated the same specification for France, Germany and Italy. In each case we use the domestic real interest rate. For France and Italy we obtain qualitatively similar estimates as for the Euro area. For Germany however, the estimate of the interest rate sensitivity is not significant and the lags of output do not exhibit an accelerator-type pattern.

It remains to discuss the deterministic steady state of this model. In steady state, the output gap is zero and the long-term real rate is equal to the equilibrium real rate, r^*. This equilibrium rate is determined as a function of the parameters of the aggregate demand curve (9), s.t. $r^* = \delta_0/\delta_3$. The steady-state value of inflation is determined exclusively by monetary policy. Since the overlapping contracts specifications of the wage-price block do not impose any restriction on the steady-state inflation rate, steady-state inflation will be equal to the inflation target, π^*, in the policy rule.

7 Evaluating Monetary Policy Rules for the Euro Area

In a recent paper, Gerlach and Schnabel (1999) argue that average interest rates in the EMU countries in 1990-98, with the exception of the period of exchange market turmoil in 1992-93, moved very closely with average output gaps and inflation as suggested by the Taylor rule. Here, we explore the inflation and output dynamics that would arise in our model of the Euro area under such a rule. To this end we present the impulse responses of inflation and output to unexpected demand and supply shocks and also simulate a disinflation. In this exercise, the coefficients in the interest rate rule (10) are set equal to the values proposed in Taylor (1993), that is $\alpha_r = 0$, $\alpha_\pi = 0.5$ and $\alpha_q = 0.5$. The contract wage specification used is the RW-S specification preferred by Fuhrer and Moore. For aggregate demand we use the equation estimated with the German real interest rate. In each case, initial conditions correspond to the steady state.

We start with an unexpected short-run supply shock, that is in our framework, a shock to the contract equation. The standard deviation of this type of shock was estimated as part of the structural estimation of the overlapping contracts model in section 5. The effect of such standard deviation shock under a Taylor rule is depicted in **Figure 8.** The four panels shows deviations of the output gap, the inflation rate, the short-term nominal interest rate as well as the long-term nominal and real rates from steady state. The shock occurs in period 10. As a result inflation increases over the next four quarters by almost a full percentage point.

Monetary policy responds to this increase in inflation by raising short-term nominal interest rates sufficiently so as to increase the long-term real interest rate. This policy tightening induces a slowdown in aggregate demand which lasts for about four years. Since future aggregate demand affects contract wage setting and thru this channel the inflation rate, inflation eventually returns back to target and even undershoots for a few periods. The aggressiveness of the policy response, and the depth of the resulting shortfall of aggregate demand are directly influenced by the relative strength of the inflation and output response coefficents in the interest rate rule. The policymaker is faced with a tradeoff, namely whether to achieve a faster reduction in the inflation deviation by accepting a larger output deviation. Due to the assumption of rational expectations this tradeoff depends on all the model parameters.

As shown in **Figure 9**, in the case of an unexpected demand shock the policymakers' is faced with a simpler decision. As shown in the top left panel in response to a 1 percent surprise increase in demand, the output gap increases by a little more than a full percent over the first three quarters and then returns back to potential over the course of the next two years. As contract wage setting setting takes into account expected future aggregate demand conditions this improvement in expected demand leads to an increase in inflation of about one percentage point over the same period. Monetary policy responds by raising short nominal rates enough to increase the long-term real rate and return output to potential and inflation to target. Of course, the higher the coefficients in the policy rule, the faster both output and inflation would return to steady state. There is no conflict of interest in that respect. One concern, which may restrain policymakers from reacting much more aggressively, is regarding the degree of interest rate variability.

Finally, **Figure 10** reports a disinflation by 2 percentage points. It is modelled as a fully credible change in the policymaker's inflation target, π^* in the policy rule (10), from 2 percent to 0 percent. Initial conditions are consistent with a steady state inflation rate of 2 percent. As can be seen in the lower-left, the reduction in the inflation target leads to an increase in short-term nominal interest rates. As a result, the long-term real rate increases and the output gap turns negative. The slowdown in aggregate demand induces disinflationary conditions and the inflation rate declines to the new steady state level over the course of 2 1/2 years. The change in policy target also has a direct effect on inflation thru inflation expectations, which accelerates the disinflation and reduces the associated output loss. This effect is directly apparent from the behavior of long-term nominal and real interest rates. Because the disinflation is fully credible, long-term nominal rates which embody expectations of future short rates decline throughout the full perod. Nevertheless, the initial increase in short rates still implies an initial increase in the long-term real rate.

8 Conclusions

In this paper we estimate a small-scale empirical model of the Euro area where short-run real effects of monetary policy arise due to overlapping wage contracts. The paper primarily focussed on exploring the empirical fit of alternative overlapping wage contract specifications, because they are the key ingredient for the short-run inflation-output variability tradeoff faced by monetary policymakers. We investigate the nominal wage contracting model due to Taylor (1980) as well as three different versions of the relative real wage contracting model first proposed by Buiter and Jewitt (1981) and investigated empirically with U.S. data by Fuhrer and Moore (1995a). Contrary to Fuhrer and Moore, who reject the nominal contracting model and prefer a version of the relative contracting model which induces a higher degree of inflation persistence, we find that both types of contracting models fit the data for the Euro area. The best fitting specification, however, is a version of the relative contracting model, which is theoretically more plausible than the simplified version preferred by Fuhrer and Moore.

To investigate the validity of our results, we also estimate the contracting models for France, Germany and Italy separately. We find that the relative contracting model dominates in countries which transitioned out of a high inflation regime such as France and Italy, while the nominal contracting model fits German data better. Thus, an optimist may conclude that the independent European Central Bank will face a similar environment in the future as the Bundesbank did in Germany and pick the nominal contracting specification, while a pessimist, who suspects that stabilizing Euro area inflation will require higher output losses, may want to pick the relative contracting specification. A robust monetary policy strategy, however, should perform reasonably well in both cases. The use of this type of model for policy analysis was illustrated by investigating its impulse responses to demand and price shocks and a disinflation under Taylor's interest rate rule.

References

Amato, J. and T. Laubach, 1999, Monetary policy in an estimated optimization-based model with sticky prices and wages, unpublished working paper.

Anderson, G. S., 1987, A procedure for differentiating perfect-foresight-model reduced-form coefficients, Journal of Economic Dynamics and Control, 11, 465-81.

Anderson, G. S. and G. R. Moore, 1985, A linear algebraic procedure for solving linear perfect foresight models, Economics Letters, 17, 247-52.

Angeloni, I. and L. Dedola, 1999, From the ERM to the Euro: New evidence on economic and policy convergence among EU countries, European Central Bank, Working Paper No. 4, May.

Angeloni, I., V. Gaspar and O. Tristani, 1999, The monetary policy strategy of the ECB, in, D. Cobham and G. Zis, eds., From EMS to EMU, London: Macmillan.

Bernanke, B. and F. S. Mishkin, 1997, Inflation targeting: A new framework for monetary policy? Journal of Economic Perspectives, 11(2), 97-116.

Blanchard, O. J. and C. Kahn, 1980, The solution of linear difference models under rational expectations, Econometrica, 48(5), 1305-1311.

Branch, M. A. and A. Grace, 1996, Optimization toolbox: For use with MATLAB, Natick: The Math Works.

Brayton, F., A. Levin, R. W. Tryon and J. C. Williams, 1997, The evolution of macro models at the Federal Reserve Board, Carnegie-Rochester Conference Series on Public Policy, 42, 115-67.

Bryant, R. C., P. Hooper and C. Mann, eds., 1993, Evaluating policy regimes: New research in empirical macroeconomics, Washington DC: Brookings Institution.

Buiter, W. H. and I. Jewitt, 1981, Staggered wage setting with real wage relativities: Variations on a theme of Taylor, The Manchester School, 49, 211-228, reprinted in, W. H. Buiter (ed.), 1989, Macroeconomic Theory and Stabilization Policy, Manchester University Press.

Cecchetti, S., M. McConnell and G. Perez Quiros, 1999, Policymakers' revealed preferences and the output-inflation variability tradeoff: Implications for the European System of Central Banks, Manuscript, June.

Clarida, R., J. Gali and M. Gertler, 1998, Monetary policy rules and macroeconomic stability: Evidence and some theory, forthcoming, Quarterly Journal of Economics.

Clarida, R., J. Gali and M. Gertler, 1998, ???, European Economic Review.

Clarida, R., J. Gali and M. Gertler, 1999, The science of monetary policy, forthcoming, Journal of Economic Literature.

Coenen, G., 2000, Asymptotic confidence bands for the estimated autocovariance and autocorrelation functions of vector autoregressive models, European Central Bank, Working Paper No. 9, January.

Duffie, D. and K. J. Singleton 1993, Simulated moments estimation of Markov models of asset prices, Econometrica, 61, 929-952.

Erceg, C., D. Henderson, and A. Levin, 1999, Optimal monetary policy with staggered wage and price contracts, forthcoming in, Journal of Monetary Economics.

Fagan, G., J. Henry and R. Mestre, 1999, An area-wide model (AWM) for the EU11, European Central Bank, mimeo.

Fuhrer, J. C., 1994, Optimal monetary policy and the sacrifice ratio, in J. C. Fuhrer (ed.), Goals, guidelines, and constraints facing monetary policymakers, Boston: Federal Reserve Bank of Boston.

Fuhrer, J. C., 1997a, Inflation/output variance trade-offs and optimal monetary policy, Journal of Money, Credit, and Banking, 29(2), 214-234.

Fuhrer, J. C., 1997b, Towards a compact, empirically-verified rational expectations model for monetary policy analysis, Carnegie-Rochester Conference Series on Public Policy, 47, 197-230.

Fuhrer, J. C., and G. R. Moore, 1995a, Inflation persistence, Quarterly Journal of Economics, 110(1), 127-59.

Fuhrer, J. C. and G. R. Moore, 1995b, Monetary policy trade-offs and the correlation between nominal interest rates and real output, American Economic Review, 85, 219-239.

Gali, J. and M. Gertler, 1999, Inflation dynamics: A structural econometric analysis, Journal of Monetary Economics, 44, 195-222.

Gallant, R. A. and G. Tauchen, 1996, Which Moments to Match? Econometric Theory, 12, 657-681.

Gerlach, S. and L.E.O. Svensson, 1999, Money and inflation in the Euro area: A case for monetary indicators? Manuscript, July.

Gerlach, S. and G. Schnabel, 1999, The Taylor rule and interest rates in the EMU area, Manuscript, July.

Goodfriend, M. and R. King, 1997, The new neoclassical synthesis and the role of monetary policy, NBER Macroeconomics Annual.

Gouriéroux, C. and A. Monfort, 1995, Testing, encompassing and simulating dynamic econometric models, Econometric Theory, 11, 195-228.

Gouriéroux, C. and A. Monfort, 1996, Simulation-based inference methods, Oxford: Oxford University Press.

