FOLLOWING GERMANY'S LEAD: USING INTERNATIONAL MONETARY LINKAGES TO ESTIMATE THE EFFECT OF MONETARY POLICY ON THE ECONOMY

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Abstract—Forward-looking behavior on the part of the monetary authority makes it difficult to estimate the effect of monetary policy interventions on output. We present instrumental variables estimates of the impact of interest rates on quarterly real output for several European countries, using German interest rates as the instrument. These estimates confirm a strong forward-looking bias in least squares estimates that persists even conditional on standard controls for the history of the system. Due to the potential for correlation of output shocks across countries, we interpret our estimates as lower bounds for the effect of monetary policy on real output.

I. Introduction

A RECURRING question in economics is the extent to which monetary policy interventions affect the real economy. Assessing the magnitude of these effects empirically is inherently difficult because of central-bank efforts at anticipating trends in growth and inflation. As discussed in the literature, this forward-looking aspect of monetary policy imparts a downward bias on the estimates of the real impact of interventions.[1]

One solution to this problem is to isolate sources of variation in monetary policy that are not themselves correlated with the economic outcomes of interest. Although such variation is rare in a macroeconomic setting, institutional arrangements may occasionally constrain a central bank's behavior and lead to deviations from systematic forward-looking monetary policy. Thus, these institutional features can sometimes lead to recurring "natural" experiments in monetary policy.[2]

In this paper, we present instrumental variables (IV) estimates of the effect of monetary policy on real output for several European countries in the pre-EMU period, using German interest rates as the instrument. We thereby exploit quasi-experimental variation in interest rates generated by the adherence of countries to the interest rates of Germany as an "anchor" country within a system of fixed exchange rates.[3] The economic and institutional environment in Europe during the time frame of our study allows us to give our IV estimator a clear economic interpretation. We develop a simple framework that allows us to explicitly address several threats to interpretation, in particular the role of common output shocks.

Our primary findings are as follows. First, as other researchers (such as Clarida, Galí, & Gertler, 1998) have found, we document a strong and precisely estimated correlation of home interest rates with German interest rates. This relationship is sufficiently strong that the important statistical concerns with weak instruments are not an issue here (Stock, Wright, & Yogo, 2002). Second, IV estimates confirm powerful real effects of monetary policy. Our estimates suggest that a 5 percentage point increase in interest rates leads to a recessionary contraction in annual real growth of 2 to 3 percentage points. Third, IV estimates are decisively more negative than OLS—typically three to four times as large in magnitude—suggesting a strong forward-looking component of European monetary policy in the pre-EMU period.[4] Fourth, for large countries with low trade integration with Germany—where any remaining bias of IV due to common output shocks is expected to be low—IV estimates are most negative and the difference between OLS and IV is highest.

For countries whose output shocks are correlated with those of Germany, our IV approach will not fully eliminate the endogenous component of monetary policy. We argue

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[1] The empirical literature on the effect of monetary policy on the economy is vast; see Christiano, Eichenbaum, and Evans (1999). Isolating exogenous variation in monetary policy to estimate its effect has been a major theme in the literature. Andersen and Jordon (1968) is an early paper showing how policy endogeneity can affect estimates of the impact of monetary policy on the economy in the context of a "St. Louis Equation." A recent paper by Romer and Romer (2004) follows in the tradition of Friedman and Schwartz (1963), seeking to identify periods of exogenous shifts in monetary policy. A large number of papers use the vector autoregression (VAR) approach developed by Sims (1972, 1980a, 1980b) to estimate the effect of monetary policy. For recent summaries, see, for example, Watson (1994) and Stock and Watson (2001).

[2] This type of approach is discussed in a recent paper by Tenreyro and Barro (2007), who argue that currency arrangements such as dollarization can be used as an instrument for the effect of the exchange rate regime on bilateral outcomes between dollarizing countries.

[3] Many European countries followed Germany's lead in setting their monetary policy during our sample period, 1973–1998, making Germany effectively the anchor country (von Hagen & Fratianni, 1990). Theoretical arguments in favor of fixed exchange rates in Europe are reviewed by Giavazzi and Giovannini (1989), among others, and go back to arguments made in the context of the Mundell-Fleming model (Mundell, 1963; Fleming, 1962), and the original work on optimal currency areas of Mundell (1961). Within the European context, several authors stressed the benefits of avoiding the problem of time inconsistency, building on work by Kydland and Prescott (1977) and Barro and Gordon (1983). An elegant recent survey and treatment of the subject is found in Alesina and Barro (2002) within the context of currency unions.

[4] The magnitude of the OLS-IV difference is not diminished by the inclusion of standard controls of the recent history of the system. We interpret this as suggestive evidence that, for the European countries we study, controlling for the recent history of the system does not adequately capture central bankers' information sets.

that this leads our estimates to be lower bounds for the true real effects of monetary policy. Specifically, we show that under mild assumptions the probability limit of the IV estimate is a convex combination of the true parameter and the probability limit of the OLS estimate, with the degree of output correlation as the weighting factor. If output shocks were positively correlated, IV would understate the true consequences of monetary policy interventions, but would be expected to outperform OLS. Since our IV estimates are markedly and significantly more negative than OLS, it follows that if our estimates are biased, we are still understating the real consequences of monetary policy interventions.

Our estimation strategy borrows from a long literature on quasi-experimental identification in labor economics. In contrast to most of that literature, motivated by the macroeconomic setting we explicitly recognize the limitation of our instruments and use the explicit derivation of the remaining bias to gain further insights into the underlying identification problem. On the macroeconomic side, our paper relates to two recent strands of literature. On the one hand, our paper is related to the so-called narrative approach, since we strive to work with a known source of changes in interest rates. On the other hand, our approach can be interpreted a restricted VAR with an alternative way of measuring policy innovations. We make this relationship explicit in the paper, and discuss simple VAR estimates motivated by the alternative identification assumptions explored in the paper.

Our results are also relevant for several literatures not directly concerned with estimating the real effects of monetary policy interventions. First, our empirical results show the effect of an anchor country's interest rate movements on economic outcomes in countries pegging their exchange rate to that of the anchor country. Recent papers have discussed the extent to which international monetary linkages may limit a country's ability to conduct independent monetary policy (Shambaugh, 2004; Obstfeld, Shambaugh, & Taylor, 2004, 2005). Second, our first-stage estimates are closely related to recent estimates of the reaction functions of European central banks in Clarida and Gertler (1997) and Clarida et al. (1998).[5] Third, our results also relate to an extensive literature examining the cost and benefits of fixed exchange rates, particularly in reference to the EMS and EMU.[6]

The remainder of the paper is organized as follows. Section II describes the identification strategy and describes how our approach may be understood as a structural VAR with prior restrictions on the dynamic effects of interest rates on output. In that section, we also describe the dynamic interpretation of our estimate, which is the economic consequence of an episode of tightening. The main empirical results of the paper are presented in section III, and section IV concludes.

[5] Taylor (1993) discusses the concept of an interest rate policy rule. See Woodford (2003) for a comprehensive analysis of optimal interest rules. An early theoretical and empirical assessment of interest rate targeting goes back to Barro (1988).

[6] This literature is summarized by Eichengreen (1990) and Wyplosz (1997), among others.

