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INFLATION DYNAMICS AND INTERNATIONAL LINKAGES: A MODEL OF THE UNITED STATES, THE EURO AREA AND JAPAN

BY GÜNTER COENEN AND VOLKER WIELAND

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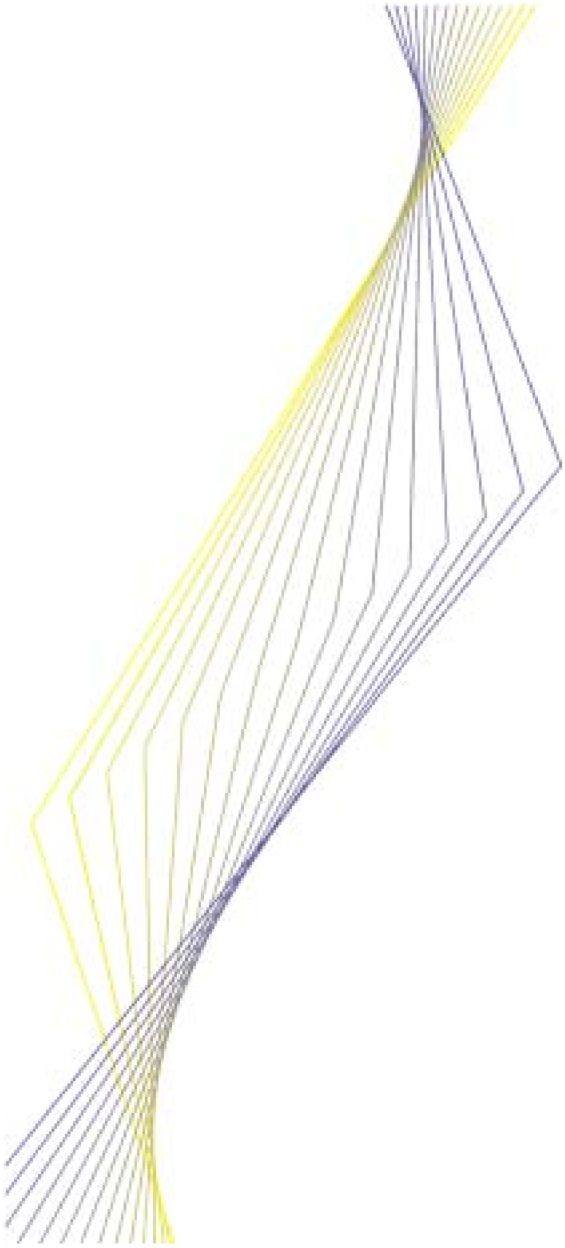

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INFLATION DYNAMICS AND INTERNATIONAL LINKAGES: A MODEL OF THE UNITED STATES, THE EURO AREA AND JAPAN[1]

BY GÜNTER COENEN[2]
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September 2002

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Abstract

In this paper we estimate a small macroeconometric model of the United States, the euro area and Japan with rational expectations and nominal rigidities due to staggered contracts. Comparing three popular contracting specifications we find that euro area and Japanese inflation dynamics are best explained by Taylor-style contracts, while Buiter-Jewitt/Fuhrer-Moore contracts perform somewhat better in fitting U.S. inflation dynamics. We are unable to fit Calvo-style contracts to inflation dynamics in any of the three economies without allowing either for ad-hoc persistence in unobservables or a significant backward-looking element. The completed model matches inflation and output dynamics in the United States, the euro area and Japan quite well. We then use it to evaluate the role of the exchange rate for monetary policy. Preliminary results, which are similar across the three economies, indicate little gain from a direct policy response to the exchange rate.

JEL Classification System: E31, E52, E58, E61

Keywords: macroeconomic modelling, nominal rigidities, inflation persistence, international linkages, monetary policy rules

Non-Technical Summary

In this paper we estimate a small three-country model with nominal rigidities and rational expectations to fit inflation and output dynamics in the three major world economies, the United States, the euro area and Japan. We then use this model to study the performance of alternative monetary policy strategies focusing on the role of the exchange rate for monetary policy.

We begin our empirical analysis by investigating whether inflation dynamics in the United States, the euro area and Japan are consistent with staggered nominal contracts and rational expectations. We consider the three specifications of nominal rigidity that have been most popular in the recent empirical literature, the staggered nominal contracts models of Calvo (1983) and Taylor (1980, 1993a) with random-duration and fixed-duration contracts respectively, as well as the relative real-wage contracting model proposed by Buiter and Jewitt (1981) and estimated by Fuhrer and Moore (1995a,b). Our findings can be summarized as follows: we are unable to estimate the inflation equation based on Calvo's specification except if we allow for ad-hoc persistence in supply shocks or some degree of adaptive expectations. Taylor's specification, which explicitly depends on lagged prices and output gaps, performs better. We are able to estimate statistically signifcant and economically meaningful parameters for all three countries consistent with rational expecations and a maximum contract length of one year. However, in case of U.S. data Taylor's specification does not pass a test of overidentifying restrictions. Fuhrer and Moore's specification obtains the best fit to U.S. data, but is dominated by Taylor's specification for Japanese and euro area data.

Concerning output dynamics we specify an open-economy aggregate demand equation that relates output to the ex-ante real interest rate and the real exchange rate as well as some additional lags of the output gap. Given estimates of the aggregate demand equation and the historical policy rule for the United States, the euro area and Japan we proceed to investigate the importance of international linkages and spillovers for the model's empirical fit as well as optimal policy design. The empirical fit of the model is quite good and the historical structural shocks implied by the multi-country model are essentially white noise. However, international linkages only seem to play a minor role. Comparing the three-country open-economy version to a version with three closed economies we find that in both cases model-generated output and inflation autocorrelations and cross-correlations are quite close to those implied by the data. International spillovers of domestic shocks turn out to be rather small when exchange rates are flexible and short-term interest rates are set according to policy rules that focus on stabilizing domestic variables.

Finally, we investigate the role of the exchange rate in the design of monetary policy rules. We compute optimized simple interest rate rules in the open-economy model and also in a variation of this model that allows for a direct effect of the real exchange rate on inflation. We find that it is largely sufficient to respond to output, inflation and lagged interest rates. Little seems to be gained from an explicit response of nominal interest rates to the exchange rate.

1 Introduction

In this paper we estimate a small three-country model with nominal rigidities and rational expectations to fit inflation and output dynamics in the three major world economies, the United States, the euro area and Japan. We then use this model to study the performance of alternative monetary policy strategies focusing on the role of the exchange rate for monetary policy. Our approach to building a macroeconometric model to be used for policy analysis is oriented on the following three principles. First, the model should fit the data under rational expectations. Thus, we avoid the assumption that policymakers can persistently fool market participants. Secondly, the model should explain the predictable dynamics of key macroeconomic variables such as output and inflation. Thus, we avoid ad-hoc specification of persistence in unobserved error processes. Thirdly, in specifying the final model equations we try to incorporate elements of models derived under optimising behavior of representative agents but we give priority to the preceding two principles.

We begin our empirical analysis by investigating whether inflation dynamics in the United States, the euro area and Japan are consistent with staggered nominal contracts and rational expectations. We consider the three specifications of nominal rigidity that have been most popular in the recent empirical literature[1], the staggered nominal contracts models of Calvo (1983) and Taylor (1980, 1993a) with random-duration and fixed-duration contracts respectively, as well as the relative real-wage contracting model proposed by Buiter and Jewitt (1981) and estimated by Fuhrer and Moore (1995a,b). Each contracting specification implies a different structural inflation equation that we try to fit to the data taking reduced-form output dynamics as given and assuming rational expectations regarding future output and inflation. In case of Calvo- and Taylor-style contracts these inflation equations have been shown to be consistent with optimising behavior of representative households and monopolistically competitive firms.[2] The inflation equation resulting from Fuhrer-Moore-

[1] See for example the recent papers by Coenen and Wieland (2000), Galí and Gertler (1999), Neiss and Nelson (2001), Roberts (1997), Sbordone (2002) or Rudd and Whelan (2002).

[2] Concerning Calvo-style contracts see for example Rotemberg and Woodford (1997), Clarida, Galí and Gertler (1999) and Sbordone (2002). Concerning Taylor-style contracts see for example Chari, Kehoe and

style contracts, however, has been criticized for lacking such a foundation in optimizing behavior.[3]

Our empirical findings can be summarized as follows: we are unable to estimate the inflation equation based on Calvo's specification except if we allow for ad-hoc persistence in supply shocks or some degree of adaptive expectations. Taylor's specification, which explicitly depends on lagged prices and output gaps, performs better. We are able to estimate statistically signifcant and economically meaningful parameters for all three countries consistent with rational expecations and a maximum contract length of one year. However, in case of U.S. data Taylor's specification does not pass a test of overidentifying restrictions. Fuhrer and Moore's specification obtains the best fit to U.S. data, but is dominated by Taylor's specification for Japanese and euro area data.[4] Thus, for the final inflation equations of our multi-country model we pick Taylor's specification for Japan and the euro area, and Fuhrer and Moore's specification for the United States.

Concerning output dynamics, we are not aware of a possible specification that would satisfy all three modelling principles. Although there is an active and rapidly growing literature on closed and open-economy models, which are consistent with optimizing behavior of representative households and firms, these models do not yet seem able to match hump-shaped output dynamics without introducing persistence in unobservables. Instead we specify an open-economy aggregate demand equation that relates output to the ex-ante real interest rate and the real exchange rate as well as some additional lags of the output gap.[5] While these lags do not yet have micro-foundations, we prefer to incorporate such predictable output dynamics explicitly in the output equation rather than assuming per-

McGrattan (2000) and King and Wolman (1999).

[3] Christiano, Eichenbaum and Evans (2001) provide such a foundation for a specification of price and wage contracts with partial indexation, which implies an inflation equation that is quite similar to Fuhrer-Moore's relative real wage contracting model.

[4] These findings confirm earlier results in Coenen and Wieland (2000) for the euro area and Fuhrer and Moore (1995) for the United States, although with our sample the Taylor specification performs better on U.S. data than in Fuhrer and Moore's earlier investigation. Recent work by Guerrieri (2002) even suggests that with a longer sample Taylor-style contracts are not rejected for U.S. data.

[5] With this approach we follow Taylor (1993a). The resulting estimated model also has many similarities to the calibrated model considered by Svensson (2000).

sistence in unobservable shock processes. Clearly, these lags will have implications for the design of monetary policy rules in either case.