Gouriéroux, C., A. Monfort and E. Renault, 1993, Indirect inference, Journal of Applied Econometrics, 8, Special Issue: Econometric Inference Using Simulation Techniques, S85-S118.

Huh, C. and K. Lansing, 1999, Expectations, credibility and disinflation in a small macroeconomic model, forthcoming in, Journal of Economics and Business.

Levin A., 1991, The macroeconomic significance of nominal wage contract duration, Discussion Paper No. 91-08, San Diego: University of California.

Levin A., V. Wieland and J.C. Williams, 1999, Robustness of Simple Policy Rules Under Model Uncertainty, in Taylor 1999.

McCallum, B. T., 1999, Issues in the design of monetary policy rules, forthcoming in, J. B. Taylor and M. Woodford (eds.), Handbook of Macroeconomics.

McCallum, B. T., 1999, Analysis of the monetary transmission mechanism: methodological issues, Manuscript, February.

McCallum, B. T. and E. Nelson, 1999, Performance of operational policy rules in an estimated semiclassical structural model, in Taylor 1999.

Monticelli, C. and O. Tristani, 1999, What does the single monetary policy do? A SVAR benchmark for the European Central Bank, European Central Bank, Working Paper No. 2, May.

Newey, W. K. and K. D. West, 1987, A simple, positive semi-definite, heteroskedasticity and autocorrelation consistent covariance matrix, Econometrica, 55, 703-708.

OECD, 1999, OECD economic outlook No. 65, Organisation for Economic Co-Operation and Development, June.

Orphanides, A, D. Small, D. Wilcox and V. Wieland, 1997, A quantitative exploration of the opportunistic approach to disinflation, Finance and Economics Discussion Series, 97-36, Board of Governors of the Federal Reserve System, June.

Orphanides, A. and V. Wieland, 1998, Price stability and monetary policy effectiveness when nominal interest rates are bounded at zero, Finance and Economics Discussion Series, 98-35, Board of Governors of the Federal Reserve System, June.

Peersman, G. and F. Smets, 1999, The Taylor rule: A useful monetary policy benchmark for the Euro area?, International Finance, 2(1), 85-116, April.

Reifschneider, D., R. Tetlow and J.C. Williams, 1999, Aggregate disturbances, Monetary policy, and the macroeconomy: The FRB/US perspective, Federal Reserve Bulletin, 85(1) January, 1-19.

Roberts, J., 1997, Is inflation sticky?, Journal of Monetary Economics, 39, 173-196.

Rotemberg, J. and M. Woodford, 1997, An optimization-based econometric framework for the evaluation of monetary policy, NBER Macroeconomics Annual, Cambridge: MIT Press.

Rotemberg, J. and M Woodford, 1999, Interest-rate rules in an estimated sticky price model, in Taylor, 1999.

Rudebusch, G. and L. E.O. Svensson, 1999, Policy rules for inflation targeting, in Taylor 1999.

Sack, B. and V. Wieland, 1999, Interest rate smoothing and optimal monetary policy: A review of recent empirical evidence, forthcoming in, Journal of Economics and Business.

Sbordone, A., 1998, Price and unit labor costs: A new test of price stickiness, Manuscript, Rutgers University.

Smith, A. A., 1993, Estimating nonlinear time-series models using simulated vector autoregressions, Journal of Applied Econometrics, 8, Special Issue: Econometric Inference Using Simulation Techniques, S63-S84.

Svensson, L. E. O., 1999, Inflation targeting as a monetary policy rule, Journal of Monetary Economics, Vol. 43(3), 607-654.

Toda, H. Y. and T. Yamamoto, 1995, Statistical inference in vector autoregressions with possibly integrated processes, Journal of Econometrics, 66, 225-250.

Taylor, J. B., 1979, Estimation and control of a macroeconomic model with rational expectations, Econometrica, 47(5), 1267-86.

Taylor, J. B., 1980, Aggregate dynamics and staggered contracts, Journal of Political Economy, 88, 1-24.

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

Taylor, J. B., 1993b, Discretion versus policy rules in practice, Carnegie-Rochester Conference Series on Public Policy, 39, December, 195-214.

Taylor, J. B., (ed.), 1999, Monetary policy rules, Chicago: NBER and University of Chicago Press.

White, H., 1994, Estimation, inference and specification analysis, Cambridge: Cambridge University Press.

Wieland, V., 1998, Monetary policy under uncertainty about the natural unemployment rate, Finance and Economics Discussion Series, 98-22, Board of Governors of the Federal Reserve System.

Williams, J. C., 1999, Simple rules for monetary policy, Finance and Economics Discussion Series, 99-12, Board of Governors of the Federal Reserve System, February.

Woodford, M., 1999, Optimal monetary policy inertia, Manuscript, January.

A Estimation by Indirect Inference

To estimate the structural parameters of the overlapping contracts specifications we apply the indirect inference methods proposed by Smith (1993), Gouriéroux, Monfort and Renault (1993) and developed further in Gouriéroux and Monfort (1995, 1996). Indirect inference is a simulation-based procedure for calibrating a structural model with the objective of finding values of its parameters such that its dynamic characteristics match the dynamic properties of the observed data as summarized by the estimated parameters of an approximating probability model.[22] The latter should fit the empirical dynamics reasonably well, but need not necessarily nest the structural model. The unconstrained VAR model, which we used as a descriptive device to explore the dynamic interaction of the inflation and output data in section 4 is a natural candidate for such an approximating probability model.

We start with a brief review of standard Maximum-Likelihood (ML) methods used by Taylor (1993a) and Fuhrer and Moore (199a) for estimating the structural models at hand in subsection A.1, and then motivate and describe the indirect estimation method in subsection A.2. We also discuss the asymptotic properties of the indirect estimator as well as a global specification test which may be used to assess whether the structural model is consistent with the data. Finally, subsection A.3 provides more detail on the implementation of the indirect estimation method.

A.1 The ML Estimator

Both, theTaylor and Fuhrer-Moore models have a stationary reduced-form vector autoregressive representation

$$z_t \quad = \quad B_1\, z_{t-1} + \cdots + B_p\, z_{t-p} + B_0\, \epsilon_t, \tag{A.1}$$

where the K-dimensional vector of endogenous variables $z_t = [\,x_t',\, y_t'\,]'$ comprises a k-dimensional vector of observable variables y_t and a $(K-k)$-dimensional vector of non-observable variables x_t such as the contract wage. The K-dimensional vector ϵ_t is serially uncorrelated with mean zero and positive semi-definite covariance matrix. The covariance matrix is assumed to be diagonal with its non-zero elements normalized to unity. The $(K \times K)$-dimensional coefficient matrices $B_i = B_i(\theta)$ $(i = 0, 1, \ldots, p)$ are non-linear functions of an m-dimensional vector of structural parameters $\theta \in \Theta$, where $\Theta \subset \mathbf{R}^m$ denotes the feasible parameter space. The maximum lag length p of the endogenous variables z_t is equal to the maximum length of the wage contracts minus one. We refer to (A.1) as our constrained VAR(p) model, although it is more general than a standard VAR model in that it contains the unobservable variables x_t.

[22] See Duffie and Singleton (1993) and Gallant and Tauchen (1996) for related approaches relying on selected sample moments or the scores of the approximating probability model respectively.

The reduced-form representation of the structural model can alternatively be characterized by a family of parametric conditional density functions

$$\mathbf{F}(\theta) \;=\; \Big\{ f(z_{-p+1},\ldots,z_0;\theta),\; \{\, f(y_t \,|\, z_{t-p},\ldots,z_{t-1};\theta)\,\}_{t=1}^{\infty} : \theta \in \Theta \Big\}$$

with an element of $\mathbf{F}(\theta)$ being denoted by $\mathbf{f}(\theta)$. The observed data $\{\, y_t \,\}_{t=-p+1}^{T}$ is then presumed to be a sample from $f(z_{-p+1},\ldots,z_0;\theta)\,\prod_{t=1}^{T} f(y_t \,|\, z_{t-p},\ldots,z_{t-1};\theta)$ for some $\theta \in \Theta$. Usually, the structural parameter vector θ may be estimated by Maximum-Likelihood (ML) methods relying on Kalman filtering techniques (see Fuhrer and Moore (1995a, 1995b)). In particular, if we condition on fixed pre-sample values $z_{-p+1},\ldots,z_0$, the conditional ML estimator $\hat{\theta}_T$ for θ is

$$\hat{\theta}_T \;=\; \arg\max_{\theta\in\Theta} \sum_{t=1}^{T} \ln f(y_t \,|\, z_{t-p},\ldots,z_{t-1};\theta).$$

Under appropriate regularity conditions the ML estimator $\hat{\theta}_T$ is consistent for the "true" structural parameter vector θ_0,

$$\operatorname*{plim}_{T\to\infty} \hat{\theta}_T \;=\; \theta_0,$$

and asymptotically normal,

$$\sqrt{T}\,(\hat{\theta}_T - \theta_0) \;\xrightarrow{d}\; \mathrm{N}[0, \Sigma_{\hat{\theta}}(\theta_0)],$$

where $\Sigma_{\hat{\theta}}(\theta_0)$ is the asymptotic covariance matrix of $\sqrt{T}\,(\hat{\theta}_T - \theta_0)$.

In practice, however, the ML estimator may be sensitive with respect to the choice of the partially unobserved initial conditions $z_{-p+1},\ldots,z_0$ and may lack robustness if the presumed data-generating process $\mathbf{f}(\theta)$ is mis-specified.

A.2 The Indirect Estimator

Instead of estimating the parameter vector θ of the structural model directly, indirect estimation starts from an approximating probability model – henceforth auxiliary model – which is capable of summarizing the dynamic characteristics of the sequence of observed data $\{\, y_t \,\}_{t=-p+1}^{T}$.

In the case at hand, the unconstrained k-dimensional VAR(p) model

$$y_t \;=\; A_1\, y_{t-1} + \cdots + A_p\, y_{t-p} + u_t \tag{A.2}$$

with u_t being serially uncorrelated with mean zero and positive-definite covariance matrix $\mathrm{E}[u_t u_t'] = \Sigma_u$ is a natural candidate for approximating the structural model, recalling that the latter has a constrained VAR(p) representation as given by (A.1).

Throughout, we assume that the unconstrained VAR(p) model (A.2) is stable, i.e.

$$\det(I_k - A_1 z - \cdots - A_p z^p) \;=\; 0 \quad \Rightarrow \quad |z| > 1,$$

where $|\cdot|$ denotes the absolute value operator.