II. Identification Strategy

An important step in obtaining unbiased estimates of the effect of monetary policy on output growth is the isolation of innovations in monetary policy that are not themselves correlated with the evolution of the economy. This has proven difficult since central banks typically set their interest rate in response to current and expected future evolutions in output growth and inflation. However, central banks may pursue policy goals that are not directly related to output innovations. For example, countries often peg their currency to that of an anchor country to obtain credibility, stabilize financial markets, or reduce inflation. Some central banks choose anchors for their monetary policy in order to detach interventions from output stabilization. In this paper, we argue that alternative goals provide potential estimation strategies for consistent estimation of some aspects of the effects of monetary policy on the real economy.

Suppose the central bank sets monetary policy taking into account expected future inflation and output growth according to the reaction function

$$i_t = \beta_0 + \beta_1 \hat{y}_{t|t-1} + \beta_2 \hat{\pi}_{t+1|t-1} + \beta_3 z_t + v_t, \tag{1}$$

where the interest rate (i_t) is taken to be the central bank's main policy tool, $\hat{y}_{t|t-1} = E[y_t|\Omega_{t-1}]$ and $\hat{\pi}_{t+1|t-1} = E[\pi_{t+1}|\Omega_{t-1}]$ denote the monetary authority's forecast of real output growth and the lead of inflation based on information available as of date $t - 1$ and assuming no change in stance, and v_t is an orthogonal policy disturbance. Such a reaction function has been proposed by Clarida, Galí, and Gertler (2000) based on Taylor (1993), but a forward-looking component of monetary policy is implicit in many classic discussions of monetary policy (such as Bernanke & Blinder, 1992; Bernanke & Mihov, 1998; or Romer & Romer, 1989).[7]

In addition, the central bank may adopt a target z_t for interest rates that is independent of the evolution of domestic output and inflation. Often, central banks peg their exchange rate to that of a leader country, effectively limiting their monetary policy independence in an environment of flexible capital flows. For example, it is well known that European countries followed the interest rate policies of the German Bundesbank. Similarly, many countries in Asia or Latin America have targeted U.S. interest rates at various moments in the last decades.

To develop the point that these external goals may aid in identifying the effect of monetary policy innovations, con-

[7] Equation (1) has also become an integral part of recent theoretical models of monetary policy and the open economy such as Benigno (2004), Engel and West (2006), or Galí and Monacelli (2005).

sider a common regression specification in the literature for a linear relationship between real output growth (y_t) and the interest rate (i_t):

$$y_t = \alpha_0 + \theta i_t + \phi_1' W_{t-1} + u_t, \tag{2}$$

where θ represents the short-run causal effect of interest rates on the real economy, and W_{t-1} may include other variables such as inflation as well as lags of variables in the system. The fundamental identification problem of the effect of monetary policy arises because u_t may be correlated with i_t, leading to a bias in ordinary least squares estimates of equation (2). If the central bank follows the reaction function in equation (1), this is likely to be the case since in addition to true policy innovations short-term interest rates are also set in reaction to current and future trends in output growth.

An important literature addresses this identification problem. A common approach is to impose assumptions on the central bank's reaction function to recover the true underlying policy disturbances v_t. Thereby, researchers often assume a convention on the timing of monetary policy decisions that ensures interest rates only react to past information on the economy. This assures that the ordinary least squares estimator of θ will be consistent since it implies that conditional on W_{t-1}, the interest rate i_t is uncorrelated to the error component; that is, we have that

$$C[u_t, i_t | W_{t-1}] = 0. \tag{3}$$

If the central bank's reaction function is well represented by equation (1), it is thus crucial for the implementation of this approach that the variables at the researcher's disposal are sufficient for the central bank's information set.

This approach has allowed the existing literature to uncover important insights of the relationship of monetary policy and the economy. However, a potential critique is that monetary policy innovations are inferred from a residual without explicit information on the actual stance of monetary policy. We argue that the extended reaction function (1) can lead to an alternative strategy to identification of monetary policy shocks on the economy complementary to the conventional approach.

In particular, if the additional target z_t is uncorrelated with the central bank's expectation of future output or inflation realizations, then it leads to changes in the interest rate that are uncorrelated with the disturbance in equation (2). Instead of equation (3), the orthogonality condition becomes

$$C[u_t, z_t | W_{t-1}] = 0, \tag{4}$$

which is the generic condition for the validity of an IV estimator. The system of equations corresponding to the IV estimate consists of equation (2) and an equation for the interest rate. Using the policy reaction function (1), the first-stage regression can be written as

$$i_t = \beta_0 + \phi_2' W_{t-1} + \beta_3 z_t + \eta_t, \tag{5}$$

where the error η_t is the sum of ν_t and an error reflecting the differences in the information of the researcher and the central bank. If z_t is uncorrelated with u_t, the error in equation (2), then z_t generates quasi-experimental variation in i_t that allows for consistent estimation of the causal effect of nominal interest rates on the economy.

This approach has some key advantages. First, it uses an explicit source of variation to estimate the effect of monetary policy innovations on the economy. This allows for an assessment of the potential bias of estimates of the effects of monetary policy that may arise because of forward-looking monetary policy by comparing our OLS and IV estimates. Second, it allows one to relax the assumption that the central bank only reacts to past output growth. The new estimates are thus consistent under a wider range of assumptions of central-bank behavior.

However, it is also clear that this estimation strategy does not come without costs. First, one has to impose the assumption that the external goal z_t is independent of domestic output innovations and that it has no direct effects on the domestic economy. We will address this concern below at length. Second, the approach limits the ability to analyze the dynamic effects of monetary policy. Nonetheless, we show below that the estimated parameter still has an economically meaningful interpretation. Since most of the literature has estimated the effect of interest rates on the economy using VARs, we will also discuss under what circumstances the main idea of the paper—the use of *observable* policy innovations to estimate the effect of monetary policy on the economy—can also be implemented in a VAR context.

The identification strategy we pursue may be understood as a highly parsimonious structural VAR with a focus on identification of a single parameter. Suppressing intercepts, a VAR comparable to the two-equation system used here can be written as

$$\begin{pmatrix} 1 & -\theta \\ 0 & 1 \end{pmatrix} \begin{pmatrix} y_t \\ i_t \end{pmatrix} = \begin{pmatrix} \phi_1 \\ \phi_2 \end{pmatrix} W_{t-1} + \begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix}.$$

Identification within this system of equations is usually specified in terms of the matrix on the contemporaneous correlations alone, leaving the structure of lags completely flexible.[8] A common identifying assumption is to exclude contemporaneous feedback of output growth on the interest rate (or of interest rates on output, which is equivalent in statistical terms), setting to 0 the lower-left-hand parameter of the matrix of contemporaneous correlations. This identi-

[8] In the earlier literature on VARs, identification of the effects was mainly based on the ordering of variables and a factorization of the error matrix to achieve a recursive system. In the case of structural VARs, restrictions on the matrix of contemporaneous correlations are determined by economic theories; the resulting system can but need not be recursive; see, for example, Bernanke (1986) or Blanchard and Watson (1986).

fication strategy is equivalent to imposing the assumption in equation (3) conditional on all of the lags of the system.[9]

Augmenting the above system by an equation for the German interest rate and freeing up the zero restriction yields

$$\begin{pmatrix} 1 & -\theta & 0 \\ \lambda & 1 & \beta \\ 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} y_t \\ i_t \\ z_t \end{pmatrix} = \begin{pmatrix} \phi_1 \\ \phi_2 \\ \phi_3 \end{pmatrix} W_{t-1} + \begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \\ \varepsilon_{3t} \end{pmatrix},$$

where now identification of θ can rely on the zero restriction in the upper-right corner of the matrix of contemporaneous correlations. In this system of equations the restriction $\lambda = 0$ is not needed for identification.[10]

The feasibility of our approach hinges on the assumption that z_t is independent of output innovations, and that its only effect is through the interest rate. The validity of these assumptions is best discussed in the specific context of our empirical implementation, the subject to which we now turn.