Given estimates of the aggregate demand equation and the historical policy rule for the United States, the euro area and Japan we proceed to investigate the importance of international linkages and spillovers for the model's empirical fit as well as optimal policy design. The empirical fit of the model is quite good and the historical structural shocks implied by the multi-country model are essentially white noise. However, international linkages only seem to play a minor role. Comparing the three-country open-economy version to a version with three closed economies we find that in both cases model-generated output and inflation autocorrelations and cross-correlations are quite close to those implied by the data. International spillovers of domestic shocks turn out to be rather small when exchange rates are flexible and short-term interest rates are set according to policy rules that focus on stabilizing domestic variables.

Finally, we investigate the role of the exchange rate in the design of monetary policy rules. We compute optimized simple interest rate rules in the open-economy model and also in a variation of this model that allows for a direct effect of the real exchange rate on inflation. We find that it is largely sufficient to respond to output, inflation and lagged interest rates. Little seems to be gained from an explicit response of nominal interest rates to the exchange rate. Finally, we also investigate the extent of possible gains from international monetary coordination.

The paper proceeds as follows. Section 2 presents the supply side of the model, that is, the alternative staggered contracts specifications, and reports the empirical findings for the United States, the euro area and Japan. In section 3, we discuss the determination of aggregate demand, the role of monetary policy and international linkages. We also report estimation results regarding aggregate demand equations and forward-looking policy rules. Section 4 reviews the empirical fit of the complete multi-country model, while section 5 investigates the extent of international spillovers. In section 6 we analyze the role of the exchange rate for monetary policy rules, while section 7 concludes.

2 The supply side: inflation dynamics and staggered contracts

2.1 Calvo-style contracts

Calvo-style random-duration contracts have been the workhorse of the recent theoretical literature on monetary policy in models with nominal rigidities and optimizing representative households and firms. Typically, intermediate goods firms that are monopolistically competitive are assumed to set prices on a staggered basis. The duration of a given firm's price x_t is random but the probability that a firm keeps its price fixed in a given period (or gets to change its price) is constant. As shown by Rotemberg and Woodford (1997) and others a log-linearized version of the optimal price setting rule for x_t together with the definition of the aggregate price index p_t implies a log-linear relationship between inflation ($\pi_t = p_t - p_{t-1}$), expected inflation and marginal cost. Assuming output is proportional to marginal cost this relationship implies the so-called 'New Keynesian' Phillips curve,

$$\pi_t = \kappa \, \mathrm{E}_t[\, \pi_{t+1} \,] + \gamma \, q_t, \tag{1}$$

where $q_t = y_t - y^*$ denotes the gap between current output y_t and the 'natural' output level y_t^* that would occur with completely flexible prices. κ refers to the discount factor.

Recent empirical studies with U.S. and euro area data have typically rejected equation (1) and have shown that some degree of ad-hoc serial correlation in supply shocks or share ω of price setters with backward-looking 'rules-of-thumb' behavior is necessary to fit the data.[6] Thus, we will also consider two empirical extensions of the inflation equation implied by Calvo-style contracts. First an extension with purely rational expectations but serially correlated supply shocks:

$$\begin{aligned} \pi_t &= \kappa \, \mathrm{E}_t[\, \pi_{t+1} \,] + \gamma \, q_t + u_t \qquad (2) \\ u_t &= \rho \, u_{t-1} + \sigma_{\epsilon_\pi} \, \epsilon_\pi, \end{aligned}$$

[6] Some authors have argued that the New-Keynesian Phillips curve fits the data if one uses unit labor cost, which is a more direct measure of marginal cost than the output gap, in the equation (cf. Galí and Gertler (2000) and Sbordone (2002). However even in those cases allowing for ad-hoc persistence in the error process or an backward-looking element seems to be necessary to fit the data.

where ρ measures the degree of serial correlation in the supply shock u_t. The innovation ϵ_π to this supply shock is assumed to be serially uncorrelated with zero mean and unit variance and scaled by the parameter σ_{ϵ_π}. Secondly, a hybrid version with weighted forward-looking and backward-looking expectations,

$$\pi_t = (1-\omega)\,\mathrm{E}_t[\,\pi_{t+1}\,] + \omega\,\pi_{t-1} + \gamma\,q_t + \sigma_{\epsilon_\pi}\,\epsilon_\pi, \tag{3}$$

where ω refers to the share of backward-looking price-setters.

2.2 Taylor-style contracts

In contrast to Calvo's model, Taylor-style contracts are of fixed duration. The original motivation for this type of nominal rigidity was the existence of long-term nominal wage contracts. Understood as a source of nominal wage rigidity, Taylor-style fixed duration contracts imply that the aggregate wage level can be expressed as a weighted average of current and previously negotiated contract *wages* x_{t-i} $(i = 0, 1, \ldots, \eta(x))$, which are still in effect. Fuhrer and Moore (1995a) and others have treated the aggregate price and aggregate wage indices interchangeably, which is consistent with a fixed markup. In this case the aggregate price level p_t can be related directly to contract wages:

$$p_t = \sum_{i=0}^{\eta(x)} f_i\, x_{t-i}. \tag{4}$$

The weights f_i $(i = 1, \ldots, \eta(x))$ on contract wages from different periods are assumed to be non-negative, non-increasing and time-invariant and need to sum to one.[7] The parameter $\eta(x)$ corresponds to the maximum contract length. Workers negotiate long-term contracts and compare the contract wage to past contracts that are still in effect and future contracts that will be negotiated over the life of this contract. The contract wage x_t in Taylor's model is determined as follows:

$$x_t = \mathrm{E}_t\left[\sum_{i=0}^{\eta(x)} f_i\, p_{t+i} + \gamma \sum_{i=0}^{\eta(x)} f_i\, q_{t+i}\right] + \sigma_{\epsilon_x}\,\epsilon_{x,t}, \tag{5}$$

[7] As discussed in Taylor (1993a) this assumption is consistent with the existence of wage contracts of different length in constant proportions.

where $q_t = y_t - y_t^*$ again denotes the output gap. Thus, the contract wage 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 deviation of output from its potential over this period. Since the price indices p_{t+i} reflect contemporaneous and preceding contract wages, (5) implies that wage setters look at an average of nominal contract wages negotiated in the recent past and expected to be negotiated in the near future when setting the current contract wage. The sensitivity of contract wages to excess demand is measured by γ. The contract wage shock $\epsilon_{x,t}$, which is assumed to be serially uncorrelated with zero mean and unit variance, is scaled by the parameter σ_{ϵ_x}.

More recently, Chari, Kehoe and McGrattan (2000) have derived further micro-foundations for an equation such as (5). However, in their model x_t stands for contract *prices* rather than contract *wages*. They build a representative agent model with monopolistically competitive firms and impose a Taylor-contract-style constraint that firms set prices for a fixed number of periods and do so in a staggered fashion. In particular, each period, $1/\eta$ firms are assumed to choose new prices that are then fixed for η periods.[8] A log-linear approximation of a stripped-down version of their equilibrium implies a contract price equation that coincides with Taylor's contract wage equation (5). Thus, they are able to express the parameter γ as a function of deeper technology and preference parameters.

For empirical purposes it will be more convenient to rewrite equations (4) and (5) in terms of the quarterly inflation rate π_t and the real contract wage $x_t - p_t$. Thus, inflation rates can be used in estimation. The contract wages, which are unobservable, can be inferred from past output and inflation data given an assumption regarding initial conditions.[9] In comparison to the inflation equation (1) that results under Calvo-style contracts, it is important to note that *the inflation equation resulting from Taylor-style contracts will depend explicitly on lagged inflation and on lagged output gaps* if the maximum contract length exceeds two quarters.

[8] Thus, in this model the nominal rigidity occurs in intermediate goods markets, while there is labor and capital markets clearing.

[9] For a more detailed discussion see Coenen and Wieland (2000).

2.3 Fuhrer-Moore style contracts

The distinction between Taylor-style nominal wage contracts and Fuhrer-Moore's relative real wage contracts concerns the definition of the wage indices that form the basis of the intertemporal comparison underlying the determination of the current nominal contract wage. Thus, Fuhrer and Moore's specification should not be understood as a real wage rigidity, but rather as an alternative nominal rigidity.

Workers negotiating their nominal wage are assumed to compare the implied real wage with the real wages on overlapping contracts in the recent past and near future. As a result, the expected real wage under contracts signed in the current period is set with reference to the average real contract wage index expected to prevail over the current and the next three quarters:

$$x_t - p_t = \mathrm{E}_t \left[\sum_{i=0}^{\eta(x)} f_i \, v_{t+i} + \gamma \sum_{i=0}^{\eta(x)} f_i \, q_{t+i} \right] + \sigma_{\epsilon_x} \, \epsilon_{x,t}, \tag{6}$$

where $v_t = \sum_{i=0}^{\eta(x)} f_i \, (x_{t-i} - p_{t-i})$ refers to the average of real contract wages that are effective at time t.

Fuhrer and Moore (1995a) prefer this specification to Taylor-style contracts because it gives more weight to past inflation and consequently provides a better fit to the observed degree of inflation persistence in the United States. It has been criticised, however, for lacking explicit microfoundations that are available for the inflation equations resulting from Calvo- or Taylor-style contracts.

2.4 Estimation

We estimate the three different staggered contracts models in two stages. In the first stage, we fit unconstrained VAR models to output and inflation in the three economies. In the second stage we use the unconstrained VARs as auxiliary models in estimating the structural parameters of the staggered contracts specifications by indirect inference methods. These estimates are obtained assuming that market participants form rational expectations of future output and inflation. The estimation methodology is described in more detail in the appendix.

The data that we use in the first stage to estimate the unconstrained VARs comprises real GDP and the GDP deflator for the United States and Japan and area-wide averages of those same variables for the euro area.[10] In constructing output gaps we need a measure of potential output. For the United States and Japan we have investigated various alternatives such as log-linear trends with and without breaks, as well as estimates that can be recovered from output gap estimates of the OECD and the Congressional Budget Office. The results we will focus on in the following are based on the OECD's estimate, however our findings are quite robust to the alternatives we have considered. For the euro area we stick to the log-linear trends used in our earlier paper (cf. Coenen and Wieland (2000)).[11] A chart of the data used for estimation is shown in **Figure A** in the appendix.[12]

The estimation results with U.S., euro area and Japanese data are summarized in **Table 1**. We were unable to obtain statistically significant and economically meaningful estimates of the inflation equation (1) implied by Calvo-style contracts for any of the three economies. This is perhaps not so surprising given the recently documented failure of this specification to fit U.S. inflation dynamics (at least in the version which imposes proportionality of marginal cost and output). Thus, the first three rows refer to our estimates for the extended versions of Calvo's specification (that is, equation (2) and equation (3), respectively), which incorporate either some share of backward-looking price-setters ($0 < \omega < 1$) or positive serial correlation in supply shocks ($\rho > 0$).