The unconstrained VAR(p) model, in turn, can be characterized by a parametric family of conditional densities

$$\mathbf{F}(\beta) \;=\; \Big\{ \tilde{f}(y_{1-p}, \ldots, y_0; \beta), \; \{ \tilde{f}(y_t \,|\, y_{t-p}, \ldots, y_{t-1}; \beta) \}_{t=1}^{\infty} : \beta \in B \Big\}$$

with an element of $\mathbf{F}(\beta)$ being denoted by $\mathbf{f}(\beta)$ and where

$$\beta \;=\; \big[\operatorname{vec}(A_1, \ldots, A_p)', \operatorname{vech}(\Sigma_u)' \big]',$$

is the n-dimensional parameter vector of the auxiliary VAR(p) model with $n = pk^2 + k(k+1)/2$. $B \subset \mathbf{R}^n$ denotes the feasible parameter space. The vec($\cdot$)-operator stacks the columns of a matrix in a column vector and the vech($\cdot$)-operator stacks the elements on and below the principal diagonal of a square matrix.

Given a sample $\{ y_t \}_{t=1}^{T}$ with fixed pre-sample values $y_{-p+1}, \ldots, y_0$, its dynamic characteristics can be summarized by computing the conditional Quasi-Maximum-Likelihood (QML) estimator $\hat{\beta}_T$ for β,[23]

$$\hat{\beta}_T \;=\; \arg\max_{\beta \in B} \sum_{t=1}^{T} \ln \tilde{f}(y_t \,|\, y_{t-p}, \ldots, y_{t-1}; \beta).$$

Under general regularity conditions the QML estimator $\hat{\beta}_T$ is consistent for the "true" reduced-form parameter vector β_0,

$$\operatorname*{plim}_{T \to \infty} \hat{\beta}_T \;=\; \beta_0,$$

and asymptotically normal,

$$\sqrt{T}\,(\hat{\beta}_T - \beta_0) \;\xrightarrow{d}\; \mathrm{N}[0, \Sigma_{\hat{\beta}}(\beta_0)],$$

[23] As layed out in Section 4, we assume that the available data $\{ Y_t \}_{t=-p+1}^{T}$ are a sample of the 2-dimensional vector of variables $Y = [\,\Pi,\, Q\,]'$ being generated by the linear model

$$Y_t \;=\; \alpha_0 + \alpha_1\, t + y_t$$

with $\{ y_t \}$ following a stable VAR(p) process as represented by (A.2) above. This type of model has been advocated by Toda and Yamamoto (1995) for conducting statistical inference in vector autoregressions with possibly integrated variables.

Substituting (A.2), it becomes obvious that $\{ Y_t \}$ is assumed to follow a VAR(p) process around a deterministic linear trend,

$$Y_t - \alpha_0 - \alpha_1\, t \;=\; A_1\,(Y_{t-1} - \alpha_0 - \alpha_1\,(t-1)) + \cdots + A_p\,(Y_{t-p} - \alpha_0 - \alpha_1\,(t-p)) + u_t.$$

Since we are merely interested in the parameters of the VAR(p) model, i.e. $A_1, \ldots, A_p$ and Σ_u, we proceed in two steps. First, we detrend the data by a simple projection technique to obtain the sample $\{ y_t \}$. Second, using this sample, we compute the QML estimates of the parameters of interest.

where $\Sigma_{\hat{\beta}}(\beta_0) = \mathcal{H}(\beta_0)^{-1}\mathcal{I}(\beta_0)\,\mathcal{H}(\beta_0)^{-1}$ is the asymptotic covariance matrix of $\sqrt{T}\,(\hat{\beta}_T - \beta_0)$. $\mathcal{I}(\beta_0)$ denotes the asymptotic information matrix and $\mathcal{H}(\beta_0)$ is the asymptotic expected Hessian of the appropriately normalized quasi-log-likelihood function evaluated at β_0.[24]

If the auxiliary model $\mathbf{F}(\beta)$ approximates the structural model $\mathbf{F}(\theta)$ sufficiently well, it makes sense to estimate the vector of structural parameters θ indirectly by minimizing the "distance" between the structural model $\mathbf{F}(\theta)$ and the auxiliary model fitted to the data, i.e. $\mathbf{f}(\hat{\beta}_T) \in \mathbf{F}(\beta)$. Thus, the indirect estimator brings the information in $\hat{\beta}_T$ to bear on the task of estimating the structural parameter vector θ.

To make this approach operational Gouriéroux, Monfort and Renault (1993) and Gouriéroux and Monfort (1995, 1996) start from the well-known Kullback-Leibler information criterion (KLIC),

$$\mathrm{KLIC}(\beta,\theta) \quad = \quad \mathrm{E}_\theta\left[\ln\left(\frac{f(y_t|z_{t-p},\ldots,z_{t-1};\theta)}{\tilde{f}(y_t|y_{t-p},\ldots,y_{t-1};\beta)}\right)\right],$$

and introduce the so-called binding function $\mathrm{b}\colon \Theta \to B$,

$$\begin{aligned}\mathrm{b}(\theta) &= \arg\min_{\beta\in B}\ \mathrm{KLIC}(\beta,\theta)\\ &= \arg\max_{\beta\in B}\ \mathrm{E}_\theta\Big[\ \ln\tilde{f}(y_t|y_{t-p},\ldots,y_{t-1};\beta)\Big],\end{aligned}$$

which, for a given $\theta \in \Theta$, identifies the density $\mathbf{f}(\beta) \in \mathbf{F}(\beta)$ minimizing the distance between the particular density $\mathbf{f}(\theta) \in \mathbf{F}(\theta)$ and the elements of the family of densities $\mathbf{F}(\beta)$.

For a given $\beta \in B$, the expectation of the logarithm of the conditional density $\ln\tilde{f}(y_t|y_{t-p},\ldots,y_{t-1};\beta)$ has to be determined with respect to the conditional density $f(y_t|z_{t-p},\ldots,z_{t-1};\theta)$ from $\mathbf{f}(\theta) \in \mathbf{F}(\theta)$. While this expectation cannot be obtained analytically in general, it may easily be approximated for given parameters θ and β by the simulated sample moment

$$\frac{1}{S}\sum_{s=1}^{S}\ln\tilde{f}(y_s(\theta)|y_{s-p}(\theta),\ldots,y_{s-1}(\theta);\beta),$$

where the simulated sample $\{\,y_s(\theta)\,\}_{s=-p+1}^{S}$ is generated by the structural model characterized by $\mathbf{f}(\theta) \in \mathbf{F}(\theta)$.

Obviously, the approximation of the binding function $\mathrm{b}(\theta)$ coincides with the QML estimator $\hat{\beta}_S(\theta)$ of the auxiliary model using the simulated sample,

$$\hat{\beta}_S(\theta) \quad = \quad \arg\max_{\beta\in B}\ \frac{1}{S}\sum_{s=1}^{S}\ln\tilde{f}(y_s(\theta)|y_{s-p}(\theta),\ldots,y_{s-1}(\theta);\beta),$$

where, without loss of generality, the simulated conditional quasi-log-likelihood function has been normalized by $1/S$.

[24] See White (1994) for a thorough treatment of QML theory and covariance estimation.

Subsequently, a simulation-based indirect estimator $\hat{\theta}_{S,T}$ for θ is obtained by minimizing a criterion function $Q_{S,T} : B \times \Theta \to \mathbf{R}_+$ which is defined as a quadratic form measuring the distance between the empirical QML estimator $\hat{\beta}_T$ and the simulated QML estimator $\hat{\beta}_S(\theta)$:

$$\hat{\theta}_{S,T} \quad = \quad \arg\min_{\theta \in \Theta} Q_{S,T}(\hat{\beta}_T, \theta)$$

with

$$Q_{S,T}(\hat{\beta}_T, \theta) \quad = \quad \left(\hat{\beta}_T - \hat{\beta}_S(\theta) \right)' W_T(\hat{\beta}_T) \left(\hat{\beta}_T - \hat{\beta}_S(\theta) \right),$$

where $W_T(\hat{\beta}_T)$ is a positive definite $(n \times n)$-dimensional weighting matrix which possibly depends on the QML estimate $\hat{\beta}_T$. The two subscripts S and T indicate that the particular object depends on both the empirical sample $\{ y_t \}_{t=-p+1}^{T}$ and the simulated sample $\{ y_s(\theta) \}_{s=-p+1}^{S}$.

Under appropriate assumptions, the following results can be established.

Proposition: (Asymptotic properties of the indirect estimator)

i. The indirect estimator $\hat{\theta}_{S,T}$ is consistent for θ_0,

$$\operatorname*{plim}_{T \to \infty} \hat{\theta}_{S,T} \quad = \quad \theta_0,$$

and asymptotically normal,

$$\sqrt{T}\,(\hat{\theta}_{S,T} - \theta_0) \quad \xrightarrow{d} \quad \mathrm{N}[0, (1 + c^{-1})(\mathcal{B}'\, W\, \mathcal{B})^{-1} \mathcal{B}'\, W\, \Sigma_{\hat{\beta}}\, W\, \mathcal{B}\, (\mathcal{B}'\, W\, \mathcal{B})^{-1}],$$

where $S = cT$ with $c \in \mathbf{N}$, $\mathcal{B} = \mathcal{B}(\theta_0) = (\partial \,/\, \partial\theta')\, \mathrm{b}(\theta_0)$, $\Sigma_{\hat{\beta}} = \Sigma_{\hat{\beta}}(\beta_0) = \lim_{T\to\infty} \mathrm{Var}[\sqrt{T}(\hat{\beta}_T - \beta_0)]$ and $W = W(\beta_0) = \mathrm{plim}_{T\to\infty} W_T(\hat{\beta}_T)$ with $\beta_0 = \mathrm{plim}_{T\to\infty} \hat{\beta}_T$.

ii. The asymptotically efficient indirect estimator $\hat{\theta}^*_{S,T}$ is obtained by using a consistent estimate $\left(\hat{\Sigma}_{\hat{\beta},T} \right)^{-1}$ of the optimal asymptotic weighting matrix $W^* = \Sigma_{\hat{\beta}}^{-1}$ and is asymptotically normal,

$$\sqrt{T}\,(\hat{\theta}^*_{S,T} - \theta_0) \xrightarrow{d} \mathrm{N}[0, (1 + c^{-1})(\mathcal{B}'\, \Sigma_{\hat{\beta}}^{-1} \mathcal{B})^{-1}].$$

iii. If the auxiliary model $\mathbf{F}(\beta)$ nests the structural model $\mathbf{f}(\theta_0) \in \mathbf{F}(\theta)$, then the asymptotically efficient indirect estimator $\hat{\theta}^*_{S,T}$ is as efficient as the ML estimator $\hat{\theta}_T$.

iv. Under the null hypothesis that the structural model $\mathbf{f}(\theta_0) \in \mathbf{F}(\theta)$ is correctly specified, the statistic

$$\mathcal{Z}_{S,T}(\hat{\beta}_T, \hat{\theta}^*_{S,T}) \quad = \quad \frac{S\,T}{S + T}\, Q^*_{S,T}(\hat{\beta}_T, \hat{\theta}^*_{S,T})$$

with

$$Q^*_{S,T}(\hat{\beta}_T, \hat{\theta}^*_{S,T}) \;=\; \min_{\theta \in \Theta} \left(\hat{\beta}_T - \hat{\beta}_S(\theta)\right)' \left(\hat{\Sigma}_{\hat{\beta},T}\right)^{-1} \left(\hat{\beta}_T - \hat{\beta}_S(\theta)\right),$$

is asymptotically χ^2*-distributed with* $n - m$ *degrees of freedom.*

Proof: Parts *i.*, *ii.* and *iv.* of the proposition are alternatively established in Smith (1993), Gouriéroux, Monfort and Renault (1993) and Gouriéroux and Monfort (1996), chapter 4.5. Part *iii.* follows from Gallant and Tauchen (1996), pp. 665-666.