We exploit the fact that the German central bank was the leader for monetary policy for many European countries indirectly since the breakdown of Bretton Woods, and directly since the conception of the European Monetary System (EMS) in 1979. A long literature suggests that German interest rates have a strong impact on interest rates of other European countries.[11] Clarida et al. (1998) and Clarida and Gertler (1997) describe how the Bundesbank's reaction function is similar to that of the Federal Reserve. On the other hand, Clarida et al. (1998) show that the German interest rate plays an important role in the reaction function of France, Italy, and the United Kingdom.[12] Based on these considerations, we use the German interest rate as an instrument for the nominal rate of other European countries and provide explicit estimates of relationship (5). That countries constrain the scope of their domestic monetary policy also receives empirical support from recent work by Shambaugh (2004) and Obstfeld et al. (2004, 2005), who show that the comovement of interest rate changes is higher under pegged exchange rates than under floating rates.[13]

European economies are closely linked by trade flows and financial markets. This leads to two potential concerns with our instrumental variable strategy.

First, the German interest rate may have a direct effect on the domestic economy of the follower country beyond its impulse running through the domestic interest rate. This is more likely to be a concern for smaller countries that are more dependent on trade from Germany and may be hit by a contraction of German demand in response to a rise in German interest rates. Since these are the same countries for which correlated shocks are more likely, we will repeat our estimation strategy for a sample of countries with high and low importance of trade with Germany relative to domestic GDP. By the foregoing arguments, the remaining bias of IV estimates should be small for larger countries whose economies are less integrated with Germany. Note that these are the same countries that we expect to have residual autonomy in making monetary policy choices. Thus, we would expect the correlation of interest rates to be lower and the difference between OLS and IV to be larger for these countries.

Second, this implies output and inflation innovations are likely to be correlated across countries (Frankel & Rose, 1998). Such correlation will lead IV estimates to have a remaining bias. In our empirical application, we will include lags of domestic output growth and inflation to absorb sources of comovement in interest rates due to economic factors.

We next develop a simple framework for understanding the impact of correlated output shocks on the quality of IV estimates. We show that under general conditions and for a broad range of parameters, the probability limit of IV is closer to the target parameter than that of OLS. To substantiate this point, we first restate without covariates the two key equations from the simultaneous-equations model outlined above, for a representative "home country" (such as, France or the United Kingdom):

$$y_t = \alpha_0 + \theta i_t + u_t, \tag{6}$$

$$i_t = \beta_0 + \beta_3 z_t + \eta_t. \tag{7}$$

Next, we introduce two new equations. The first equation is a model for the effect of German interest rates on German output growth, analogous to equation (6):

$$y_t^* = \alpha_0^* + \theta^* z_t + u_t^*, \tag{8}$$

where asterisks represent German variables and parameters. The second equation is a population conditional linear projection of home-country output shocks on German output shocks,

[9] In standard VAR analysis, this assumption means that the reduced form of the system may be estimated consistently by least squares. The structural disturbances are then obtained from the reduced-form residuals by method-of-moments techniques. These in conjunction with the parameter estimates of the lag structure are then used for further analysis.

[10] Hamilton (1994) discusses a system such as that represented above.

[11] As noted above, estimates of the degree of leadership differ in the literature (for example, Giavazzi & Giovannini, 1987, or von Hagen & Fratianni, 1990); this is a substantively important point, but does not compromise our methodology.

[12] The authors do not analyze the role of European exchange rates in the Bundesbank's reaction function, nor do they explicitly compare the role of exchange rates versus interest rates in the other countries' functions.

[13] Their approach is similar to ours in that they also try to estimate the correlation between countries' interest rates to that of an anchor country. Invoking uncovered interest-rate parity, these authors argue that estimating interest relationships in levels is inappropriate if interest rates of the anchor country are highly persistent. This is less likely to be a problem in our application, since for part of the period capital controls were in place and the time horizon we consider is relatively short. Additionally, we argue below that our approach may be understood as uncovering a cointegrated relationship.

$$u_t = \delta u_t^* + \omega_t, \tag{9}$$

where we assume that ω_t and z_t are uncorrelated. Equation (9) amounts to a model for the bias of IV. The model allows for the invalidity of z_t, with δ quantifying the magnitude of the bias. For example, when $\delta = 0$, IV is consistent, and when $\delta > 0$, IV will typically be biased toward OLS. Equations (6)–(9) allow us to characterize the departure of the IV and OLS probability limit from their target parameter θ.

To see how, first note that the model in equations (6) and (7) implies that the home-country OLS bias is

$$\begin{aligned} B_{OLS} &\equiv \text{plim}\hat{\theta}_{OLS} - \theta = \frac{C[i_t, u_t]}{V[i_t]} \\ &= \beta_3 \frac{C[z_t, u_t]}{V[i_t]} + \frac{C[\eta_t, u_t]}{V[i_t]} \\ &= \beta_3 \frac{C[z_t, i_t]}{V[i_t]} B_{IV} + \frac{C[\eta_t, u_t]}{V[i_t]} \\ &= \frac{F}{F+T-2} B_{IV} + \frac{C[\eta_t, u_t]}{V[i_t]} \\ &= R^2 B_{IV} + \frac{C[\eta_t, u_t]}{V[i_t]}, \end{aligned} \tag{10}$$

where T is the total sample size, F is the population F-statistic on the exclusion of z_t from equation (7), and R^2 is the regression population R^2 from equation (7).[14] Note that the F-statistic in question in the display is the F-statistic that assumes i.i.d. data. Second, note that equations (8) and (9) imply the home-country IV bias is

$$\begin{aligned} B_{IV} &\equiv \text{plim}\hat{\theta}_{IV} - \theta = \frac{C[z_t, u_t]}{C[z_t, i_t]} = \delta \frac{C[z_t, u_t^*]}{C[z_t, i_t]} \\ &= \delta \frac{C[z_t, u_t^*]}{V[z_t]} \frac{V[z_t]}{C[z_t, i_t]} = \frac{\delta}{\beta_3} B_{OLS}^*, \end{aligned} \tag{11}$$

where $B_{OLS}^* \equiv C[z_t, Y^*_t]/V[z_t] - \theta^*$ is the OLS bias for Germany. Third, assume that

$$B_{OLS}^* = \frac{C[\eta_t, u_t]}{V[i_t]}. \tag{A0}$$

In words, this means we are assuming that the OLS bias for Germany would be equal to the OLS bias for the home country, if the home country did not follow Germany's monetary policy.