We were not able to obtain meaningful and significant estimates for a version with purely rational expectations and serial correlation in supply shocks with U.S. data. Instead we report estimates of a hybrid version with a 50% share of price setters with adaptive

[10] The euro area data, which are averages of member country data using fixed GDP weights at PPP rates, have been obtained from the ECB area-wide model database (see Fagan et al. (2001)).

[11] The reason being that available OECD output gap estimates do not allow the construction of potential output series that would be appropriate for our quarterly euro area data.

[12] For the euro area the chart shows the de-trended inflation series. Historical euro area inflation contains a downward trend due to the gradual policy-driven convergence of inflation rates in Italy and France to German levels during the EMS. For further discussion we refer the reader to Coenen and Wieland (2000) and the sensitivity analysis in this paper.

Table 1: Estimated Contracting Specifications

Calvo	κ	ω	ρ		γ	σ_ϵ	p-value $^{(d)}$
United States $^{(a,b)}$	-	0.4797 (0.0103)	-		0.0041 (0.0011)	0.0017 0.0001	0.0068 [2]
Euro Area $^{(a,c)}$	0.99	-	0.6322 (0.0568)		0.0206 (0.0067)	0.0012 (0.0001)	0.2601 [2]
Japan $^{(a,b)}$	0.99	-	0.8863 (0.0536)		0.0071 (0.0114)	0.0007 (0.0003)	$< 10^{-5}$ [2]
Taylor	f_0	f_1	f_2	f_3	γ	σ_{ϵ_x}	p-value $^{(d)}$
United States $^{(a,b)}$	0.2535 (0.0164)	0.2535	0.2534 (0.0171)	0.2396	0.0095 (0.0056)	0.0054 0.0003	$< 10^{-15}$ [3]
Euro Area $^{(a,c)}$	0.2846 (0.0129)	0.2828 (0.0111)	0.2443 (0.0131)	0.1883	0.0158 (0.0059)	0.0042 (0.0003)	0.2658 [2]
Japan $^{(a,b)}$	0.3301 (0.0303)	0.2393 (0.0062)	0.2393	0.1912	0.0185 (0.0057)	0.0068 (0.0006)	0.0162 [3]
Fuhrer-Moore	f_0	f_1	f_2	f_3	γ	σ_{ϵ_x}	p-value $^{(d)}$
United States $^{(a,b)}$	0.6788 (0.0458)	0.2103 (0.0220)	0.0676 (0.0207)	0.0432	0.0014 (0.0008)	0.0004 (0.0001)	0.8749 [2]
Euro Area $^{(a,c)}$	0.7664 (0.0136)	0.1712 (0.0121)	0.0546 (0.0063)	0.0078	0.0014 (0.0003)	0.0002 (0.0000)	0.1644 [2]
Japan $^{(a,b)}$	0.8986 (0.0428)	0.0828 (0.0338)	0.0149 (0.0086)	0.0037	0.0001 (0.0001)	0.0001 (0.0001)	0.0027 [2]

Notes: $^{(a)}$ Simulation-based indirect estimates using a VAR(3) model of quarterly inflation and the output gap as auxiliary model. Estimated standard errors in parentheses. $^{(b)}$ Output gap measure constructed using OECD data. $^{(c)}$ Inflation in deviation from linear trend and and output in deviation from log-linear trend. $^{(d)}$ Probability value associated with the test of overidentifying restrictions. Number of overidentifying restrictions in brackets.

expectations as shown in the first two rows of **Table 1**. With euro area and Japanese data we succeeded in estimating inflation equations with purely rational expectations as long as we allow for a significant degree of serial correlation in supply shocks and restrict the discount rate κ to 0.99. The ad-hoc persistence parameter ρ is estimated to be 0.63 and 0.89 respectively.[13] Finally, we note that only the euro area estimates of the extended Calvo specification pass a test of overidentifying restrictions (see last column).

Estimation results for Taylor-style contracts as well as Fuhrer and Moore's preferred[14] relative-real-wage contracts are reported in the remaining rows of **Table 1**. The inflation process under those specifications is defined by equations (4), (5) and (6). The first 3 columns contain estimates of three of the contract weights (f_0, f_1, f_2). The fourth weight, f_3, is determined by the constraint that the weights sum to one. The fifth column contains the estimate of the contract slope parameter γ, while the sixth column reports the estimate of the scaling factor of contract wage shocks.

Starting with the results for Taylor contracts we note that the estimated sensitivity parameter γ has the appropriate signs and are statistically significant for all three economies (at the 5% level for the euro area and Japan and the 10% level for the United States). Unfortunately, the test of overidentifying restrictions rejects this specification for U.S. data and also for Japanese data.[15] With regard to Fuhrer-Moore contracts we obtain statistically significant and economically meaningful estimates for the United States and the euro area. As indicated by the p-value for the test of overidentifying restrictions Fuhrer-Moore contracts perform better in terms of matching U.S. inflation dynamics, but worse in terms of euro area inflation dynamics than Taylor contracts. Furthermore, the Fuhrer-Moore specification is rejected more strongly in the case of Japan. Our findings broadly confirm earlier investigations by Fuhrer and Moore (1995) concerning the United States[16] and Coenen and Wieland (2000) concerning the euro area, albeit with different methodology, different data sample and more flexibility in the estimation of the contract weights f_i.

Sensitivity studies. We have experimented somewhat with the choice of potential output

[13] We also obtain estimates of the hybrid version with adaptive expectations, which are not reported in the table (ω equals 0.38 for the euro area and 0.48 for Japan).

[14] For a more detailed discussion of variations of relative real wage contracts see Coenen and Wieland (2000).

[15] With regard to Japan, however, we have been able to estimate a version of Taylor contracts for an alternative output gap measure that passes this test. These sensitivity studies are discussed further below.

[16] Although in our case Taylor's specification does not fail as miserably.

measure for Japan and the United States. With regard to Japan we found that Taylor-style contracts cannot be rejected when a log-linear trend with break is used. However, we prefer to stick with the parameter estimates obtained based on OECD potential output data. Our findings for the United States are confirmed when using a log-linear trend or CBO estimates of potential output. For the euro area, we stick with the assumptions of a linear trend in inflation and a log-linear trend in output. We have subjected these assumptions to a barrage of sensitivity tests in Coenen and Wieland (2000).[17] A further difference to our earlier work and that of Fuhrer and Moore is that we relax the constraint on the contract weights. This constraint implied that the weights were determined by a single parameter s such that $f_i = .25 + (1.5 - i)\, s,\ s \in (\,0, 1/6\,]$. Estimates for the United States, the euro area and Japan with this constraint are reported in **Table A** in the appendix. Our broad conclusions are unaffected.

Following the three modelling principles discussed in the introduction we choose Taylor-contracts for the supply-side of our model in the euro area and Japan. To match U.S. inflation dynamics we have to compromise on one of the principles, i.e. either pick Taylor's specification which induces insufficent inflation persistence or Fuhrer and Moore's specification which is not quite consistent with optimizing behavior. In this paper, we proceed with Fuhrer-Moore's contracts for the supply side of the U.S. economy.

3 The demand side: output dynamics, monetary policy and international linkages

3.1 Model equations

On the demand side of our model, we need to specify the determination of the output gap, the transmission of monetary policy and international linkages. As discussed in the introduction we take a semi-structural approach, which embodies rational expecations regarding

[17] We also refer the reader to this earlier paper with regard to our reasons for the controversial assumption of a linear trend in inflation. This trend in euro area inflation is due to the gradual convergence process undergone by Italy, France and other high-inflation countries. Thus it may be best understood as resulting from a gradual change in the inflation target of those countries, which should not be attributed to structural wage or price rigidities or other sources of shocks.

future interest rates, inflation and exchange rate changes, in the spirit of macroeconometric models such as Taylor (1993a) or the Federal Reserve's FRB/US model, but does not explicitly incorporate all the restrictions that would be implied by optimizing behavior of a representative agent.

Equation (7) in **Table 2** relates the output gap q_t to several lags of itself, the lagged ex-ante long-term real interest rate r_{t-1} and the trade-weighted real exchange rate ${e_t}^w$. The demand shock $\epsilon_{d,t}$ in equation (7) is assumed to be serially uncorrelated with mean zero and unit variance and is scaled with the parameter σ_{ϵ_d}. A possible 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 contemporaneous value of the real interest rate to allow for a transmission lag of monetary policy. The trade-weighted real exchange rate enters the aggregate demand equation because it influences net exports.[18]

Next we turn to the financial sector and relate the long-term real interest rate to the short-term nominal interest rate, which is the principal instrument of monetary policy. Three equations determine the various interest rates. The short-term nominal interest rate i_t is set according to the interest rate rule defined by equation (8) in **Table 3**. According to this rule policymakers change the nominal interest rate in response to inflation deviations from the policymaker's target π^* and output deviations from potential. This specification accommodates both forecast-based rules (with forecast horizons $\theta > 0$) and outcome-based rules ($\theta = 0$). The inflation measure $\pi_t^{(4)}$ is the annual average inflation rate and the interest rate is annualized. Furthermore, the real equilibrium rate r^* provides a reference point for the policy rule. Note also that this rule simplifies to the one proposed by Taylor (1993a) if $\theta = 0$ and $\rho = 0$.

As to the term structure that is defined in (9), we rely on the accumulated forecasts of the short rate over $\eta(l)$ quarters which, under the expectations hypothesis, will coincide

[18] For now we omit a direct channel through which foreign output affects domestic output, but we plan to explore this channel in future work.