Among the assumptions underlying this proposition is the identification condition that the equation $\mathrm{b}(\theta) = \beta_0$ has a unique root at $\theta_0 \in \Theta$. This condition resembles the standard identification condition in a non-linear equation system. The necessary (order and rank) condition for identification is $n \geq m$. If $n > m$ holds, $n - m$ overidentifying restrictions are imposed when estimating θ. For the asymptotically efficient indirect estimator $\hat{\theta}^*_{S,T}$ these overidentifying restrictions can be tested by means of the statistic $\mathcal{Z}_{S,T}(\hat{\beta}_T, \hat{\theta}^*_{S,T})$. Its probability value $P(\mathcal{Z} > z)$ gives the probability that the re-scaled minimized criterion function takes a value larger than that obtained in the estimation exercise. Hereby, a probabilistic assessment of the consistency of the structural model and the data is obtained.

A.3 Implementation

To estimate the parameters of the different wage contracting specifications, it is necessary to close the model in a way that makes it possible to solve for expected future output gaps. Rather than specifying a complete macroeconomic model for forecasting future output gaps that enter the contract wage equations, we instead use the reduced-form output gap equation from the unconstrained VAR(p) model fitted to the data. Thus, we confine ourselves to the parameters of the inflation equation in conducting indirect inference. This approach is very much in the spirit of the limited-information ML procedure applied by Fuhrer and Moore (1995a). They use the output gap and interest rate equations from a three-dimensional VAR model to generate the output gap forecasts in the relative real wage contracting equation. Similarly, Taylor (1993a) used such a reduced-form output gap equation in applying limited-information ML methods for estimating the structural parameters of the nominal wage contracting equation.

Let $\mathcal{S}$ denote the ($(pk+1) \times (pk^2 + k(k+1)/2)$)-dimensional $(0,1)$ selection matrix which picks the elements of $\beta \in B$ corresponding to the inflation equation, then our asymptotically efficient indirect estimator $\hat{\theta}^*_{S,T}$ for $\theta \in \Theta$ is obtained by minimizing the criterion function

$$Q^*_{S,T}(\hat{\beta}_T, \theta) \;=\; \left(\hat{\beta}_T - \hat{\beta}_S(\theta)\right)' \mathcal{S}' \left(\mathcal{S}\,\hat{\Sigma}_{\hat{\beta},T}\,\mathcal{S}'\right)^{-1} \mathcal{S} \left(\hat{\beta}_T - \hat{\beta}_S(\theta)\right),$$

where $\hat{\beta}_T$ and $\hat{\beta}_S(\theta)$ are the empirical and the simulated QML estimates of the parameters of the unconstrained VAR(p) model and $(\hat{\Sigma}_{\hat{\beta},T})^{-1}$ is a consistent estimate of the optimal asymptotic weighting matrix $W^* = (\Sigma_{\hat{\beta}})^{-1}$.

In order to minimize the criterion function $Q^*_{S,T}(\hat{\beta}_T, \theta)$ with respect to $\theta \in \Theta$, we rely on the sequential dynamic programming algorithm provided by the MATLAB *Optimization Toolboox*.[25] This algorithm allows us to take account of the constraints imposed on the parameter space Θ, thereby guaranteeing the existence of a unique rational expectations solution. When minimizing $Q^*_{S,T}(\hat{\beta}_T, \theta)$, we repeatedly have to simulate samples of the observable variables $\{ y_s(\theta) \}_{s=-p+1}^{S}$ from our structural model in order to compute the simulated QML estimate $\hat{\beta}_S(\theta)$. These samples are generated by drawing a normally distributed random sequence $\{ \epsilon_s \}_{s=-(\tilde{p}+p)+1}^{S}$ and recursively computing the accompanying sequences of endogenous variables $\{ z_s(\theta) \}_{s=-(\tilde{p}+p)+1}^{S}$ by using the structural model's reduced-form representation (A.1) for varying parameter vectors $\theta \in \Theta$. The recursions may start from arbitrary initial values $\bar{z}$, but a sufficiently large number of simulated values, say $\tilde{p} = 100 - p$, should be discarded in order to guarantee that the effect of the initial values die out. The sequences of the observable variables $\{ y_s(\theta) \}_{s=-p+1}^{S}$ are subsequently retained from the sequences $\{ z_s(\theta) \}_{s=-p+1}^{S}$. Note that in repeated simulations the employed random number generator must always start from the same random seed. Similarly, always the same initial values $\bar{z}$ must be chosen.

Reporting standard errors for the indirect estimate $\hat{\theta}^*_{S,T}$ requires estimation of the asymptotic covariance matrix $\Sigma_{\hat{\theta}^*} = (1 + c^{-1})(\mathcal{B}'\,\mathcal{S}'\,(\mathcal{S}\,\Sigma_{\hat{\beta}}\,\mathcal{S}')^{-1}\mathcal{S}\,\mathcal{B})^{-1}$. This matrix may be consistently estimated by replacing the unknown quantities by consistent estimates,

$$\hat{\Sigma}_{\hat{\theta}^*,S,T} \;=\; (1 + c^{-1}) \left(B_S(\hat{\theta}^*_{S,T})'\,\mathcal{S}' \left(\mathcal{S}\,\hat{\Sigma}_{\hat{\beta},T}\,\mathcal{S}' \right)^{-1} \mathcal{S}\, B_S(\hat{\theta}^*_{S,T}) \right)^{-1}$$

with $B_S(\theta) = (\partial/\partial\theta')\,\hat{\beta}_S(\theta)$ being computed by finite difference methods.

Our experience has been, however, that it is more convenient to estimate the asymptotic covariance matrix by the inverse of the appropriately normalized Hessian of the criterion function that is returned by the numerical algorithm,

$$\hat{\Sigma}_{\hat{\theta}^*,S,T} \;=\; (1 + c^{-1}) \left(\frac{1}{2}\,\frac{\partial Q^*_{S,T}(\hat{\beta}_T, \hat{\theta}^*_{S,T})}{\partial\theta\,\partial\theta'} \right)^{-1}.$$

[25] See Branch and Grace (1996) for technical details.

Table 1: Estimates of the Unconstrained VAR Model for the Euro Area

A_1		A_2		A_3		$\Sigma_u \times 10^4$	
A. VAR(2) :							
0.4879	0.3890	0.0989	-0.2190			0.9871	
(0.0963)	(0.1709)	(0.0899)	(0.1688)			(0.1518)	
0.0481	1.1236	-0.2159	-0.1605			-0.0686	0.2736
(0.0571)	(0.0928)	(0.0366)	(0.1094)			(0.0528)	(0.0584)
B. VAR(3) :							
0.4763	0.4038	0.0995	-0.2272	0.0181	-0.0105	0.9826	
(0.0968)	(0.1661)	(0.1035)	(0.2644)	(0.1043)	(0.1774)	(0.1556)	
0.0430	1.0758	-0.1804	-0.0450	-0.0426	-0.0740	-0.0701	0.2728
(0.0561)	(0.1041)	(0.0502)	(0.1345)	(0.0533)	(0.0767)	(0.0527)	(0.0619)

Note: Estimates of the asymptotic standard errors in parentheses with the asymptotic information matrix being estimated by the Newey-West (1987) estimator with the lag truncation parameter set equal to 3.

Table 2: Estimates of the Unconstrained VAR(3) Model for France, Germany and Italy

A_1		A_2		A_3		$\Sigma_u \times 10^4$	
A. France:							
0.4898	0.4551	-0.0645	-0.4035	0.1262	0.0506	3.1883	
(0.1201)	(0.3697)	(0.1465)	(0.5265)	(0.0990)	(0.3305)	(0.6399)	
0.0069	1.1412	-0.0804	-0.0661	0.0291	-0.1499	-0.2390	0.3151
(0.0274)	(0.1135)	(0.0384)	(0.1496)	(0.0415)	(0.0889)	(0.1020)	(0.0524)
B. Germany:							
0.0334	0.4474	0.2035	-0.0508	0.1402	-0.1270	3.5367	
(0.0873)	(0.1867)	(0.0990)	(0.2240)	(0.0899)	(0.2125)	(0.4390)	
-0.0190	0.7480	-0.1061	0.1380	-0.0197	0.0692	-0.2614	1.1826
(0.0658)	(0.0812)	(0.0561)	(0.0891)	(0.0622)	(0.0901)	(0.2357)	(0.1623)
C. Italy:							
0.7137	0.5620	-0.0715	-0.5580	0.0074	0.0502	4.4426	
(0.1186)	(0.3391)	(0.2425)	(0.7027)	(0.0986)	(0.4881)	(1.1728)	
0.0005	1.3220	-0.0215	-0.3212	-0.0711	-0.0362	0.2931	0.4077
(0.0313)	(0.1312)	(0.0367)	(0.1855)	(0.0273)	(0.0877)	(0.1389)	(0.0848)

Note: Estimates of the asymptotic standard errors in parentheses with the asymptotic information matrix being estimated by the Newey-West (1987) estimator with the lag truncation parameter set equal to 3.

Table 3: Estimates of the Staggered Contracts Models for the Euro Area

	Relative Real Wage Contracts			Nominal Wage Contracts (NW)
	RW	RW-C	RW-S	
A. VAR(2): [a]				
s	.0658	.1344	0	0
	(.0283)	(.0330)	(—)[b]	(—)
γ	.0016	.0026	.0126	.0070
	(.0000)	(.0006)	(.0033)	(.0025)
σ_{ϵ_x}	.0002	.0009	.0018	.0027
	(.0000)	(.0002)	(.0001)	(.0001)
$P(\mathcal{Z} > z)$ [c]	.3265 [2]	.0510 [2]	.0197 [2]	.5743 [2]
A. VAR(3):				
s	.1276	.1372	.0742	.0456
	(.0401)	(.0129)	(.0245)	(.0465)
γ	.0022	.0046	.0212	.0115
	(.0011)	(.0008)	(.0048)	(.0053)
σ_{ϵ_x}	.0003	.0012	.0024	.0038
	(.0001)	(.0002)	(.0003)	(.0005)
$P(\mathcal{Z} > z)$	.7993 [4]	.3326 [4]	.2602 [4]	.3186 [4]

Notes: [a] Estimated standard errors in parantheses. [b] Estimate on the boundary of the parameter space. [c] Marginal probability value of the test of overidentifying restrictions. Number of overidentifying restrictions in brackets.