Equations (10) and (11), combined with assumption (A0), allow us to characterize the settings in which IV outperforms OLS in bias terms:

$$\begin{aligned} B_{IV} < B_{OLS} &\Leftrightarrow B_{IV} < R^2 B_{IV} + \frac{C[\eta_t, u_t]}{V[i_t]} \\ &\Leftrightarrow B_{IV}(1 - R^2) < B_{OLS}^* \\ &\Leftrightarrow \frac{\delta}{\beta_3} B_{OLS}^*(1 - R^2) < B_{OLS}^* \\ &\Leftrightarrow \delta < \beta_3 \frac{1}{1 - R^2} = \beta_3 \frac{F + T - 2}{T - 2}. \end{aligned} \tag{12}$$

This inequality states that IV will be less biased than OLS whenever output shocks covary less than $\beta_3/(1 - R^2)$. Note that $\beta_3/(1 - R^2)$ measures the strength of the first-stage relationship and can be calculated empirically. In our data, $\beta_3/(1 - R^2)$ with no controls is 1.41 on average, ranging from 0.42 to 2.67. We find it implausible that output shocks would be so highly correlated. Thus, while the structure we have outlined here is restrictive, we view these calculations as strongly suggesting that IV should outperform OLS in bias terms in this application.

A similar expression can be developed, under slightly different assumptions, that is appropriate for regressions with covariates. Rewrite

$$y_t = \alpha_0 + \theta i_t + \phi_1' W_{t-1} + u_t, \tag{6$'$}$$

$$i_t = \beta_0 + \beta_3 z_t + \phi_2' W_{t-1} + \eta_t, \tag{7$'$}$$

$$y_t^* = \alpha_0^* + \theta^* z_t + \phi_1^{*\prime} W_{t-1} + u_t^*, \tag{8$'$}$$

$$u_t = \delta u_t^* + \phi_3' W_{t-1} + \omega_t, \tag{9$'$}$$

where we now assume that ω_t and z_t are uncorrelated conditional on W_{t-1}. These equations imply a slightly restated version of equation (10):

$$\begin{aligned} B_{OLS} &= \frac{C[i_t, u_t | W_{t-1}]}{V[i_t | W_{t-1}]} \\ &= \beta_3 \frac{C[z_t, u_t | W_{t-1}]}{V[i_t | W_{t-1}]} + \frac{C[\eta_t, u_t | W_{t-1}]}{V[i_t | W_{t-1}]} \\ &= \beta_3 \frac{C[z_t, i_t | W_{t-1}]}{V[i_t | W_{t-1}]} B_{IV} + \frac{C[\eta_t, u_t | W_{t-1}]}{V[i_t | W_{t-1}]} \\ &= \frac{F}{F + T - k} B_{IV} + \frac{C[\eta_t, u_t | W_{t-1}]}{V[i_t | W_{t-1}]} \\ &= \frac{R_U^2 - R_R^2}{1 - R_R^2} B_{IV} + \frac{C[\eta_t, u_t | W_{t-1}]}{V[i_t | W_{t-1}]}, \end{aligned} \tag{10$'$}$$

where F is the population F-statistic for the exclusion of z_t, k is the number of total regressors, $R_U{}^2$ is the population R^2 for the full regression, and $R_R{}^2$ is the population R^2 for the "short" regression that excludes z_t, with each concept pertaining to the regression model in equation (7$'$). As for the

[14] Equation (10) can be derived directly, but also follows from considering a first-order version of some of the equations developed in Hausman and Hahn (2003).

unconditional case, the relevant F-statistic is the F-statistic that assumes i.i.d. data.

Equations (6′) through (9′) also imply that equation (11) holds, with each concept being conditional on W_{t-1}, rather than being unconditional. Finally, suppose that (A0) holds conditional on W_{t-1}.

Under these conditions, the covariate-adjusted analog to equation (12) is

$$B_{IV} < B_{OLS} \Leftrightarrow \beta_3 \frac{1 - R_R^2}{1 - R_U^2} = \beta_3 \frac{F + T - k}{T - k}. \qquad (12')$$

In our data with controls (four lags of GDP and inflation), $\beta_3(1 - R_R^2)/(1 - R_U^2)$ is 1.41 on average, ranging from 0.46 to 2.38.

In our empirical analysis, we will use equations (12) and (12′) to relate the relative bias to countries' macroeconomic relationships with Germany. For example, if a country is heavily dependent on trade with Germany (relative to its GDP), then shocks that hit Germany will be directly transmitted to the domestic economy as German supply and demand for goods adjust. In this case, forward-looking monetary choices by the Bundesbank will be correlated with a country's GDP growth, making it more difficult to differentiate between the OLS and IV estimates. Factors governing the degree of a country's monetary independence also determine the relative bias between IV and OLS. For example, the wider exchange rate bands are in a target zone, the more domestic interest rates can temporarily deviate from those of the anchor country. If larger "effective" exchange rate bands imply higher exchange rate volatility, we expect the size of the OLS-IV difference (IV estimate) to be positively related to volatility.

A possible drawback of our approach is that θ is a measure of the short-run causal effect of a change in interest rates on economic growth. Typically, W_{t-1} contains several lags of the interest rate, and researchers have been interested in the entire dynamic path of the effect of interest rate shocks.

However, we can demonstrate the probability limit of our reduced-form parameter under a dynamic data-generating process (DGP). Suppose that in place of equation (2) the DGP is

$$y_t = \alpha_0 + \theta_0 i_t + \theta_1 i_{t-1} + \cdots + \theta_p i_{t-p} + \phi_1' W_{t-1} + u_t.$$

The probability limits are

$$\hat{\theta}_{OLS} \stackrel{p}{=} \theta_0 + \theta_1\gamma_1 + \theta_2\gamma_2 + \cdots + \theta_p\gamma_p + \frac{\beta_3 C[z_t, u_t|W_{t-1}] + C[\eta_t, u_t|W_{t-1}]}{V[i_t|W_{t-1}]}, \qquad (13)$$

$$\hat{\theta}_{IV} \stackrel{p}{=} \theta_0 + \theta_1\gamma_1^{IV} + \theta_2\gamma_2^{IV} + \cdots + \theta_p\gamma_p^{IV} + \frac{C[z_t, u_t|W_{t-1}]}{C[z_t, i_t|W_{t-1}]}, \qquad (14)$$

where $\gamma_j = C[i_t, i_{t-j}|W_{t-1}]/V[i_t|W_{t-1}]$, j = 1, 2, . . . , p are the autocovariances of interest rates, and $\gamma_j^{IV} = C[z_t, i_{t-j}]|W_{t-1}]/C[z_t, i_t|W_{t-1}]$ are the instrumental variable analogs.

Consider briefly the interpretation of the summary parameter $\theta \equiv \theta_0 + \theta_1\gamma_1 + \cdots + \theta_p\gamma_p$. The parameter summarizes (i) the instantaneous effect of monetary policy on the real economy, θ_0, and (ii) the historical effect of monetary policy on the real economy, or θ_j for $j = 1, 2, \ldots p$. The weight γ_j applied to the historical influence of monetary policy has a natural interpretation—it measures the extent to which a monetary tightening in period t predicts that monetary policy was tight in period $t - j$. In short, the summary parameter θ measures the general effect of an episode of tight monetary policy of a given magnitude. Thus, while our approach does not allow us to trace out the entire dynamic effect of monetary policy on the real economy, it does allow us to identify a parameter of interest to policymakers.

Importantly, the implicit OLS and IV weighting functions are (under a mild assumption) equal—that is, $\gamma_j = \gamma_j^{IV}$. This follows immediately from a few lines of algebra. Linearly project the German interest rate onto the national interest rate for period t, and plug these linear projections into the definition of γ_j^{IV}. The "mild assumption" mentioned holds that the residual from this projection is orthogonal to lagged home-country interest rates, which we view as innocuous since the projection residual is by definition orthogonal to current home-country interest rates.