Table 2: Aggregate Demand, Interest Rates and Exchange Rates

Aggregate Demand	$q_t = \delta(L)\, q_{t-1} + \phi\,(r_{t-1} - r^*) + \psi\, e^w_t + \sigma_{\epsilon_d}\, \epsilon_{d,t},$ where $\delta(L) = \sum_{j=1}^{\eta(q)} \delta_j\, L^{j-1}$	(7)
Monetary Policy Rule	$i_t = \rho(L)\, i_{t-1} + (1-\rho(1))\,(\, r^* + \mathrm{E}_t[\, \pi^{(4)}_{t+\theta}\,]\,) + \alpha\, \mathrm{E}_t[\, \pi^{(4)}_{t+\theta} - \pi^*\,] + \beta\, q_t + \sigma_{\epsilon_p}\, \epsilon_{p,t},$ where $\rho(L) = \sum_{j=1}^{\eta(i)} \rho_j\, L^{j-1}$ and $\pi^{(4)}_t = p_t - p_{t-4}$	(8)
Term Structure	$l_t = \mathrm{E}_t \left[\frac{1}{\eta(l)} \sum_{j=1}^{\eta(l)} i_{t+j-1} \right]$	(9)
Real Interest Rate	$r_t = l_t -\, 4\, \mathrm{E}_t \left[\frac{1}{\eta(l)}\, (p_{t+\eta(l)} - p_t) \right]$	(10)
Trade-Weighted Real Exchange Rate	$e^{w,(i)}_t = w_{(i,j)}\, e^{(i,j)}_t + w_{(i,k)}\, e^{(i,k)}_t$	(11)
Open Interest Parity	$e^{(i,j)}_t = \mathrm{E}_t \left[e^{(i,j)}_{t+1} \right] + \left(i^{(j)}_t -\, 4\, \mathrm{E}_t \left[p^{(j)}_{t+1} - p^{(j)}_t \right] \right) - \left(i^{(i)}_t -\, 4\, \mathrm{E}_t \left[p^{(i)}_{t+1} - p^{(i)}_t \right] \right)$	(12)

Notes: q: output gap; r: long-term real interest rate; r^*: equilibrium real interest rate; e^w: trade-weighted real exchange rate; ϵ_d: aggregate demand shock; i : short-term nominal interest rate; π^*: inflation target; $\pi^{(4)}$: year-on-year inflation; ϵ_p: monetary policy shock; l: long-term nominal interest rate; e: bilateral real exchange rate.

with the long rate forecast for this horizon. The term premium is assumed to be constant and equal to zero. We then obtain the long-term ex-ante real interest rate (defined in (10)) by subtracting inflation expectations over the following $\eta(l)$ quarters.

The trade-weighted real exchange rate is defined by equation (11). The superscripts (i, j, k) are intended to refer to the economies within the model without being explicit about the respective economy concerned. Thus, $e^{(i,j)}$ represents the bilateral real exchange rate between countries i and j, $e^{(i,k)}$ the bilateral real exchange rate between countries i and k, and consequently equation (11) defines the trade-weighted real exchange rate for country i. The bilateral trade-weights are denoted by $(w_{(i,j)}, w_{(i,k)}, \ldots)$. Finally, equation (12) constitutes the open interest rate parity condition with respect to the bilateral exchange rate between countries i and j in real terms. It implies that the difference between today's real exchange rate and the expectation of next quarter's real exchange rate is set equal to the expected real interest rate differential between countries i and j.

In the deterministic steady state of this model the output gap is zero and the long-term real interest rate equals its equilibrium value r^*. The equilibrium value of the real exchange rate is normalized to zero. Since the overlapping contracts specifications of the wage-price block do not impose any restriction on the steady-state inflation rate, it is determined by monetary policy alone and equals the target rate π^* in the policy rule.

3.2 Estimation

In estimating the demand side of our model we take an equation-by-equation approach that is simpler than the indirect inference approach used for the supply side. The reason is that the indirect inference approach would require including two more variables in the unconstrained VAR, the interest rate and the exchange rate. This proved rather difficult in our earlier work on the euro area. We now proceed in parallel for the United States, the euro area and Japan and estimate the parameters of the aggregate demand equation (6) by means of the Generalized Method of Moments (GMM). To do so, we first construct the ex-post real long-term rate by replacing expected future with realized values in equations (8) and (9). Then we estimate the parameters by GMM using lagged values of output, inflation and interest rates and real exchange rates as instruments.[19] The estimation results[20] for the aggregate demand equation using U.S., euro area and Japanese data are reported in **Table 3**.

The estimates obtained with U.S. data indicate a hump-shaped output pattern with a positive coefficient on the first lag of output that is greater than one and negative coefficients

[19]Note also that in estimation we use the CPI-based real effective exchange rates rather than the bilateral real effective rates calculated on the basis of the constructed weights. In solving the model we will stick to the endogeneously determined bilateral rates.

[20]Note the sample periods are as follows: U.S. (80:Q1-98:Q4), euro area (80:Q1-98:Q4) and Japan (80:Q1-97:Q1). The differences in length are due to differences in data availability, initial lags, and leads used in constructing long-term rates. As to the term structure equation we used a horizon of two years for the U.S. and the euro area but three years for Japan. In all three equations we used the HP-detrended real effective exchange rate in estimation.

Table 3: Estimated Aggregate Demand Equations: United States, Euro Area and Japan

	δ_1	δ_2	δ_3	ϕ	ψ	σ_{ϵ_d}	p-value $^{(e)}$
United States $^{(a,b)}$	1.2184	-0.1381	-0.2116	-0.0867	0.0188	0.0071	0.9685 [19]
	(0.0320)	(0.0672)	(0.0532)	(0.0193)	(0.0061)		
Euro Area $^{(a,c,d)}$	1.0521	0.0779	-0.1558	-0.0787	0.0188	0.0054	0.9665 [19]
	(0.0381)	(0.0417)	(0.0342)	(0.0335)	(0.0047)		
Japan $^{(a,b)}$	0.9071			-0.0781	0.0122	0.0068	0.9990 [21]
	(0.0124)			(0.0272)	(0.0053)		

Notes: $^{(a)}$ GMM estimates using a constant, lagged values (up to order three) of the output gap, the quartely inflation rate, the short-term nominal interest rate and the real effective exchange rate as instruments. In addition, current and lagged values (up to order two) of the foreign inflation and short-term nominal interest rates have been included in the instrument set. The weighting matrix is estimated by means of the Newey-West (1987) estimator with the lag truncation parameter set equal to the maturity implied by the definition of the long-term nominal interest rate minus one. Estimated standard errors in parentheses. $^{(b)}$ Output gap measure constructed using OECD data. $^{(c)}$ Output measured in deviation from log-linear trend. $^{(d)}$ For the euro area, the German long-term real interest rate has been used in the estimation. Similarly, German inflation and short-term nominal interest rates have been used as instruments. $^{(e)}$ Probability value associated with the test of overidentifying restrictions. Number of overidentifying restrictions in brackets.

on the following lags. We also find a negative effect of the long-term real interest rate on aggregate demand, which is highly statistically significant, and a positive effect of the trade-weighted real exchange rate. The positive sign on the exchange rate coefficient is consistent with a stimulative effect of a real depreciation on net exports since the bilateral exchange rates are defined in terms of domestic over foreign currencies.

As to the euro area, we also find hump-shaped output dynamics, a negative effect of the real interest rate and a positive effect of the real exchange rate. Note however, that we used the German constructed long-term real interest rate in estimation lacking a convincing measure of area-wide real interest rates prior to European Monetary Union. Similarly, we obtain statistically significant estimates of the exchange and interest rate sensitivities with the proper signs for aggregate demand in Japan. However, for Japan the output pattern does not imply a hump-shaped response.

Finally the p-values for the test of overidentifying restrictions that are reported in the last column of **Table 3** indicate that we cannot reject these specifications for any of the three economies.

In principle, the above aggregate demand equations together with the staggered contracts specifications estimated in the preceding section would be sufficient to evaluate the properties of alternative interest rate rules for monetary policy. However, if we want to know how well the complete model fits the data and if we want to identify the historical structural shocks that would be consistent with the complete multi-country model under rational expectations we also need to characterize historical monetary policy. To this end we fit the type of policy reaction function defined by equation (8) in **Table 2** to historical short-term nominal interest rates. Estimates are reported in **Table 4**. In line with other authors we find that a forecast-based version, which implies that short-term nominal interest rates are changed in response to variations of one-year ahead forecasts of inflation and the current output gap performs well in fitting historical interest rates. We allow for partial adjustment by introducing up to two lags of the short-term nominal interest rate in the reaction function. The reaction function for the U.S. interest rate implies a sizeable long-run policy reaction to the forecast of the inflation rate that is substantially greater than one and thereby ensures stability of the model. The response to the output gap turns out to be a good bit smaller.

Since GDP-weighted averages of European interest rates prior to EMU seem unlikely to be appropriate as a measure of the euro-area-wide historical monetary policy stance, we resort to estimating a reaction function for the German interest rate that we have already used in estimating euro area aggregate demand as discussed above. The German estimates, which are reported in the second row of **Table 4**, also indicate a stabilizing inflation response but no output response.[21] Finally, for Japan we also estimate a significant

[21] Work by Clarida, Galí and Gertler (1998) suggests that German interest rate policy since 1979 is summarized quite well by such a forecast-based interest rate rule. Clarida et al. (1998) also argue that German monetary policy had a strong influence on interest rate policy in the U.K., France and Italy throughout this period and may have led to higher interest rates in those countries than warranted by domestic conditions at the time of the EMS crisis as suggested in Wieland (1996).

Table 4: Estimated Monetary Policy Rules: United States, Germany and Japan

	ρ_1	ρ_2	α	β	σ_{ϵ_p}	p-value [(d)]
United States [(a,b)]	0.7745		0.2851	0.0840	0.0110	0.8474 [9]
	(0.0620)		(0.1081)	(0.0428)		
Germany [(a,b,c)]	1.1169	-0.3480	0.2039		0.0054	0.7155 [12]
	(0.0739)	(0.0594)	(0.0383)			
Japan [(a,b)]	1.2672	-0.4160	0.1239		0.0048	0.4057 [9]
	(0.1385)	(0.1069)	(0.0396)			

Notes: [(a)] GMM estimates using a constant, lagged values (up to order three) of the output gap, the quartely inflation rate, the short-term nominal interest rate and the real effective exchange rate as instruments. The weighting matrix is estimated by means of the Newey-West (1987) estimator with the lag truncation parameter set equal to four. Estimated standard errors in parentheses. [(b)] Output gap measure constructed using OECD data. [(c)] For Germany, lagged values (up to order three) of government consumption relative to potential output have been included in the instrument set. [(d)] Probability value associated with the test of overidentifying restrictions. Number of overidentifying restrictions in brackets.

stabilizing response to inflation but not to the output gap.[22]

4 The empirical fit of the multi-country model

Having estimated demand and supply-side equations of our model separately a natural question is how well the complete model fits the data. To assess the overall fit of the structural multi-country model we start by computing the implied historical series of structural shocks. The relevant sample period is 1980:Q1 to 1998:Q4. These structural shocks differ from the single-equation residuals, because expectations of future variables are computed to be consistent with the complete model. Interest rates are set according to the estimated forecast-based policy rules. Long-term rates satisfy the term-structure and Fisher relationships. Nominal exchange rates are flexible, but satisfy the open interest rate parity condition. Our investigation of historical structural shocks indicates that the implied

[22]The sample periods for the three regressions are as follows: U.S. (79:Q4-99:Q4), Germany (79:Q2-98:Q4) and, Japan (79:Q2-95:Q2). The choice of starting dates for the U.S. and the euro area was motivated by earlier work of Clarida et al., while the end date for Japan was chosen so as to exclude the zero-nominal-interest rate period.