Table 4: Estimates of the Staggered Contracts Models for France, Germany and Italy

VAR(3)	Relative Real Wage Contracts RW	RW-C	RW-S	Nominal Wage Contracts (NW)
A. France: [a]				
s	.1085 (.0500)	0 (—) [b]	.0564 (.0230)	.1189 (.0370)
γ	.0036 (.0020)	.0108 (.0000)	.0296 (.0066)	.0041 (.0041)
σ_{ϵ_x}	.0004 (.0001)	.0052 (.0000)	.0046 (.0005)	.0048 (.0010)
$P(\mathcal{Z} > z)$ [c]	.1156 [4]	.0073 [4]	.0002 [4]	.5435 [4]
B. Germany:				
s	.0487 (.0209)	.0376 (.0195)	0 (—)	.0501 (.0296)
γ	.0061 (.0017)	.0084 (.0013)	.0273 (.0064)	.0195 (.0057)
σ_{ϵ_x}	.0008 (.0001)	.0054 (.0007)	.0063 (.0003)	.0074 (.0007)
$P(\mathcal{Z} > z)$	$< 10^{-5}$ [4]	.0001 [4]	$< 10^{-7}$ [4]	.0026 [4]
C. Italy:				
s	1/6 (—)	.1244 (.0111)	.0970 (.0162)	n.c. [d]
γ	.0006 (.0003)	.0046 (.0010)	.0141 (.0043)	n.c.
σ_{ϵ_x}	.0002 (.0000)	.0023 (.0003)	.0038 (.0005)	n.c.
$P(\mathcal{Z} > z)$	.1575 [4]	.1574 [4]	.0709 [4]	

Notes: [a] Estimated standard errors in parantheses. [b] Estimate on the boundary of the parameter space. [c] Marginal probability value of the test of overidentifying restrictions. Number of overidentifying restrictions in brackets. [d] No convergence.

Table 5: Simulation Based Indirect Estimation of the Structural Parameters

	$T = 100$			$T = 200$			$T = 500$		
	BIAS[b]	STD	RMSE	BIAS	STD	RMSE	BIAS	STD	RMSE
A. $\pi^* = 0$:[a]									
s	13.3	53.2	54.8	22.4	33.9	40.7	13.2	24.9	28.2
γ	42.0	72.2	83.5	3.7	35.2	35.3	-0.2	20.4	20.4
σ_{ϵ_x}	-13.7	22.7	26.5	-15.9	15.7	22.3	-10.4	12.0	15.9
B. $\pi^* = \pi^*_{-1} - 0.001\,(1 + \epsilon_{\pi^*})$:[c]									
s	12.1	52.4	53.8	18.5	35.9	40.3	1.5	27.3	27.4
γ	41.6	66.3	78.2	5.3	37.6	38.0	1.1	20.9	20.9
σ_{ϵ_x}	-12.4	23.0	26.1	-13.4	16.5	21.3	-5.4	12.4	13.6

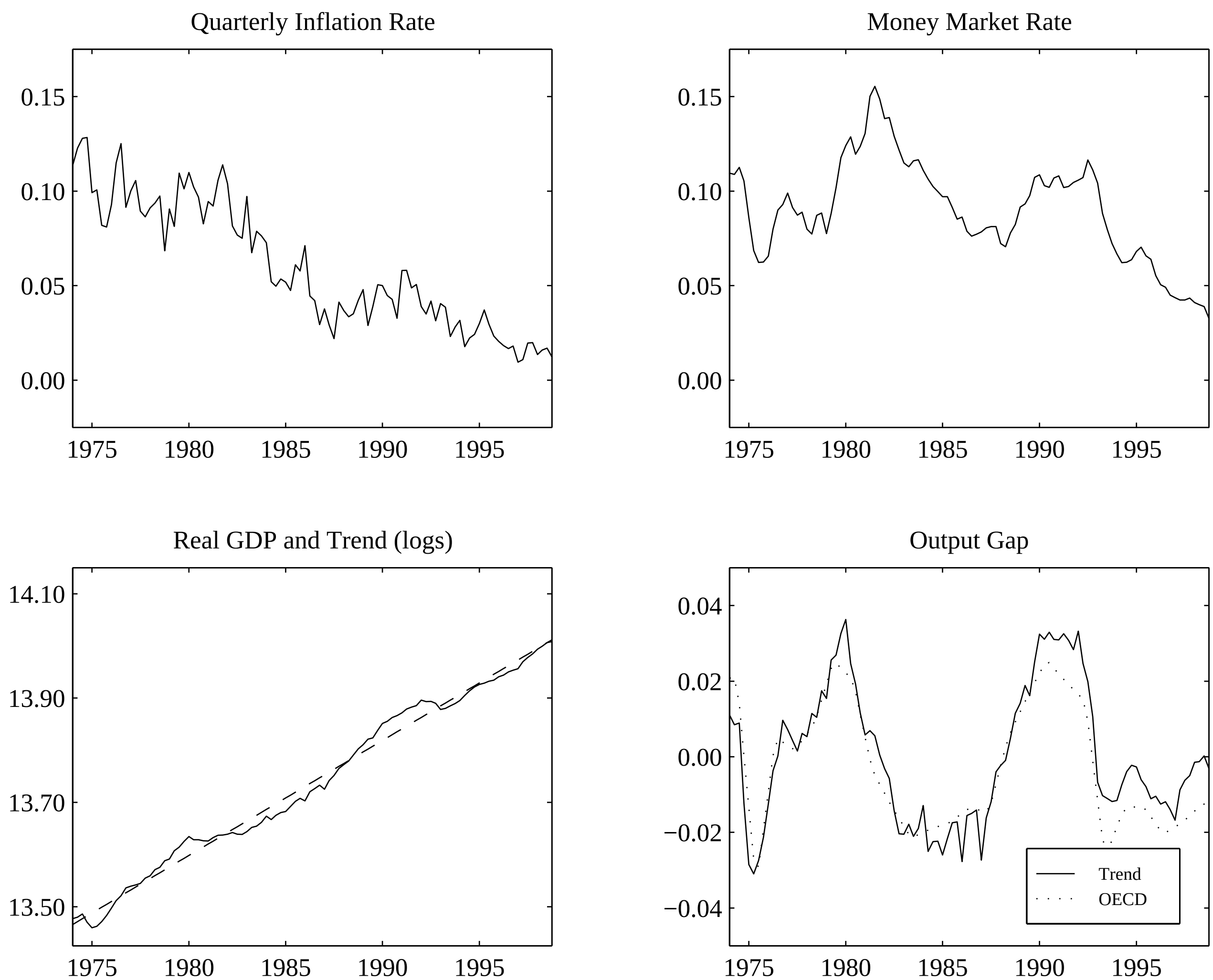
Note: [a] The structural parameters used for generating the data are $s = 0.0742$, $\gamma = 0.0212$ and $\sigma_{\epsilon_x} = 0.0024$. [b] All statistics are reported as fractions of the structural parameters (in percentage points). [c] The innovations to π^* are drawn from a standard normal distribution.

Table 6: Estimates of the IS Curve for the Euro Area, France, Germany and Italy

	δ_0	δ_1	δ_2	δ_3	$\sigma_{\epsilon_d} \times 10^4$	$P(\mathcal{J} > j)$ [b]
A. Euro Area: [a]						
A.1 area-wide rate:	0.0012	1.2347	-0.2737	-0.0364	0.3185	0.1209 [5]
	(0.0007)	(0.0916)	(0.1004)	(0.0224)		
A.2 German rate:	0.0027	1.1807	-0.2045	-0.0947	0.3176	0.2307 [5]
	(0.0012)	(0.1006)	(0.1065)	(0.0333)		
B. France:	0.0024	1.2247	-0.2708	-0.0638	0.3460	0.1977 [5]
	(0.0008)	(0.1275)	(0.1284)	(0.0234)		
C. Germany:	0.0012	0.7865	0.1395	-0.0365	1.2289	0.2518 [5]
	(0.0027)	(0.0686)	(0.0825)	(0.0874)		
D. Italy:	0.0023	1.3524	-0.3852	-0.0544	0.3913	0.4210 [5]
	(0.0009)	(0.0845)	(0.0804)	(0.0236)		

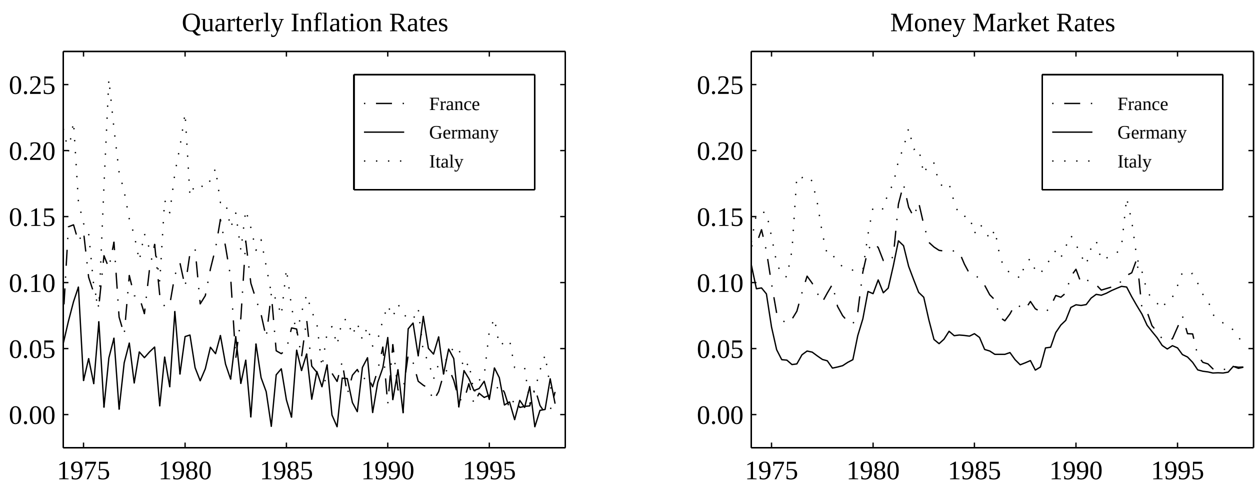
Notes: [a] GMM estimates using a vector of ones and lagged values of the output gap (q_{t-1}, q_{t-2}), the quarterly inflation rate $(\pi_{t-1}, \pi_{t-2}, \pi_{t-3})$, and the short–term nominal interest rate $(i^s_{t-1}, i^s_{t-2}, i^s_{t-3})$ as instruments. The weighting matrix is estimated by means of the Newey-West (1987) estimator with the lag truncation parameter set equal to 7. Estimated standard errors in parantheses. [b] Marginal probability value of the $\mathcal{J}$-test of overidentifying restrictions. Number of overidentifying restrictions in brackets.