The intuition for this result is straightforward and is based on the omitted variable bias formula. Essentially, our estimates of the effect of interest rates on output will reflect not just the effect of the current period's interest rate, but additionally the effect of past interest rates, since these are correlated with current period interest rates. Our estimates represent, then, not the effect of a one-period interest rate increase, but the reduced-form effect of an episode of interest rate increases.

Interest rates are too persistent to allow an estimation of a fully dynamic version of our instrumental variable estimates. Essentially, current and lagged foreign interest rates do not provide enough distinct variation to function as two separate instruments. Nevertheless, we can follow common practice and estimate the direct effect of German interest rates on domestic output growth over time using a vector autoregression framework. This approach fully exploits the additional information we bring to bear—the existence of external policy goals of the central bank—without restricting the lag structure of the model. Thus, we essentially estimate the "reduced form" version of our instrumental variable model in which *only* the foreign interest rate is allowed to enter the equation for output growth. The parameters of this model yield true causal effects by a simple extension of assumption (4) to the dynamic context.

An added advantage of this approach is that it does not yet impose the exclusion restriction that the foreign interest rates have no direct effects on domestic output growth. If one is willing to impose this assumption, the causal effect is approximately a scaled version of the reduced-form coefficients. If one believes there should be a direct effect of German interest rates, as may be the case for smaller countries, the VAR estimates capture the dynamic effect of an external shock to the domestic economy.[15]

III. Data and Empirical Results

A. *Data and Empirical Implementation*

We estimate OLS and IV regressions of the impact of nominal short-term interest rates on real output growth for twelve European countries using quarterly data from 1973 to 1998. These countries are chosen given data availability and include but are not limited to most participants in the European Monetary System (EMS).[16] Nominal GDP data are taken from the International Monetary Fund's International Financial Statistics (IFS) database and are deflated by each country's real GDP deflator (1995 = 100, also from the IFS database). To control for seasonal components we include quarterly dummies in all specifications.[17] The short-term interest rate by which we measure monetary policy is the overnight lending or call money rate from the Global Financial Database. We average end-of-month rates quarterly.[18] We also have tried using the central bank's discount rate, and the three-month T-bill rate (annualized). Our results are generally robust to the choice of interest rates used.

The main estimation equations are (2) and (5), where the level of the quarterly German overnight rate is used as an instrument for the level of the call money rate in the other European countries. It is widely accepted in the literature that the German central bank became the effective trendsetter in the stance of monetary policy for other European countries since the breakdown of the Bretton Woods system. However, while German monetary policy seems to have been a strong influence on countries' interest rates, this did not negate forward-looking behavior on the part of the monetary policy, particularly for larger countries within the EMS, and those who joined late or had wider exchange rate bands (see, for example, von Hagen & Fratianni, 1990). For these countries, we expect IV to yield more negative estimates than OLS. We also expect the remaining bias due to common output shocks to be small, raising our confidence in the IV estimates.

However, for the smaller, open countries pegged exchange rates and flexible capital markets may have left little scope for independent monetary policy.[19] On the one hand, this implies a small bias of OLS. On the other hand, many smaller countries had strong trade linkages to Germany. This leads to a remaining bias in the IV estimate through a higher correlation of shocks to output growth. Similarly, a high degree of integration is likely to lead to direct effects of German interest rate shocks on the domestic economy. Thus, IV and OLS should be more similar, and both may be hard to interpret.

Based on these considerations, we begin by presenting the results of estimating OLS and IV models for single countries and briefly discuss summary measures based on pooled regressions. We then present results from our pooled models estimated separately for samples of countries with high and low trade integration with Germany and discuss the empirical implications of our bias calculations. Last, we briefly summarize the results from a series of simple VARs.

B. *Main Empirical Results*

Basic OLS estimates of the effect of monetary policy are shown in columns 1 and 5 of table 1. Taken at face value, these estimates imply that a 1 percentage point increase in the interest rate lowers quarterly real growth only moderately: 0.094 percentage points in the Netherlands and only 0.015 percentage points in France. The average effect across countries is −0.043, and the median is −0.039. All tables report two sets of standard errors; usual heterokedasticity-robust Eicher-White standard errors are in parentheses, and Newey-West standard errors correcting for fourth-order serial correlation are in squared brackets. The two sets of standard errors are quite similar, and the choice of standard error affects our results only for very few cases. Neither seems to be overall more conservative, so we chose to report both.[20]

[15] Note that if one is willing to impose the OLS assumption in equation (3), the direct effect of both the foreign and the domestic interest rate can be estimated in the framework of a recursive VAR. However, in the case of forward-looking monetary policy and under the alternative set of assumptions we explore here, this is not feasible. In particular, if i_t is endogenous even conditional on W_{t-1}, then the coefficients on both i_t and z_t will be biased and cannot be interpreted. This is a standard problem in simultaneous-equation systems.

[16] The countries are Austria, Belgium, France, Great Britain, France, Italy, the Netherlands, Norway, Portugal, Spain, Sweden, and Switzerland. Notable exceptions due to data limitations on quarterly nominal interest rates are Denmark and Ireland.

[17] We lack complete data for quarterly GDP for Belgium, Italy, the Netherlands, Portugal, and Sweden in the 1970s. Data are missing from 1973Q1–1980Q2 (Belgium), 1973Q1–1977Q1 (Netherlands), 1973Q1–1977Q1 (Portugal), and 1973Q1–1980Q1 (Sweden). For Portugal we are also missing interest rate data from 1973Q1–1975Q3.

[18] Overnight/call money rates are missing for two countries: 1973Q1–1978Q2 (Italy) and 1973Q1–1975Q3 (Portugal).

[19] The existence of flexible capital markets was not always the case during the EMS period. As Giavazzi and Giovannini (1989) point out, the use of capital controls was predominant in many of the "weaker" currency countries. Paradoxically, Giavazzi and Giovannini find that though these controls had a tendency to break the link between interest rates (as measured by the differential in movements of onshore and offshore rates), they could not reject France's and Italy's monetary policies from being different from Germany's during the period.

[20] As suggested by a referee, recent research by Kiefer and Vogelsan (2005) suggests that robust standard errors may overstate the degree of precision. To assess this possibility, we also compared these robust standard errors with the basic OLS standard errors to make sure that the former are more conservative. This is indeed the case for countries with high precision. Moreover, for these countries confidence levels are small