Figure 1: Correlation Pattern of Historical Structural Shocks

Notes: Solid bars: Correlation functions implied by the complete multi-country model. Dotted lines: Asymptotic 95%-confidence bands.

demand and supply shocks are sufficiently close to white noise. The correlogram of these shocks is shown in **Figure 1**.

Overall, the correlograms of the historical shocks do not reveal significant serial correlation.[23] This finding provides some support that our model fits the historical sample reasonably well.

As a further test of the fit of our model we compute the implied autocorrelation func-

[23] The only exception is a significant 4-th order correlation for euro area contract wage shocks.

tions of inflation and output and compare them to the empirical autocorrelation functions implied by the unconstrained bivariate VARs for each country.[24] Such an approach has also been used by Fuhrer and Moore (1995a) and by McCallum (2001), who argued that autocorrelation functions are more appropriate for confronting macroeconomic models with the data than impulse response functions because of their purely descriptive nature.

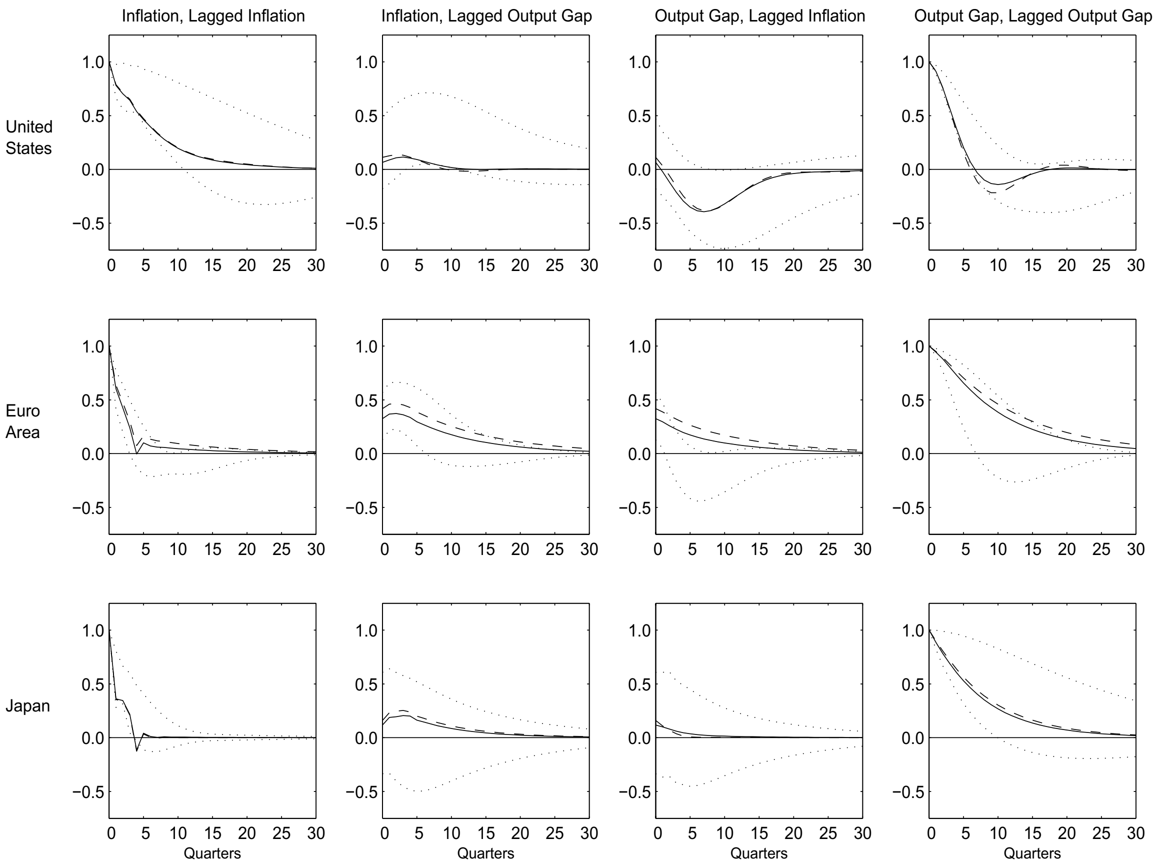
The comparison of autocorrelation functions of inflation and output in the three economies is reported in **Figure 2**. The solid lines refer to the autocorrelation functions implied by the complete multi-country model. They are derived with interest rates set according to the estimated forecast-based policy rules, flexible nominal exchange rates and aggregate demand and contract wage shocks drawn from the covariance matrix of historical structural shocks. The first panel in each row shows the autocorrelations of inflation, the second and third panel the lagged cross-correlations of inflation and output and the fourth panel the autocorrelations of output. The thin dotted lines in each panel of **Figure 2** correspond to the asymptotic 95% confidence bands associated with the autocorrelation functions of the individual bivariate unconstrained VAR(3) models used in the estimation of the staggered contracts specifications.[25]

The autocorrelation functions implied by the structural multi-country model fall within the confidence bands implied by the unconstrained VARs. Thus, the structural model appears to fit inflation and output dynamics quite well, in particular in light of the fact that the estimation was carried out with limited information methods. We summarize that the model is capable to match both, inflation persistence (panels in the first column) and output persistence (panels in the fourth column) for all three economies. As to the cross-correlations of output and inflation, we find that high output tends to lead to high inflation in subsequent quarters (second column of panels). However, the confidence bands tend to be quite wide.

[24] These are the VARs that served as approximating probability models in the estimation of the contracting parameters.

[25] 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).

Figure 2: Fitting Inflation and Output Dynamics with the Structural Model



Notes: Solid line: Autocorrelation functions implied by the complete multi-country model. Dash-dotted line: Autocorrelation functions implied by the single-country models. Dotted lines: Asymptotic 95%-confidence bands implied by bivariate unconstrained VAR's of inflation and output.

Finally, the dot-dashed lines in **Figure 2** correspond to autocorrelation functions based on three separate models of these economies without international linkages. They are based on re-estimated aggregate demand equations which do not include the effective real exchange rate.[26] Typically these correlation functions also remain within the confidence bands implied by the bivariate unconstrained VARs. This provides a first indication that accounting

[26]Estimates for the closed-economy aggregate demand equations are reported in **Table B** in the appendix.

for the exchange rate channel affecting aggregate demand is not crucial to capturing the observed degree of output and inflation persistence. Furthermore, the comparison with the autocorrelation functions implied by the multi-country model shows that the exchange rate channel introduces noticeable but relatively small changes in output and inflation persistence.

5 International linkages and spillovers

We now turn to consider the magnitude of international spillover effects within our model. We maintain the assumption of flexible nominal exchange rates, which implies that nominal interest rates can be set independently by the central banks of the three economies. However, we want to focus on the differences in adjustment to unexpected shocks that arise from differences in the structure of the economies rather than from differences in monetary policy. For this reason we assume that each central bank implements Taylor's rule:

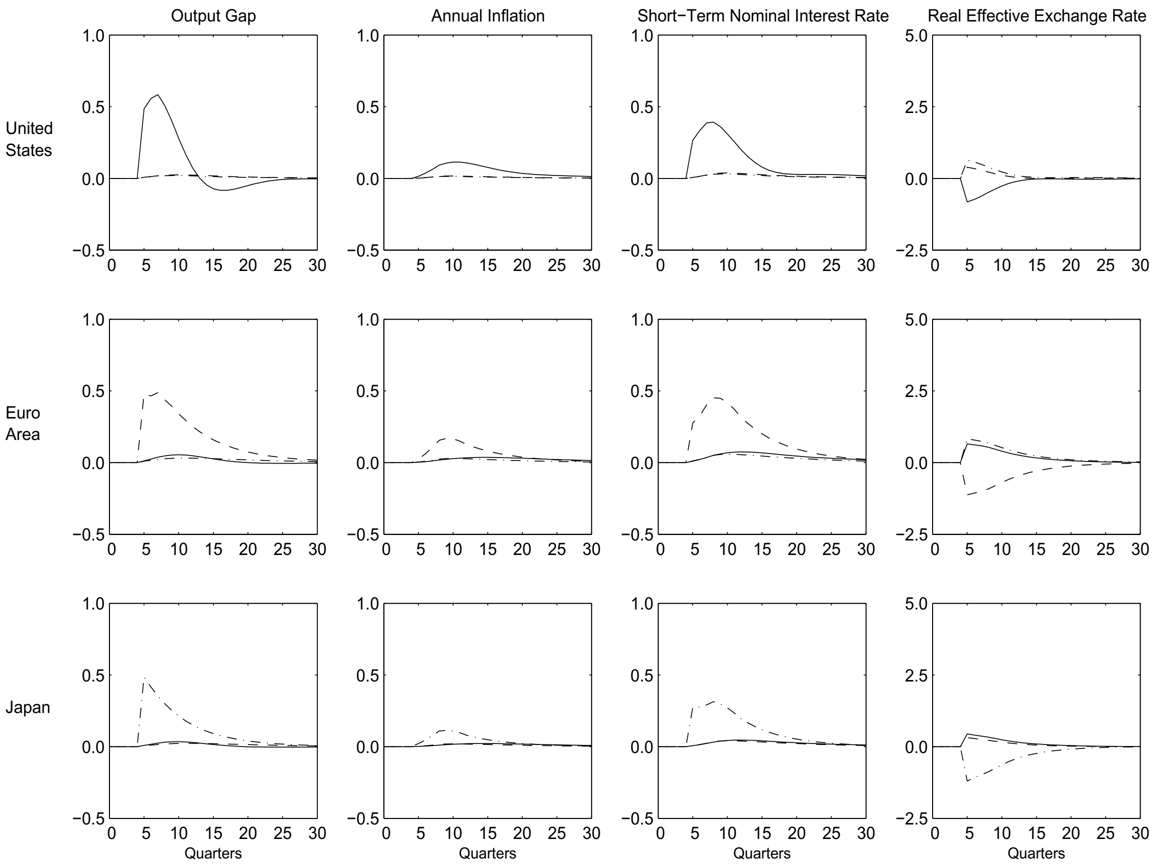
$$i_t = r^* + \pi^* + 1.5\,(\pi_t^{(4)} - \pi^*) + 0.5\,q_t, \tag{13}$$

where $\pi_t^{(4)}$ again stands for annual average inflation, π^* for the inflation target and r^* for the real equilibrium interest rate.