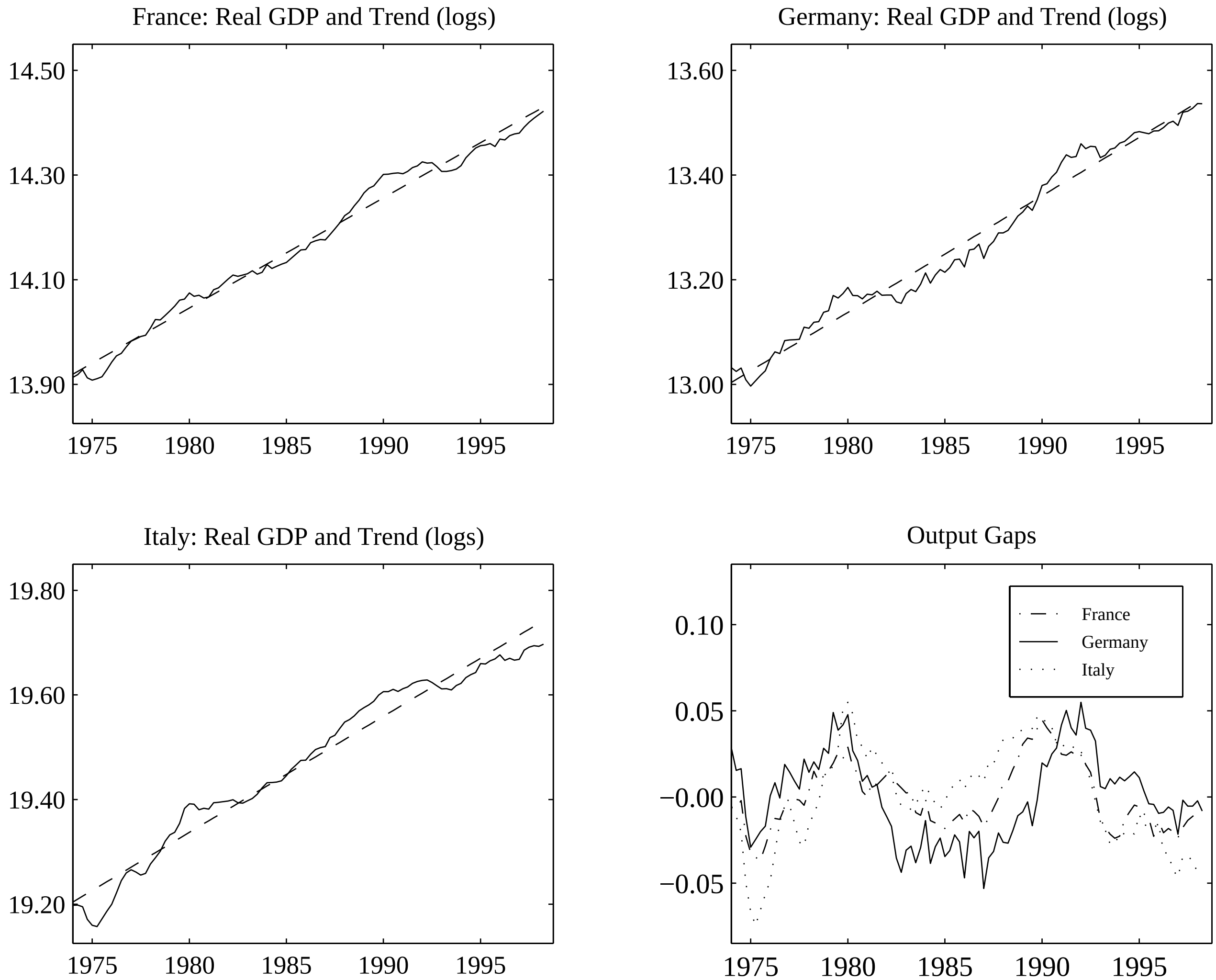
Figure 1: The Euro-Area Data



Source: ECB area-wide model database (see Fagan et al. (1999)). Aggregation over data for the member countries of the Euro area using fixed 1995 GDP weights at PPP rates. The OECD output gap is obtained by interpolating the annual figures reported in OECD (1999).

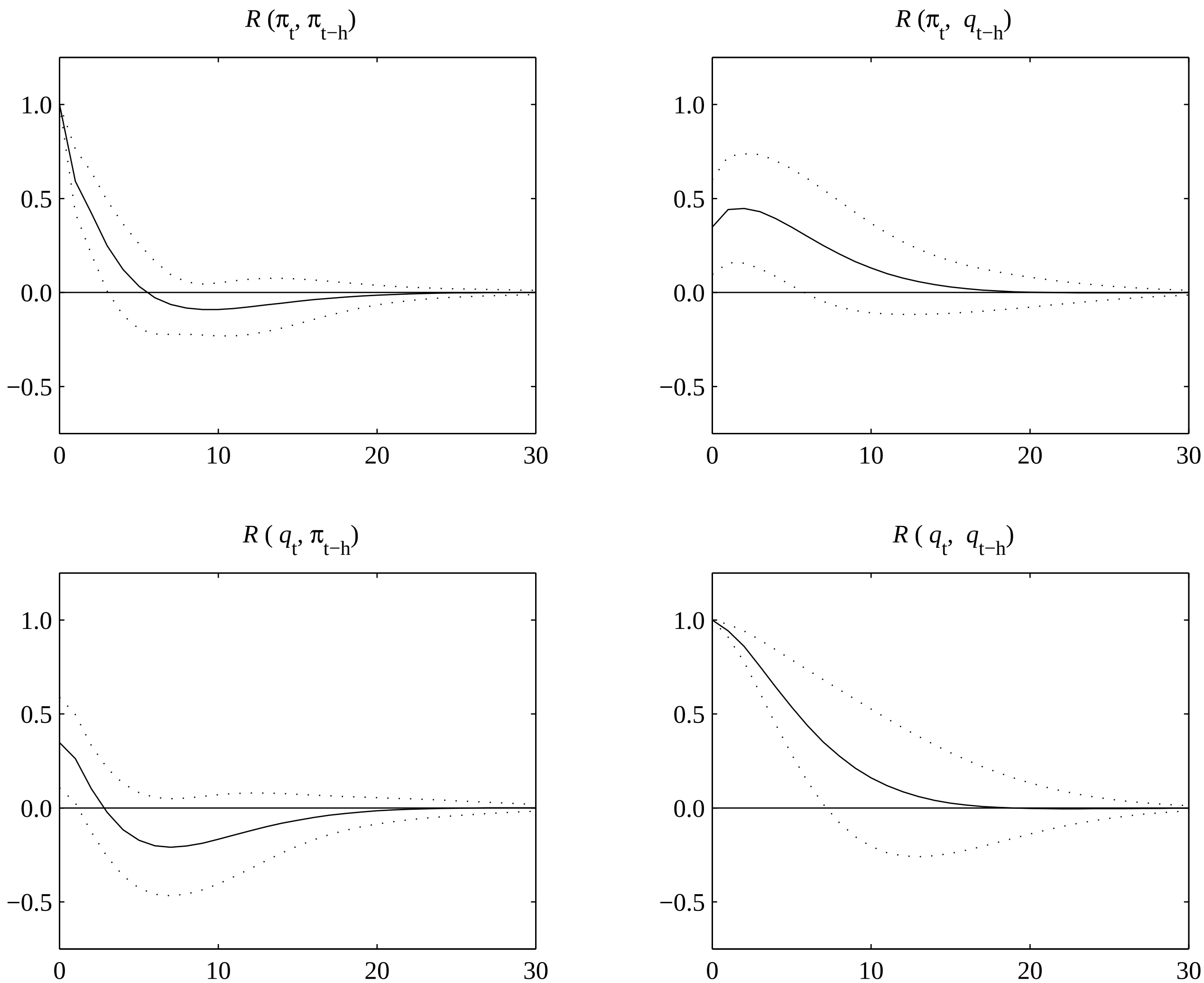
Figure 2: The Data for France, Germany and Italy





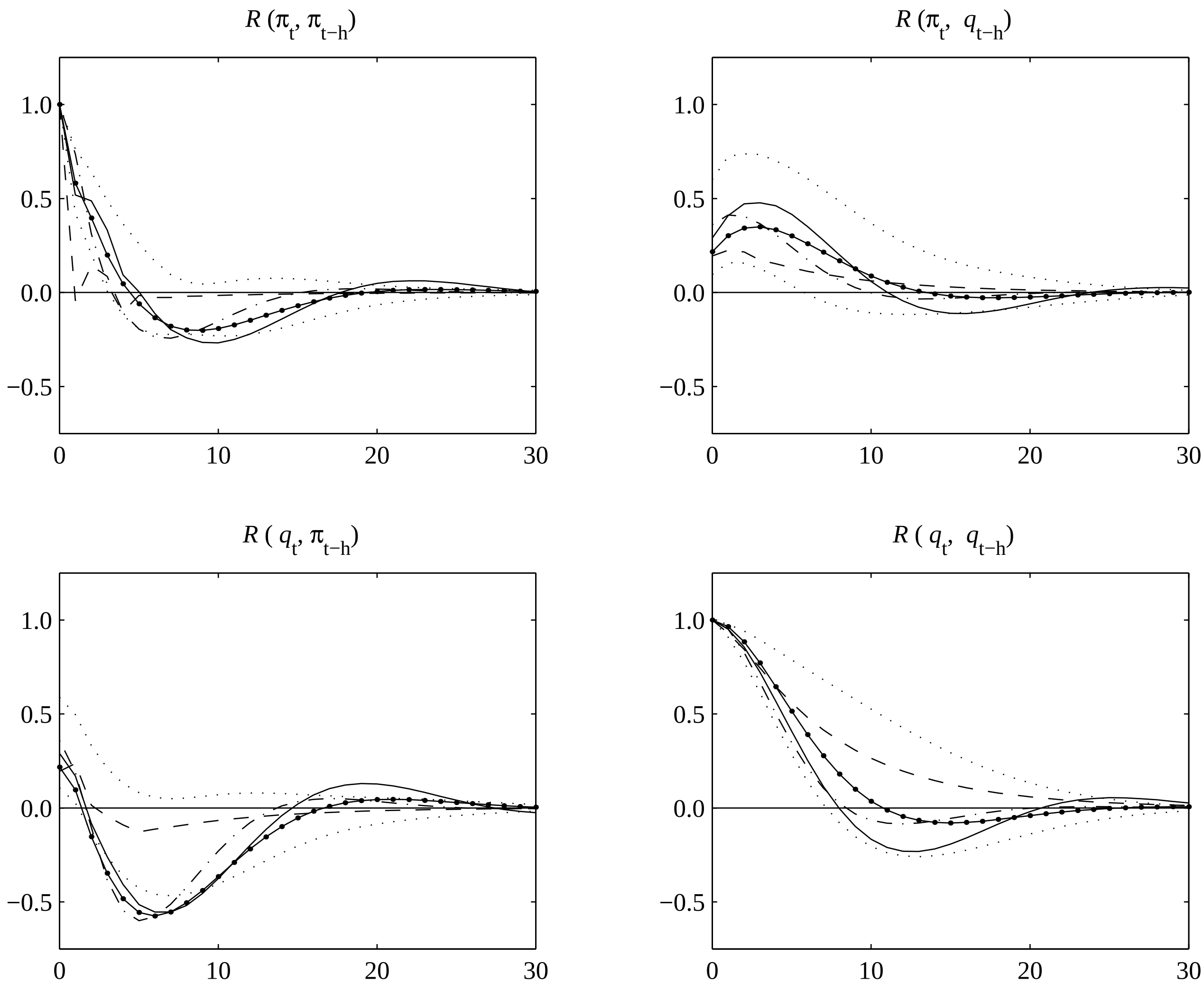
Source: ECB multi-country model database.