TABLE 1.—THE EFFECT OF INTEREST RATES ON THE REAL ECONOMY: OLS, IV, AND FIRST STAGE

	No Controls				Four Lags of Inflation and Growth			
Country	(1) OLS	(2) IV	(3) OLS − IV	(4) FS	(5) OLS	(6) IV	(7) OLS − IV	(8) FS
Germany	−0.071				−0.062			
	(0.043)				(0.048)			
	[0.042]				[0.050]			
Great Britain	−0.058	−0.197	0.139	0.750	−0.059	−0.179	0.120	1.147
	(0.031)	(0.055)	(0.046)	(0.126)	(0.025)	(0.043)	(0.035)	(0.153)
	[0.036]	[0.069]	[0.060]	[0.210]	[0.024]	[0.052]	[0.047]	[0.218]
France	−0.015	−0.074	0.059	1.264	−0.019	−0.072	0.053	1.113
	(0.015)	(0.035)	(0.032)	(0.247)	(0.020)	(0.035)	(0.029)	(0.215)
	[0.016]	[0.042]	[0.038]	[0.315]	[0.026]	[0.032]	[0.019]	[0.204]
Italy	−0.024	−0.129	0.105	0.950	−0.034	−0.168	0.134	0.583
	(0.018)	(0.047)	(0.043)	(0.199)	(0.023)	(0.062)	(0.057)	(0.165)
	[0.020]	[0.070]	[0.067]	[0.316]	[0.021]	[0.059]	[0.056]	[0.187]
Spain	−0.015	−0.180	0.165	0.488	0.002	−0.063	0.066	0.441
	(0.011)	(0.078)	(0.077)	(0.194)	(0.005)	(0.043)	(0.042)	(0.183)
	[0.015]	[0.131]	[0.131]	[0.338]	[0.004]	[0.053]	[0.053]	[0.263]
Netherlands	−0.094	−0.145	0.051	0.870	−0.087	−0.140	0.052	0.793
	(0.034)	(0.050)	(0.036)	(0.085)	(0.038)	(0.060)	(0.046)	(0.105)
	[0.029]	[0.040]	[0.028]	[0.110]	[0.038]	[0.048]	[0.030]	[0.146]
Switzerland	−0.016	−0.130	0.114	0.559	−0.015	−0.060	0.044	0.702
	(0.041)	(0.089)	(0.079)	(0.122)	(0.040)	(0.076)	(0.065)	(0.159)
	[0.046]	[0.088]	[0.074]	[0.243]	[0.043]	[0.058]	[0.039]	[0.279]
Sweden	−0.055	−0.061	0.006	1.033	−0.085	−0.112	0.027	0.707
	(0.034)	(0.047)	(0.033)	(0.075)	(0.060)	(0.076)	(0.046)	(0.064)
	[0.033]	[0.042]	[0.026]	[0.104]	[0.054]	[0.071]	[0.047]	[0.088]
Belgium	−0.031	−0.025	−0.006	1.105	−0.118	−0.174	0.056	0.910
	(0.063)	(0.131)	(0.114)	(0.200)	(0.058)	(0.136)	(0.123)	(0.237)
	[0.040]	[0.079]	[0.069]	[0.254]	[0.059]	[0.141]	[0.128]	[0.312]
Austria	−0.069	−0.065	−0.004	0.838	−0.097	−0.072	−0.025	0.787
	(0.085)	(0.090)	(0.028)	(0.060)	(0.070)	(0.090)	(0.057)	(0.063)
	[0.055]	[0.058]	[0.018]	[0.089]	[0.074]	[0.084]	[0.039]	[0.093]
Norway	−0.047	−0.175	0.128	0.596	−0.093	−0.435	0.342	0.495
	(0.066)	(0.196)	(0.184)	(0.177)	(0.075)	(0.286)	(0.276)	(0.209)
	[0.050]	[0.162]	[0.154]	[0.270]	[0.074]	[0.290]	[0.281]	[0.304]
Portugal	−0.025	−0.068	0.044	1.143	−0.047	−0.104	0.057	0.912
	(0.032)	(0.062)	(0.054)	(0.191)	(0.039)	(0.083)	(0.073)	(0.146)
	[0.024]	[0.042]	[0.035]	[0.345]	[0.039]	[0.059]	[0.044]	[0.211]
Average coefficient	**−0.043**	**−0.114**	**0.073**	**0.872**	**−0.059**	**−0.144**	**0.084**	**0.781**
Median coefficient	**−0.039**	**−0.129**	**0.059**	**0.870**	**−0.060**	**−0.112**	**0.056**	**0.787**
Standard deviation	**0.026**	**0.057**	**0.061**	**0.255**	**0.038**	**0.107**	**0.095**	**0.230**

Notes: Table gives OLS and IV estimates of the effect of nominal interest rates on quarterly real economic growth. OLS estimates in columns 1 and 5 include four season indicators ("no controls"), and four season indicators as well as four lags each of inflation and real economic growth, respectively. IV estimates in columns 2 and 6 use the same controls, but instrument home-country interest rates with German interest rates. OLS-IV difference in columns 3 and 7 give the simple difference between the OLS and IV estimates. First-stage coefficient of German interest rate in columns 4 and 8. Standard errors in parentheses are Huber-Eicher-White standard errors and are robust to heteroskedasticity. Standard errors in square brackets are fourth-order Newey-West standard errors and are robust to fourth-order autocorrelation.

To summarize the basic relationship across countries, and help to assess the impact of different specifications on the overall effect of monetary policy, we also pool our results using several alternative variables as weights.[21] Table 2 shows pooled estimates in which countries are equally weighted and weighted by 2003 GDP in U.S. dollars.[22] In calculating the pooled estimates, we restrict the first-stage and reduced-form coefficients to be equal across countries for computational reasons. We did allow for country fixed effects and separate coefficients on the lags of the system.[23]

enough that they would be significant even under liberal upward adjustment of standard errors. For countries with low precision, OLS standard errors tend to be similar; for these countries, inferences have to be made with care. However, none of the main conclusions of the paper are affected.

[21] We do not view them as an estimate of a common underlying parameter, but rather as a summary measure of the individual coefficients. In the case of fixed country-specific weights, one can show that the pooled estimates are a weighted function of the country-specific coefficients (with weights proportional to the fixed country weight in the pooled model).

[22] An earlier working-paper version (di Giovanni, McCrary, & von Wachter, 2005) also used the fraction GDP not due to trade and the volatility of the exchange rate vis-à-vis the German mark as weights, with little difference in results.

[23] It would seem sensible to allow country-specific first-stage coefficients to reflect differences in the underlying mechanism across countries. However, doing so we face a problem of multiple weak instruments very similar to that faced by Angrist and Krueger (1991), who also interact their instrument with state dummies. As discussed in the ensuing literature on

TABLE 2.—SUMMARIZING OLS, IV, AND FIRST-STAGE ESTIMATES

	Unweighted				Weighted by 2003 GDP US$			
Controls	(1) OLS	(2) IV	(3) OLS − IV	(4) FS	(5) OLS	(6) IV	(7) OLS − IV	(8) FS
(A) No controls	−0.033	−0.107	0.075	0.854	−0.030	−0.121	0.091	0.899
	(0.012)	(0.025)	(0.021)	(0.046)	(0.009)	(0.018)	(0.015)	(0.048)
	[0.010]	[0.021]	[0.019]	[0.085]	[0.011]	[0.027]	[0.025]	[0.125]
(B) One lag of growth and inflation	−0.038	−0.135	0.097	0.789	−0.027	−0.138	0.111	0.834
	(0.013)	(0.027)	(0.024)	(0.045)	(0.009)	(0.020)	(0.018)	(0.047)
	[0.014]	[0.025]	[0.021]	[0.079]	[0.013]	[0.032]	[0.029]	[0.113]
(C) Four lags of growth and inflation	−0.035	−0.152	0.117	0.710	−0.020	−0.135	0.115	0.802
	(0.013)	(0.030)	(0.027)	(0.044)	(0.009)	(0.021)	(0.019)	(0.047)
	[0.016]	[0.030]	[0.026]	[0.082]	[0.013]	[0.032]	[0.030]	[0.116]
(D) One lag of growth and inflation, different for each country	−0.044	−0.139	0.095	0.642	−0.035	−0.130	0.095	0.745
	(0.014)	(0.037)	(0.034)	(0.052)	(0.010)	(0.023)	(0.020)	(0.051)
	[0.017]	[0.040]	[0.036]	[0.082]	[0.015]	[0.032]	[0.028]	[0.115]
(E) Four lags of growth and inflation, different for each country	−0.045	−0.137	0.092	0.647	−0.036	−0.129	0.093	0.749
	(0.015)	(0.037)	(0.034)	(0.051)	(0.010)	(0.022)	(0.020)	(0.050)
	[0.017]	[0.040]	[0.036]	[0.081]	[0.016]	[0.031]	[0.027]	[0.113]
Average coefficient	**−0.039**	**−0.134**	**0.095**	**0.729**	**−0.030**	**−0.131**	**0.101**	**0.806**
Median coefficient	**−0.038**	**−0.137**	**0.095**	**0.710**	**−0.030**	**−0.130**	**0.095**	**0.802**
Standard deviation	**0.006**	**0.016**	**0.015**	**0.092**	**0.006**	**0.007**	**0.011**	**0.064**