We start by evaluating the consequences of an unexpected temporary demand shock of 0.5 percentage points of potential output in each of the three economies. The dynamic responses of output, inflation, short-term nominal interest rates and real effective exchange rates for the three economies are shown in **Figure 3**. Solid lines refer to U.S. variables, dashed lines to euro area variables and dot-dashed lines to Japanese variables.

In response to the positive demand shock in the euro area shown in the middle row of **Figure 3** euro area output rises for 2 quarters and then declines again. The positive output gap induces a temporary increase in inflation. In response, euro area nominal interest rates rise sufficiently so as to induce higher real interest rates and counterbalance the increase in output and inflation. The existence of the exchange rate channel introduces a second counterbalancing force, because the euro appreciates in nominal and real terms vis-à-vis the

Figure 3: Comparing Demand Shocks in the United States, the Euro Area and Japan

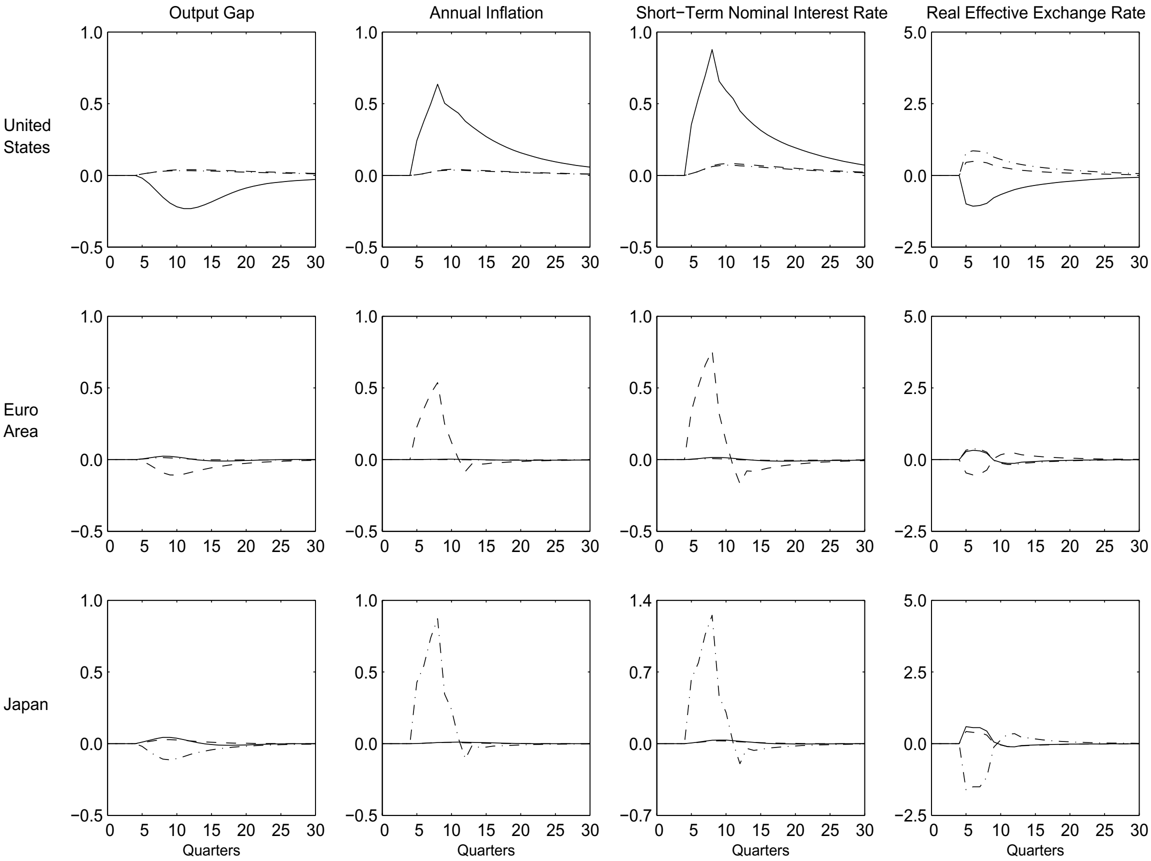


Notes: Solid line: United States. Dashed line: Euro Area. Dot-dashed line: Japan.

other currencies. The depreciation of the Yen and US$ has only a very small expansionary effect in Japan and the United States. As output returns to potential and inflation to the central bank's target, the real exchange rates also return to their equilibrium values.

The top and bottom rows of **Figure 3** show the consequences of aggregate demand shocks of 0.5 percentage points in the United States and Japan. Qualitatively, the demand shocks have the same consequences in each economy, however they exhibit some quantitative differences. For example, the inflationary impact of the demand shock is largest in the

Figure 4: Comparing Contract Wage Shocks in the United States, the Euro Area and Japan



Notes: Solid line: United States. Dashed line: Euro Area. Dot-dashed line: Japan.

United States due to the Fuhrer-Moore style overlapping contracts specification, which induces more inflation persistence. Also, output in Japan does not exhibit a hump-shaped pattern as would be expected given the estimated coefficient on the lags of the output gap in the aggregate demand equation.

We also compare the domestic and international consequences of a short-run supply shock, that is, a shock to the contract wage equations in our model. As shown in **Figure 4** a positive contract wage shock puts upward pressure on inflation. Monetary policy responds by increasing interest rates. As a result, output declines and the exchange rate appreciates. International consequences are again minor when nominal exchange rates are flexible. The output costs of stabilizing inflation are not surprisingly largest in the United States, for which we have chosen a supply-side specification with Fuhrer-Moore contracts.

6 Monetary policy rules and the role of the exchange rate

Finally we turn to assessing the stabilization performance of alternative monetary policy rules and the role of the exchange rate for monetary policy within our multi-country model. We report initial results from an exploratory analysis that indicates some interesting avenues for further work but should still be considered preliminary.

Our starting point is an evaluation of simple outcome-based interest rate rules, which respond to the annual average inflation rate, the output gap and the lagged short-term nominal interest rate. An important argument in favor of such rules is that they seem to be surprisingly robust to model uncertainty (see for example the introduction to Taylor, ed., (1999) and Levin, Wieland and Williams (1999, 2001)). However, from a perspective of monetary policymaking in open economies such rules have been criticized for lacking an explicit feedback to the exchange rate (cf. Ball (1999), Batini and Haldane (1999), Svensson (2000) and Taylor (2001)). Thus, in this paper, we consider the following class of policy rules:

$$i_t = \rho\, i_{t-1} + (1-\rho)\,(r^* + \pi_t) + \alpha\,(\pi_t^{(4)} - \pi^*) + \beta\, q_t + \xi\, e_t^w. \tag{14}$$

First, we restrict the response parameter ξ on the exchange rate to be equal to zero and choose the other three response parameters so as to minimize the policymaker's loss function. We assume that this loss function equals the weighted average of the unconditional variances of inflation and output gaps,

$$\mathcal{L} = \mathrm{Var}[\pi_t] + \lambda \mathrm{Var}[q_t]. \tag{15}$$

Following the approach in Levin et al. (1999, 2001) we minimize the loss function subject to the constraint that the volatility of the change of the nominal interest rate is no greater

Table 5: Optimized Simple Policy Rules without Exchange Rate Feedback.

	United States			Euro Area			Japan		
λ	ρ	α	β	ρ	α	β	ρ	α	β
0	0.94	0.38	0.04	0.99	0.26	0.05	1.02	0.18	0.02
1/3	0.92	0.25	0.16	0.90	0.11	0.24	0.99	0.13	0.09
1	0.91	0.16	0.20	0.89	0.00	0.35	0.95	0.04	0.18
3	0.89	0.08	0.23	0.92	0.00	0.35	0.96	0.00	0.20

Notes: For each of the countries (j) and for each preference parameter (λ), this table indicates the optimal coefficients (ρ, α and β).

than under the estimated policy rules reported in **Table 4**.

The coefficients of simple rules optimized in this manner for four alternative values of the weight $\lambda = (0, 1/3, 1, 3)$ are reported in **Table 5**. As in Levin et al (1999) we find that the optimal value of the interest-rate smoothing coefficient ρ is near unity. Also, the value of the coefficent on inflation, α, (on output, β) decreases (increases) with a greater weight on output variability in the loss function. The optimized response coefficients for the three economies are surprisingly similar given that the estimated model specifications and dynamics are quite different. Somewhat surprisingly the inflation coefficient α is almost equal to zero for the euro area and Japan when we consider a weight on output equal or greater unity.[27]

In a second step, we relax the restriction of no exchange rate feedback and optimize over all four response coefficients $(\rho, \alpha, \beta, \xi)$. The resulting coefficients and the percent change in the loss function relative to rules without exchange rate feedback are reported in **Table 6**. For each of the three economies the optimal coefficient on the exchange rate turns out to be near zero, while the associated reduction in the loss function remains negligible.

[27] We have investigated whether this result is robust to alternative values of the constraint on interest rate volatility and found that allowing for a higher degree of interest rate volatility (i.e. more aggressive policy rules) induces positive response coefficients on inflation also for higher values of the weight λ.

Table 6: Optimized Simple Rules with Exchange Rate Feedback

	United States					Euro Area					Japan				
λ	ρ	α	β	ξ	$\mathcal{L}$	ρ	α	β	ξ	$\mathcal{L}$	ρ	α	β	ξ	$\mathcal{L}$
0	0.94	0.36	0.04	-0.00	-0.01	1.00	0.27	0.05	0.00	-0.01	1.02	0.18	0.02	0.00	-0.00
1/3	0.92	0.24	0.15	-0.00	-0.01	0.91	0.13	0.26	0.01	-0.03	0.99	0.14	0.09	0.00	-0.01
1	0.91	0.17	0.21	0.00	-0.01	0.90	0.00	0.43	0.02	-0.29	0.96	0.05	0.20	0.00	-0.04
3	0.91	0.11	0.25	0.02	-0.10	0.93	0.00	0.50	0.04	-1.59	0.97	0.00	0.23	0.01	-0.25

Notes: For each of the countries (j) and for each preference parameter (λ), this table indicates the optimal coefficients (ρ, α, β and ξ) and the percentage change in the policy-makers' loss functions ($\mathcal{L}$) compared with the losses under optimized policy rules which exclude the real exchange rate.