Figure 3: Estimated Autocorrelations of the Unconstrained VAR(3) Model for the Euro Area



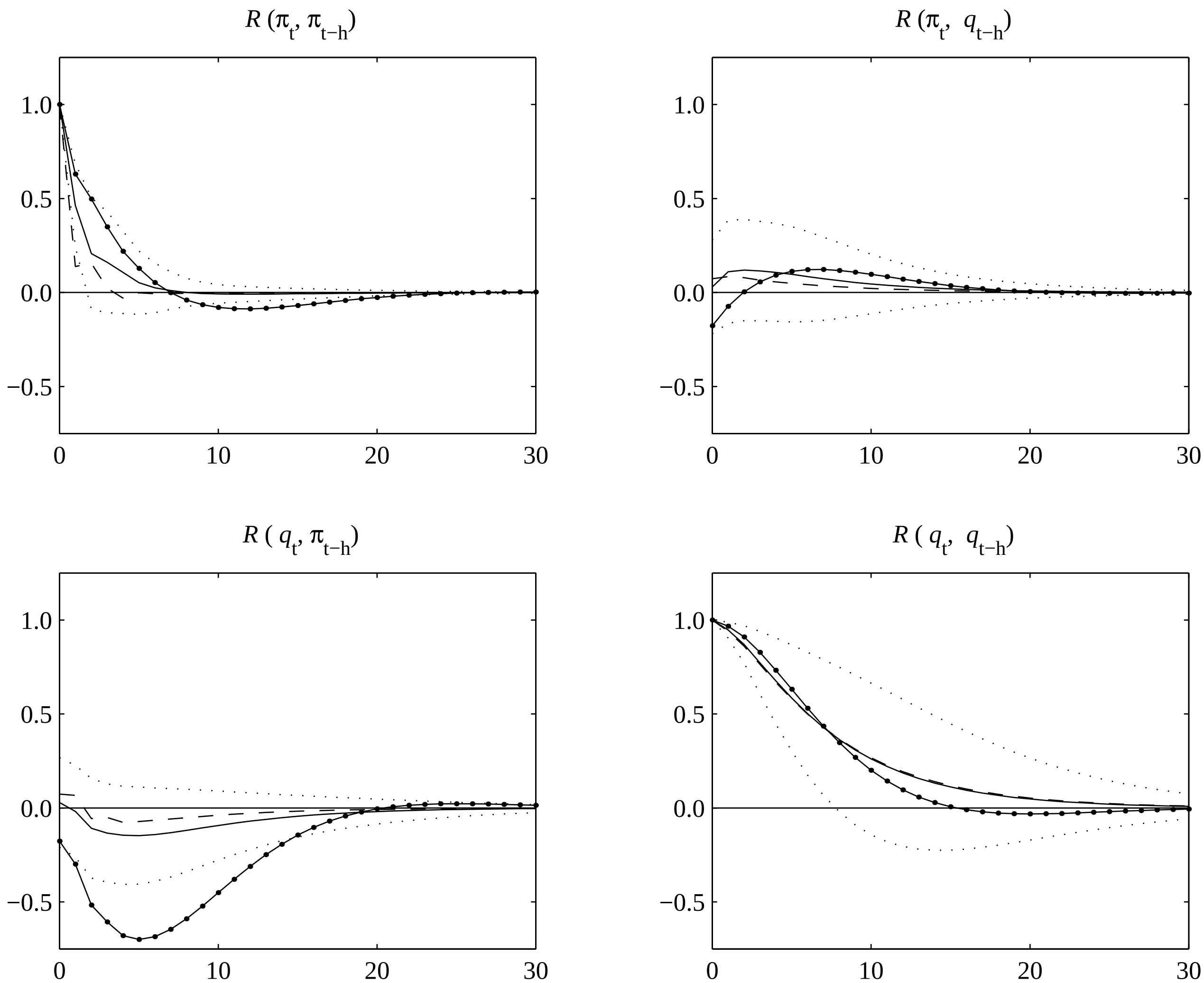
Notes: Solid line: Estimated autocorrelations. Dotted lines: Estimated autocorrelations plus/minus twice their estimated asymptotic standard errors.

Figure 4: Estimated Autocorrelations of the Constrained VAR(3) Models for the Euro Area



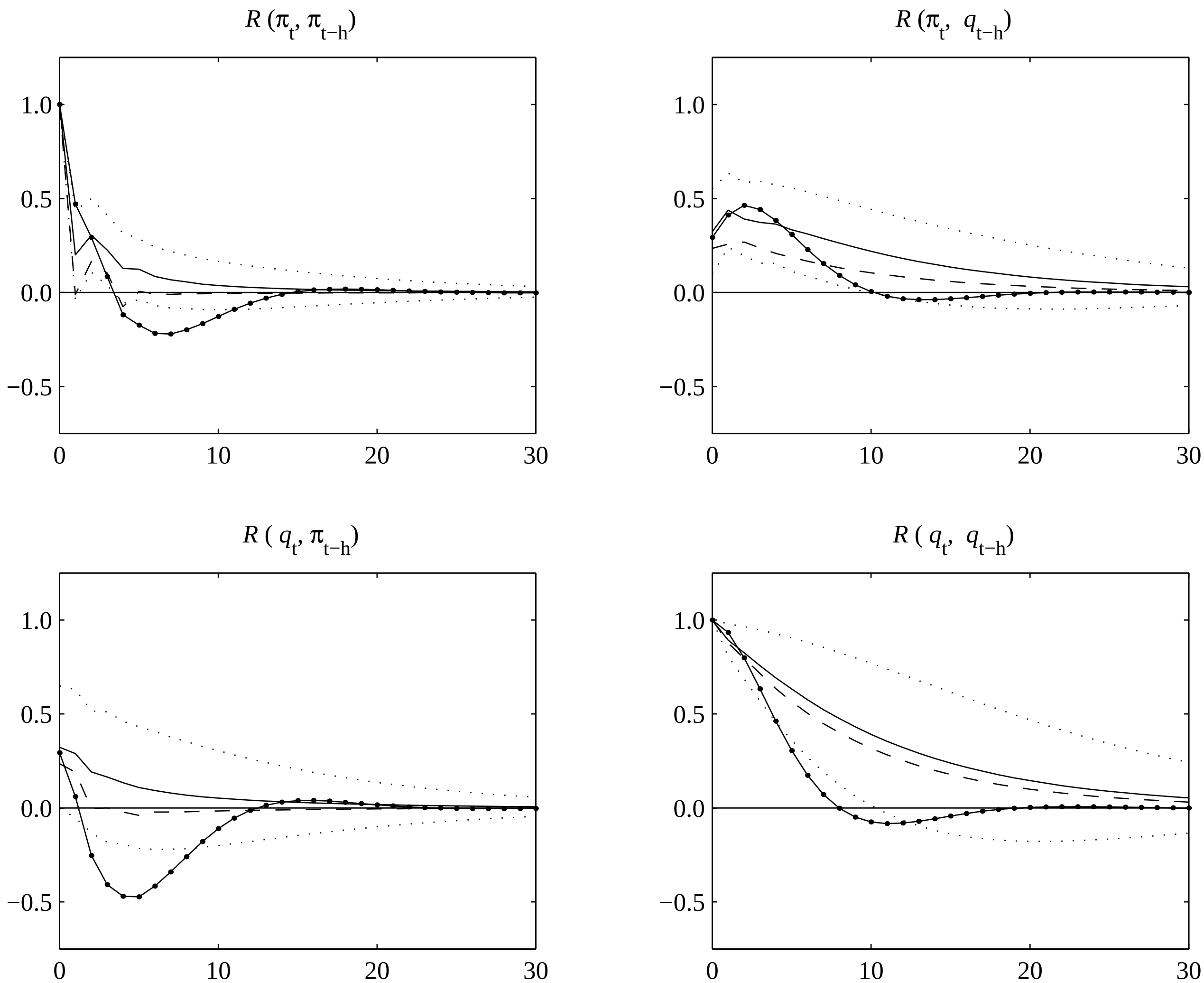
Notes: Solid line with bold dots: RW model. Dash-dotted line: RW-C model. Solid line: RW-S model. Dashed line: NW model. Dotted lines: Estimated autocorrelations of the unconstrained VAR(3) model plus/minus twice their estimated asymptotic standard errors.