Notes: Table gives pooled OLS and IV, OLS-IV, and first-stage estimates of the effect of nominal interest rates on quarterly real economic growth for all countries except Germany. Estimation includes the control variables specified under controls. Each estimate includes season indicators fully interacted with country indicators. In rows (D) and (E), lags are chosen separately for each country using significance levels. Columns 1–4 give estimates that are equally weighted. Columns 5–8 give estimates based on weights that are proportional to a country's 2003 level of GDP in U.S. dollars. The last three rows report the mean, median, and standard deviation of the coefficient estimates in rows (A) to (E). Standard errors in parentheses are Huber-Eicher-White standard errors and are robust to heteroskedasticity. Standard errors in square brackets are fourth-order Newey-West standard errors and are robust to fourth-order autocorrelation.

The average effect for the pooled OLS regressions without weights in the bottom of table 2 is −0.039 percentage points; weighting by 2003 GDP makes very little difference.

The corresponding estimates using the German interest rate as an instrument for the national interest rate are shown in columns 2 and 6 of table 1. For all countries (except Austria and Belgium), the IV estimates are more negative than the OLS estimates. This suggests that some degree of endogeneity with respect to real output growth affects most countries' interest rates. A simple interpretation of this endogeneity is that it is capturing the extent to which the monetary authority is forward looking. The pooled IV estimates in table 2 summarize this result: the IV estimate suggests that a 1 percentage point increase in interest rates (on average) causes a reduction in real output growth of 0.134 percentage points (unweighted), which exceeds the OLS estimate by a factor of three. The differences between OLS and IV are always statistically significant in the pooled models. For single countries, the difference between OLS and IV is shown in columns 3 and 7 of table 1. It is generally significant and larger for bigger countries (such as, Great Britain, France, Italy, and Spain), as expected and further discussed below.[24]

An important point that arises from the results in tables 1 and 2 is that the covariates we include do not seem to be able to capture the effects of forward-looking behavior, or substantially reduce other sources of bias in OLS estimates. In particular, if covariates were able to control for the bias arising from forward-looking monetary policy, we would have expected that OLS becomes more negative, and that the difference between OLS and IV declines. Our results suggest the opposite.

The IV estimates are based on a strong and significant "first-stage" relationship between national and German interest rates underlying the IV estimates (table 1, columns 4 and 8). This is the fundamental relationship providing us with quasi-experimental variation in interest rates. Most countries have a first-stage coefficient of at least 0.8. However, several countries, including Great Britain, Spain, and Switzerland, have first-stage coefficients on the German interest rate significantly below unity. Thus, it does not appear that our first-stage relationship is systematically biased towards unity.[25] Not surprisingly, some of the coun-

weak instruments, this risks "overfitting" the first-stage relationship and biases IV results toward OLS. However, our pooled estimates are remarkably similar to the sum of the separate estimates weighted by the inverse of their variances (the optimal method-of-moments estimator under the hypothesis of a common coefficient), suggesting to us that this limitation may not be severe.

[24] The standard errors in columns 3 and 7 of tables 1, 2, and 3 are computed as the square root of the differences in variance of IV and OLS estimates. In the case of heteroskedasticity-robust or Newey-West standard errors, this is only an approximation, since the covariance of the coefficients is then only approximately equal to the differences in the variances. The difference between OLS and IV for the pooled estimates is shown in table 2. Not surprisingly, the pooled estimates are more precise, and confirm a strong and significant difference between OLS and IV.

[25] Given the range of estimated coefficients, some significantly below unity, the limited time range, and the partial presence of capital controls during the period of study, we do not believe we are subject to the critique raised by Shambaugh (2004) discussed in section II. However, we ran

tries with low first-stage coefficients either were never part of the EMS or joined late. To directly assess the affects of changes in specifications, table 2 summarizes a variety of different specifications for first-stage regression models pooling all countries with alternative weights in columns 4 and 8. The largest pooled estimate is 0.899, the smallest is 0.642, and the average first-stage coefficient is 0.729 (unweighted) and 0.806 (weighted). We conclude that German monetary policy appears to be a strong and robust determinant of interest rates for the countries included in our sample.

C. *High- and Low-Trade Countries and Remaining Bias*

We have argued above that the IV estimates are more reliable in case of larger countries not dependent as much on trade with Germany. For these countries, we expect a lower amount of correlated shocks and less of a direct effect of German interest rates. Moreover, we would expect a greater scope of independent monetary policy. This implies a greater bias of OLS toward 0, a lower first-stage coefficient, and a bigger difference between OLS and IV.

These predictions are borne out in table 3, which shows pooled estimates separately for countries with trade with Germany as a fraction of GDP above and below the sample median. The list of countries in each group are presented in the footnote of the table. If one considers the average coefficients printed in bold, it is apparent that the OLS estimates are lower but IV estimates are larger in the low- versus the high-trade sample. This leads to large differences in the gap between IV and OLS estimates. The first stage is also slightly higher in the high-trade group (if one weights by GDP, this difference increases since larger countries have more independent monetary policy).[26] These results support the hypothesis that the instrument is valid for larger countries, but may be harder to interpret for smaller countries.[27] The lower panel replicates similar but muted results for a sample split by output correlation.[28]

To further summarize the differences in IV estimates across countries we explore the relationship between the IV estimates and proxies for the approximate bias (cf. equations [12] and [12′]). A simple way to represent the relationship between these estimates and the relevant fundamentals suggested by equations (12) and (12′) is shown in figure 1, which is based on the results from table 1. Figure 1A plots the relationship between the IV estimates and the fraction of GDP due to trade with Germany. As predicted, the IV estimates become less negative the more important a country's trade with Germany is relative to its total output. Figure 1B shows how IV estimates are more negative for countries whose currencies were more volatile, viz. the German mark. This result confirms the intuition that a more flexible exchange rate regime allowed countries more monetary independence. Hence, the use of the German rate as an instrument picks up more exogenous monetary shocks in the domestic country.