Thus, the existence of the exchange rate channel in the open economy does not seem to require an economically significant direct response of policy to the exchange rate within our multi-country model.

Table 7: Coefficients of Cooperatively Optimized Monetary Policy Rules

	United States			Euro Area			Japan		
λ	ρ	α	β	ρ	α	β	ρ	α	β
0	0.95	0.38	0.04	0.99	0.25	0.05	1.01	0.17	0.02
1/3	0.92	0.25	0.15	0.89	0.10	0.24	0.98	0.12	0.08
1	0.91	0.16	0.19	0.88	0.00	0.33	0.94	0.03	0.17
3	0.89	0.08	0.22	0.91	0.00	0.33	0.94	0.00	0.20

Notes: For each of the countries (j) and for each preference parameter (λ), this table indicates the optimal coefficients (ρ, α and β).

Another issue that has been widely debated in open-economy macroeconomics concerns the gains from international monetary policy coordination. To obtain a first quantitative assessement regarding this question we derive the optimal policy coefficients under monetary cooperation with the objective to minimize the average of the losses in the three economies. These coefficients, which are shown in **Table 7**, turn out to be surprisingly similar to the optimal coefficients under nationally-oriented monetary policies. Furthermore, as shown in **Table 8** the percentage reduction in losses that any one country could achieve by deviating from the cooperative policy unilaterally is rather small. Thus, there seems to be little to loose but also little to gain from international monetary cooperation in the context of stabilization policy within our model. However, to settle this question satisfactorily we still need to compute the non-cooperative Nash equilibrium, for which all three policy-makers minimize national (or area-wide) losses.

Table 8: Stabilisation Gains of Self-Oriented National Monetary Policies

	United States			Euro Area			Japan		
λ	$\mathcal{L}_{\mathrm{US}}$	$\mathcal{L}_{\mathrm{EA}}$	$\mathcal{L}_{\mathrm{JA}}$	$\mathcal{L}_{\mathrm{US}}$	$\mathcal{L}_{\mathrm{EA}}$	$\mathcal{L}_{\mathrm{JA}}$	$\mathcal{L}_{\mathrm{US}}$	$\mathcal{L}_{\mathrm{EA}}$	$\mathcal{L}_{\mathrm{JA}}$
0	-0.04	0.01	0.01	0.03	-0.02	0.01	0.10	0.02	-0.02
1/3	-0.07	0.06	0.03	0.14	-0.10	0.04	0.26	0.08	-0.08
1	-0.11	0.14	0.06	0.18	-0.18	0.06	0.30	0.19	-0.21
3	-0.19	0.34	0.14	0.23	-0.33	0.11	0.45	0.36	-0.46

Notes: For each of the countries (j) and for each preference parameter (λ), this table indicates the percentage point change in the policy-makers' loss functions ($\mathcal{L}_j$) when monetary policy of a single country is conducted in a self-oriented manner compared with the losses under the cooperatively optimized policy rules.

In our view, the stark results regarding the role of the exchange rate and international policy coordination within our multi-country model require further corroboration by means of sensitivity studies. As a first step in this direction, we re-consider the channel through which the exchange rate affects the domestic economies in our model. So far, we have only included an expenditure-switching effect on aggregate demand. A natural extension would be to include a direct effect of the exchange rate on prices in each economy. To this end we need to distinguish between prices for domestic goods and import prices which may be directly affected by the exchange rate depending on the degree of exchange-rate pass-through. Rather than going back and re-estimating the supply-side of our model, we simply add this channel to the existing model by defining, for each country i, the overall price level $p_t^{all,(i)}$ as the weighted sum of the domestic price level $p_t^{(i)}$ and a direct effect of the bilateral nominal exchange rates $e_t^{n,(j)}$ via import prices,

$$p_t^{all,(i)} = (1-\chi)\, p_t^{(i)} + \chi \sum_{j \neq i} w_{(i,j)}\, (e_t^{n,(j)} + p_t^{(j)}),$$

where χ measures the share of import prices and exchange-rate pass-through is assumed to be immediate and complete.[28]

Table 9: Openness and Optimized Monetary Policy Rules for the United States

	$\chi = 0.05$					$\chi = 0.10$					$\chi = 0.20$				
λ	ρ	α	β	ξ	$\mathcal{L}$	ρ	α	β	ξ	$\mathcal{L}$	ρ	α	β	ξ	$\mathcal{L}$
Domestic-Inflation Target															
0	0.88	3.79	-0.02	-0.09	-6.69	0.86	6.50	-0.04	-0.06	-1.48	0.90	4.77	-0.02	0.06	-0.75
1/3	0.81	2.80	0.31	-0.07	-0.20	0.75	5.97	0.27	0.03	-0.01	0.82	4.63	0.09	0.08	-0.06
1	0.87	1.36	0.50	0.04	-0.05	0.70	4.17	0.69	0.03	-0.00	0.73	4.20	0.32	0.15	-0.06
3	1.01	0.67	0.66	0.05	-0.13	0.83	2.15	1.31	0.22	-0.14	0.72	2.71	0.79	0.37	-0.19
Overall-Inflation Target															
0	0.85	7.86	-0.07	-0.03	-2.23	0.95	4.66	-0.04	0.07	-2.55	1.04	2.34	-0.01	0.31	-14.56
1/3	0.63	6.05	0.45	0.01	-0.00	0.83	4.31	0.14	0.09	-0.19	0.99	2.24	0.03	0.32	-2.26
1	0.67	2.99	0.84	0.03	-0.02	0.71	3.51	0.43	0.14	-0.19	0.92	2.07	0.11	0.33	-0.94
3	0.88	0.94	0.87	0.05	-0.05	0.71	1.98	0.91	0.25	-0.31	0.78	1.62	0.32	0.38	-0.48

Notes: For each of the different degrees of openness (χ), for each preference parameter (λ) and for the alternative inflation targets (π^*), this table indicates the optimal coefficients (ρ, α, β and ξ) and the percentage change in the policy-makers' loss functions ($\mathcal{L}$) compared with the losses under optimized policy rules which exclude the real exchange rate.

[28] For recent studies of exchange rate pass-through see Campa and Goldberg (2002) and Gagnon and Ihrig (2002).

Table 9 reports the coefficients of optimized policy rules in the United States given alternative values for the direct effect of exchange rates on U.S. overall inflation (i.e. $\chi = 0.05, 0.10$ or 0.20). We consider two different scenarios, one in which monetary policy targets domestic inflation only and one where it targets overall inflation. Wage-setters are assumed to look at overall inflation.

Introducing a direct effect of the exchange rate on prices clearly changes some of our earlier results. First, the optimal policy response to inflation is now rather large. Secondly, we now obtain more substantial gains from including the exchange rate in the policy rule. These gains are greatest when the policymaker targets overall inflation and the share of import prices is rather large. In our ongoing work we are investigating the robustness of these findings.

7 Conclusion

Our empirical analysis of inflation and output dynamics in the United States, the euro area and Japan has provided some new results regarding the role of staggered contracts as sources of nominal rigidity and inflation persistence. First, we find it impossible to match inflation dynamics with Calvo-style random-duration contracts except if we allow for a significant share of price setters with backward-looking behavior or for ad-hoc persistence in supply shocks. For Taylor-style fixed duration contracts, however, we obtain economically meaningful and statistically significant parameter estimates for all three economies. Fuhrer-Moore-style contracts only dominate Taylor-style contracts for U.S. inflation dynamics. One possible interpretation of this finding is that the United States just suffer from a higher degree of nominal rigidity than the euro area or Japan. A plausible alternative interpretation is that this difference in historical inflation persistence may be due to a difference in historical monetary policy. In particular, since U.S. monetary policy accommodated oil-price induced inflation increases in the 1970s, much more than Japanese or German monetary policymakers, the higher degree of historical inflation persistence may also have been caused by a lack of credibility to keep inflation under control in the late 1970s and early 1980s (see also Erceg and Levin (2001) on this point).

After completing our macro-econometric model with an admittedly more ad-hoc specification of the demand side we find that it fits inflation and output dynamics quite well. Including or excluding the real exchange rate channel on the demand side does not seem to matter much for the ability of this model to account for the observed degree of inflation and output persistence.

International spillovers turn out to be rather small as long as nominal exchange rates are flexible. Our preliminary investigation of optimized simple policy rules indicates little gain from a direct policy response to the real exchange rate. Furthermore, potential gains from international monetary coordination of stabilization policies seem rather limited. Such coordination may be more important in exceptional circumstances, for example when interest rate policy in one economy is constrained by the zero bound on nominal interest rates.[29]

[29] See for example, Svensson (2001), McCallum (2000).

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Appendix

A.1 Estimation Methodology

The unconstrained VARs that we use to summarize empirical output and inflation dynamics for the three economies under consideration take the following form,

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

where q_t refers to the output gap and π_t to inflation and the error terms $u_{\pi,t}$ and $u_{q,t}$ are assumed to be serially uncorrelated with mean zero and covariance matrix Σ_u.

We find that a maximum lag length of 3, which corresponds to a maximum contract length of 1 year, is sufficient to capture the observed degree of inflation and output persistence for the economies under consideration.

The second stage of the estimation procedure involves the estimation of the structural parameters of the staggered contracts models using the unconstrained VARs as approximating probability models. Of course, the overlapping contracts specifications alone do not represent a complete model of inflation determination. Since the contract wage equations 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 for each contract specification. A full-information estimation approach would require estimating all the structural parameters of the complete multi-country model jointly. We take a less ambitious approach and simply use the output gap equation from the unconstrained VAR (the second row in (A.1)) as an auxiliary equation for output determination.[30]

Using the output equation from the unconstrained VAR together with the wage-price block, we can solve for the reduced-form inflation and output dynamics under each staggered contracts specification.[31] In the case Taylor- and Fuhrer-Moore-Style contracts it is convenient to rewrite the wage-price block in terms of the real contract wage $(x-p)_t$ and the annualized quarterly inflation rate π_t. The reduced-form of these models 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 $(x-p)_t$ is unobservable. Given a maximum contract length of one year this constrained VAR can be written as follows:

$$\begin{bmatrix} (x-p)_t \\ \pi_t \\ q_t \end{bmatrix} = B_1 \begin{bmatrix} (x-p)_{t-1} \\ \pi_{t-1} \\ q_{t-1} \end{bmatrix} + B_2 \begin{bmatrix} (x-p)_{t-2} \\ \pi_{t-2} \\ q_{t-2} \end{bmatrix} + B_3 \begin{bmatrix} (x-p)_{t-3} \\ \pi_{t-3} \\ q_{t-3} \end{bmatrix} + B_0\, \epsilon_t, \qquad (A.2)$$

where 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 $(i = 0, 1, 2, 3)$

[30] This limited-information approach follows Taylor (1993a) and Fuhrer and Moore (1995a).