Figure 5: Estimated Autocorrelations of the VAR(3) Models for France



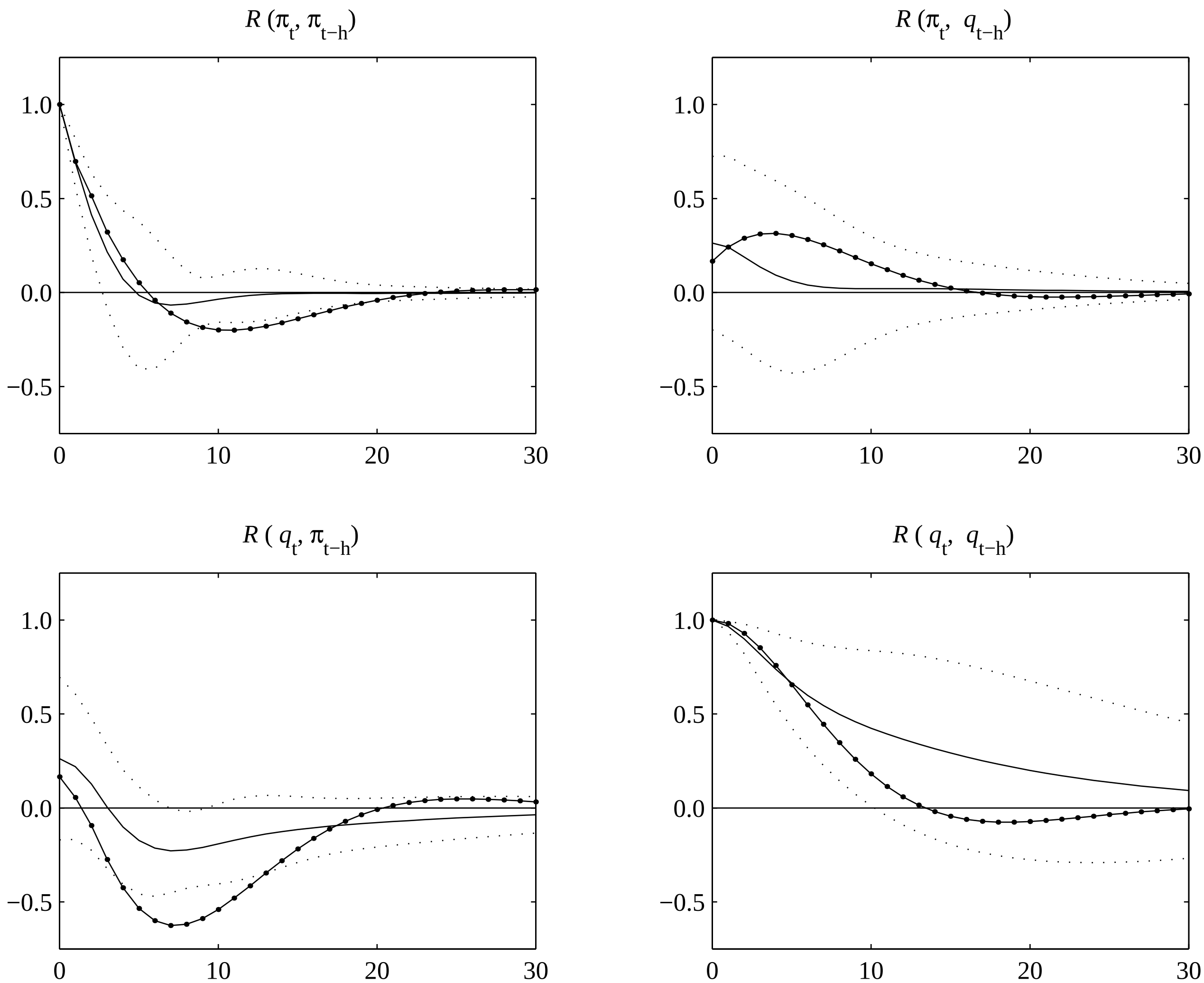
Notes: Solid line with bold dots: RW model. Dashed line: NW model. Solid line: Estimated autocorrelations of the unconstrained VAR(3) model. Dotted lines: Estimated autocorrelations of the unconstrained VAR(3) model plus/minus twice their estimated asymptotic standard errors.

Figure 6: Estimated Autocorrelations of the VAR(3) Models for Germany



Notes: Solid line with bold dots: RW model. Dashed line: NW model. Solid line: Estimated autocorrelations of the unconstrained VAR(3) model. Dotted lines: Estimated autocorrelations of the unconstrained VAR(3) model plus/minus twice their estimated asymptotic standard errors.

Figure 7: Estimated Autocorrelations of the VAR(3) Models for Italy



Notes: Solid line with bold dots: RW model. Solid line: Estimated autocorrelations of the unconstrained VAR(3) model. Dotted lines: Estimated autocorrelations of the unconstrained VAR(3) model plus/minus twice their estimated asymptotic standard errors.

Figure 8: Contract Wage Shock [1 Standard Deviation]

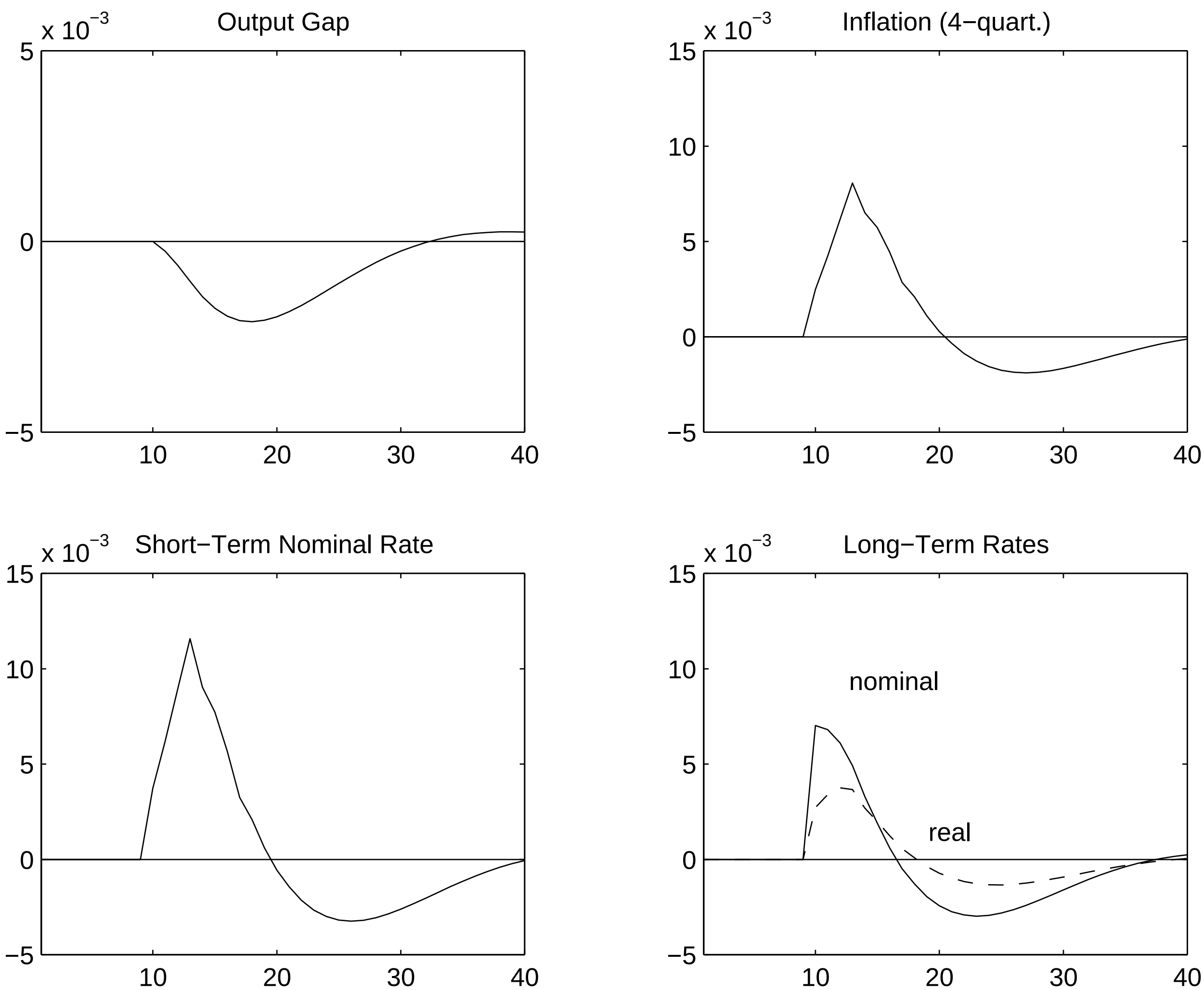


Figure 9: Demand Shock [1 Percent]

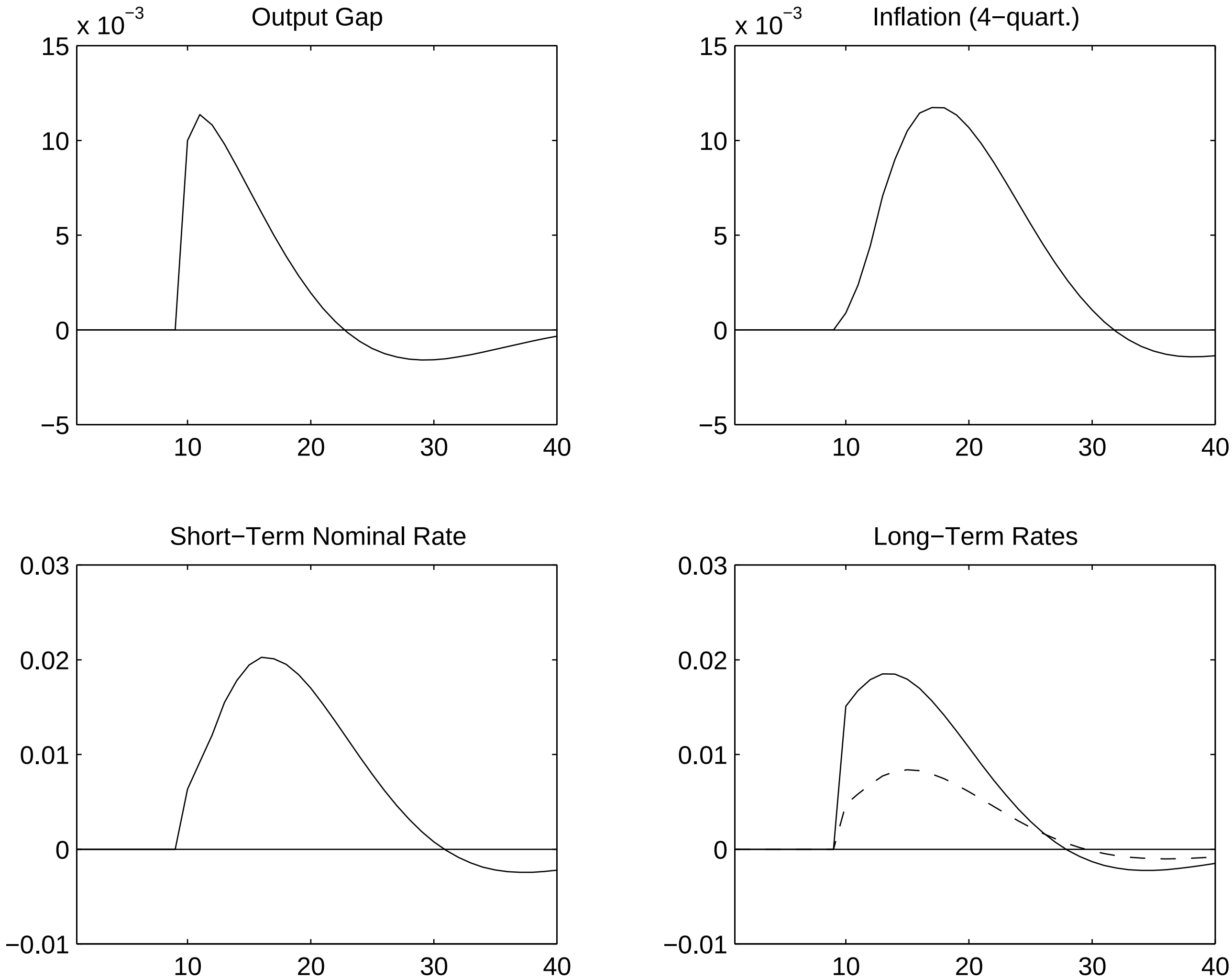


Figure 10: Disinflation [2 Percentage Points]