[31] We employ the AIM algorithm of Anderson and Moore (1985), which uses the Blanchard and Kahn (1980) method for solving linear rational expecations models, to compute model-consistent expectations.

coincide exactly with the coefficients of the output gap equation of the unconstrained VAR, with the B_0 coefficients 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 $(f_i, \gamma, \sigma_{\epsilon_x})$ as well as the coefficients of the output gap equation of the unconstrained VAR.

In fitting the constrained VAR we employ the indirect inference methods proposed by Smith (1993) and Gouriéroux, Monfort and Renault (1993) and developed further in Gouriéroux and Monfort (1996) to estimate the structural parameters f_i, γ and σ_{ϵ_x}. Indirect inference is a simulation-based procedure that provides a precise way of comparing a model to the data by comparing key characteristics, which themselves are quantities that require estimation via an auxiliary model.[32] In our case, the aim of the estimation procedure is to find values of the structural parameters such that the degree of inflation persistence exhibited by the structural model matches the persistence in the inflation data as summarized by the inflation equation of the unconstrained VAR models discussed above.

An advantage of this indirect inference procedure is that the approximating probability model (the unconstrained VAR) does not require controversial identifying assumptions. Furthermore, since the VAR parameters also determine the autocovariance functions of inflation and output, matching those parameters is essentially equivalent to matching the autocorrelations and cross-correlations of the VAR. In this sense, indirect inference based on the estimated parameters of the unconstrained VAR model is an efficient and robust way to make use of the relevant information contained in the data. By contrast, informal model calibration techniques, but also methods-of-moments based estimation, typically rely on a small set of often subjectively chosen standard deviations and autocorrelations directly inferred from the data.

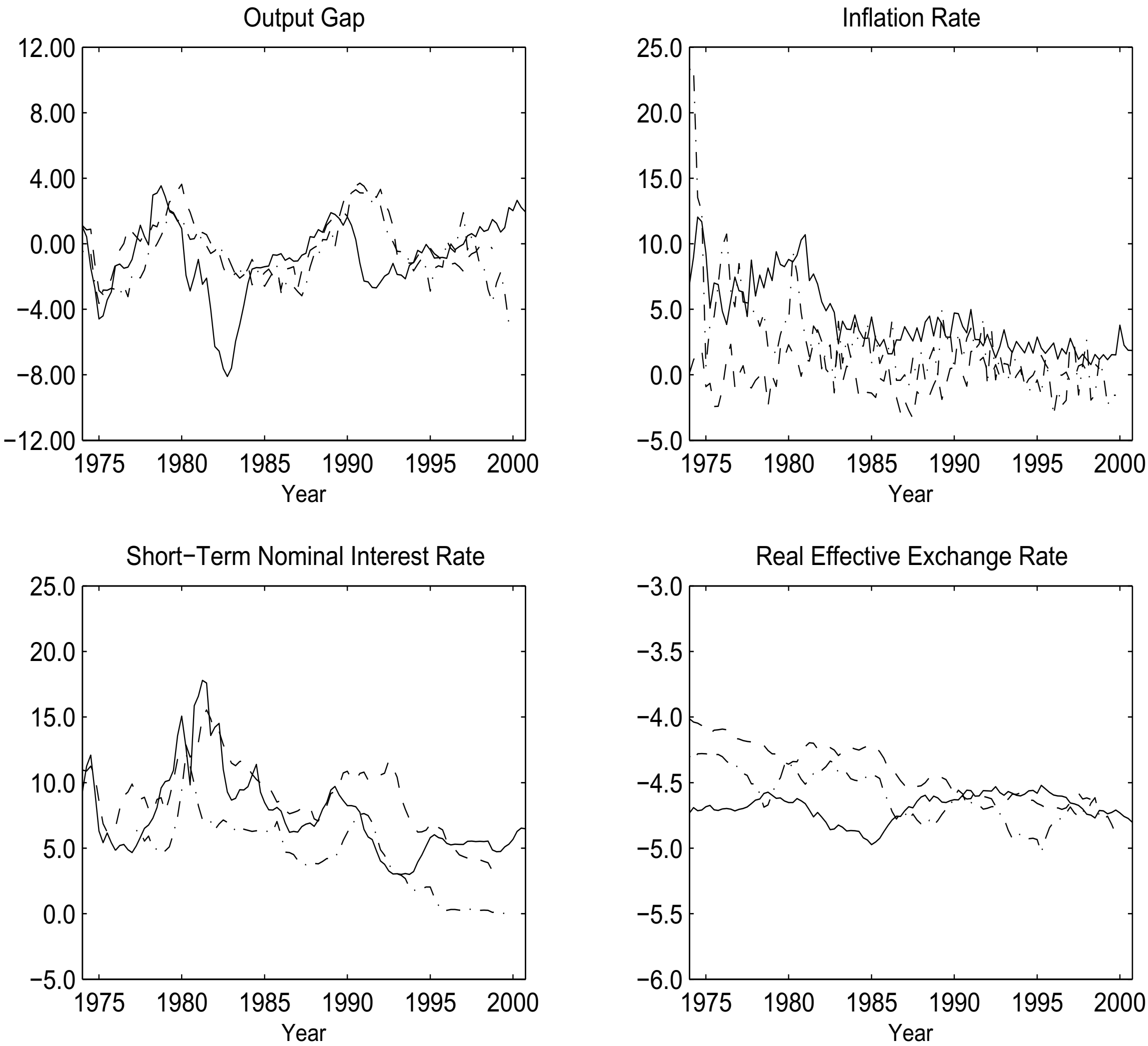
Of course, one cannot always directly match the parameters of the constrained VAR model (A.2) with the parameters of the unconstrained VAR model (A.1) because the constrained model also includes the real contract wage, which is unobservable. Instead, we first simulate the constrained VAR to generate "artificial" series for the real contract wage, the inflation rate and the output gap for given values of the structural parameters and the parameters of the reduced-form output gap equation.[33] In a second step, we then fit the unconstrained VAR model to the inflation and output gap series generated in this manner 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 parameters.

[32]Formally, indirect inference provides a rigorous statistical foundation for data-based calibration techniques, which have become increasingly popular in macroeconomic modelling in recent years. The procedure itself including its asymptotic properties, is discussed in detail in the appendix of the working paper version of Coenen and Wieland (2000). There, we also provide a comparison to the maximum-likelihood methods used by Taylor (1993a) and Fuhrer and Moore (1995a).

[33]All that is needed for simulation are three initial values for each of these variables and a sequence of random shocks. In estimation we use steady-state values as initial conditions. We drop several years of data from the simulations so as to avoid an estimation bias due to these initial conditions.

A.2 The Data

Figure A: The Data for the United States, the Euro Area and Japan



Notes: Solid line: United States. Dashed line: Euro Area. Dot-dashed line: Japan.

A.3 Additional Estimation Results

Table A: Estimated Contracting Specifications with Restrictions on Contract Weights

	s	γ	σ_{ϵ_x}	p-value $^{(d)}$
Taylor Wage Contracts				
United States $^{(a,b)}$	0	0.0093 (0.0056)	0.0053 (0.0002)	$< 10^{-13}$ [5]
Euro Area $^{(a,c)}$	0.0456 (0.0465)	0.0115 (0.0053)	0.0038 (0.0005)	0.3186 [4]
Japan $^{(a,b)}$	0.0086 (0.0153)	0.0178 (0.0055)	0.0080 (0.0004)	0.0083 [4]
Fuhrer-Moore Wage Contracts				
United States $^{(a,b)}$	0	0.0118 (0.0041)	0.0034 (0.0002)	0.6426 [5]
Euro Area $^{(a,c)}$	0.0742 (0.0245)	0.0212 (0.0048)	0.0024 (0.0003)	0.2602 [4]
Japan $^{(a,b)}$	0.0771 (0.0284)	0.0048 (0.0046)	0.0044 (0.0006)	0.0049 [4]

Notes: $^{(a)}$ Simulation-based indirect estimates using a VAR(3) model of quarterly inflation and the output gap as auxiliary model. Estimated standard errors in parentheses. $^{(b)}$ Output gap measure constructed using OECD data. $^{(c)}$ Inflation in deviation from linear trend and and output in deviation from log-linear trend. $^{(d)}$ Probability value associated with the test of overidentifying restrictions. Number of overidentifying restrictions in brackets.

Table B: Estimated Aggregate Demand Equations for the Closed Economies

	δ_1	δ_2	δ_3	ϕ	σ_{ϵ_d}	p-value$^{(e)}$
United States$^{(a,b)}$	1.2297	-0.1063	-0.2586	-0.1436	0.0072	0.5455 [5]
	(0.0678)	(0.1141)	(0.0864)	(0.0454)		
Euro Area$^{(a,c,d)}$	1.0398	0.0510	-0.1183	-0.0832	0.0054	0.3114 [5]
	(0.0985)	(0.0774)	(0.0678)	(0.0695)		
Japan$^{(a,b)}$	0.9374			-0.0815	0.0068	0.5733 [7]
	(0.0237)			(0.0498)		

Notes: $^{(a)}$ GMM estimates using a constant, lagged values (up to order three) of the output gap, the quartely inflation rate and the short-term nominal interest rate. The weighting matrix is estimated by means of the Newey-West (1987) estimator with the lag truncation parameter set equal to the maturity implied by the definition of the long-term nominal interest rate minus one. Estimated standard errors in parentheses. $^{(b)}$ Output gap measure constructed using OECD data. $^{(c)}$ Output measured in deviation from log-linear trend. $^{(d)}$ For the euro area, the German long-term real interest rate has been used in the estimation. Similarly, German inflation and short-term nominal interest rates have been used as instruments. $^{(e)}$ Probability value associated with the test of overidentifying restrictions. Number of overidentifying restrictions in brackets.

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