The differences across countries carry over to the gap between OLS and IV estimates. The OLS-IV differences, shown in columns 3 and 7 of table 1, are positive and greater for larger countries. Based on the foregoing, we would also expect them to be larger for countries that are less dependent on trade with Germany and have a more volatile bilateral exchange rate. This is shown in figures 1C and 1D. Although the cross-country heterogeneity in the OLS-IV difference is greater than that of the IV estimates, the correlations are as expected. The difference is (i) decreasing with the trade to GDP ratio (figure 1C), and (ii) increasing with exchange rate volatility (figure 1D). These correlations confirm the predictions of our simple representation of monetary policy decisions summarized in equations (12) and (12′), and suggest the gap between IV and OLS reflects at least partially the degree of endogeneity in monetary policy. However, in contrast to the results in figure 1, there do not appear to be systematic differences between countries in the covariance of home-country interest rates with the German interest rate, viz. trade to GDP ratio or exchange rate volatility.[29]

D. *Sensitivity Analysis: EMS-Period and Dynamics*

The European Monetary System came into effect in 1979 and committed countries to keep their exchange rates within bands of the German rate. This should have increased the role of leadership of the Bundesbank and affected the mechanism we exploit in our identification strategy.[30] We replicated the baseline regression for the EMS era. Overall,

several tests for nonstationarity in interest rates and cointegration which are summarized in di Giovanni et al. (2005). Overall, although we do not find that interest rates have unambiguous stochastic trends, for some specifications we cannot reject a unit root. However, for those countries we also find that the interest rate exhibits a cointegrating relationship with Germany. For example, this can be seen for Great Britain, the Netherlands, or Austria in the case of the standard Dickey-Fuller test for specifications with four lags of output growth and inflation as control variables.

[26] These results are robust to classification based on alternative measures of trade such as exports to Germany relative to GDP, or the ratio of trade to Germany and trade with the rest of the world.

[27] Note that the fact that OLS and IV are rather similar in magnitude for high-trade countries suggests there may not be a strong direct effect of German interest rates on the domestic economy.

[28] A high correlation in output growth need not imply greater economic integration. For example, countries that are subject to more large common shocks may tend to have more highly correlated output growths. This ranking also pools countries that clearly have independent monetary policy and low trade integration, such as Italy or Great Britain, in the high-correlation sample.

[29] Note that we would not have necessarily expected any systematic difference, since countries who had the option for more independence may still have an incentive to tie themselves to the German rate for other reasons (such as to foster convergence in the process of European integration).

[30] By further constraining countries' monetary policy choices, we expect the EMS to have led to more negative OLS estimates of the effect of interest rates on growth. However, von Hagen and Fratianni (1990) speculate that the Bundesbank itself may have become more lenient on inflation, since inflation's negative consequences for the German economy would be partially exported to the other countries under fixed exchange rates. This would imply lower IV estimates, since German monetary policy may have become more endogenous.

Table 3.—Pooled OLS, IV, and First-Stage Estimates by Country Groups (Unweighted)

	Low Country Sample				High Country Sample			
Controls	(1) OLS	(2) IV	(3) OLS − IV	(4) FS	(5) OLS	(6) IV	(7) OLS − IV	(8) FS
	Panel A: Trade with Germany/Output							
(A) No controls	−0.028	−0.112	0.084	0.882	−0.057	−0.098	0.042	0.801
	(0.014)	(0.032)	(0.029)	(0.068)	(0.027)	(0.037)	(0.025)	(0.041)
	[0.011]	[0.029]	[0.027]	[0.119]	[0.023]	[0.029]	[0.019]	[0.095]
(B) One lag of growth	−0.035	−0.151	0.116	0.815	−0.058	−0.107	0.049	0.777
and inflation	(0.015)	(0.035)	(0.032)	(0.066)	(0.028)	(0.041)	(0.029)	(0.043)
	[0.016]	[0.036]	[0.032]	[0.110]	[0.031]	[0.039]	[0.024]	[0.093]
(C) Four lags of growth	−0.035	−0.178	0.143	0.722	−0.039	−0.086	0.047	0.742
and inflation	(0.016)	(0.041)	(0.038)	(0.065)	(0.028)	(0.042)	(0.032)	(0.046)
	[0.018]	[0.044]	[0.040]	[0.114]	[0.035]	[0.043]	[0.025]	[0.091]
(D) One lag of growth	−0.041	−0.194	0.152	0.620	−0.061	−0.020	−0.041	0.696
and inflation,	(0.017)	(0.052)	(0.049)	(0.072)	(0.031)	(0.048)	(0.036)	(0.048)
different for each	[0.018]	[0.059]	[0.056]	[0.109]	[0.033]	[0.050]	[0.038]	[0.093]
country								
(E) Four lags of growth	−0.043	−0.191	0.148	0.627	−0.061	−0.020	−0.041	0.696
and inflation,	(0.017)	(0.051)	(0.048)	(0.071)	(0.031)	(0.048)	(0.036)	(0.048)
different for each	[0.019]	[0.058]	[0.055]	[0.107]	[0.033]	[0.050]	[0.038]	[0.093]
country								
Average coefficient	**−0.036**	**−0.165**	**0.129**	**0.733**	**−0.055**	**−0.066**	**0.011**	**0.743**
Median coefficient	**−0.035**	**−0.178**	**0.143**	**0.722**	**−0.058**	**−0.086**	**0.042**	**0.742**
Standard deviation	**0.006**	**0.034**	**0.029**	**0.115**	**0.009**	**0.043**	**0.048**	**0.047**
Controls	Panel B: Correlation of Business Cycles							
(A) No controls	−0.029	−0.097	0.068	0.757	−0.037	−0.117	0.080	0.972
	(0.020)	(0.047)	(0.042)	(0.067)	(0.011)	(0.019)	(0.015)	(0.062)
	[0.015]	[0.035]	[0.032]	[0.114]	[0.014]	[0.026]	[0.022]	[0.121]
(B) One lag of growth	−0.038	−0.126	0.088	0.703	−0.033	−0.125	0.092	0.883
and inflation	(0.021)	(0.050)	(0.045)	(0.066)	(0.012)	(0.022)	(0.018)	(0.060)
	[0.020]	[0.041]	[0.035]	[0.109]	[0.016]	[0.030]	[0.025]	[0.109]
(C) Four lags of growth	−0.042	−0.157	0.115	0.605	−0.027	−0.117	0.090	0.884
and inflation	(0.021)	(0.057)	(0.053)	(0.063)	(0.012)	(0.022)	(0.018)	(0.061)
	[0.024]	[0.051]	[0.045]	[0.111]	[0.014]	[0.027]	[0.023]	[0.118]
(D) One lag of growth	−0.044	−0.145	0.101	0.519	−0.044	−0.134	0.090	0.780
and inflation,	(0.023)	(0.078)	(0.074)	(0.079)	(0.014)	(0.027)	(0.024)	(0.064)
different for each	[0.024]	[0.083]	[0.080]	[0.111]	[0.020]	[0.032]	[0.026]	[0.116]
country								
(E) Four lags of growth	−0.047	−0.142	0.096	0.528	−0.044	−0.134	0.090	0.780
and inflation,	(0.023)	(0.076)	(0.073)	(0.076)	(0.014)	(0.027)	(0.024)	(0.064)
different for each	[0.026]	[0.082]	[0.078]	[0.106]	[0.020]	[0.032]	[0.026]	[0.116]
country								
Average coefficient	**−0.040**	**−0.134**	**0.094**	**0.623**	**−0.037**	**−0.125**	**0.089**	**0.860**
Median coefficient	**−0.042**	**−0.142**	**0.096**	**0.605**	**−0.037**	**−0.125**	**0.090**	**0.883**
Standard deviation	**0.007**	**0.023**	**0.017**	**0.106**	**0.007**	**0.008**	**0.005**	**0.081**

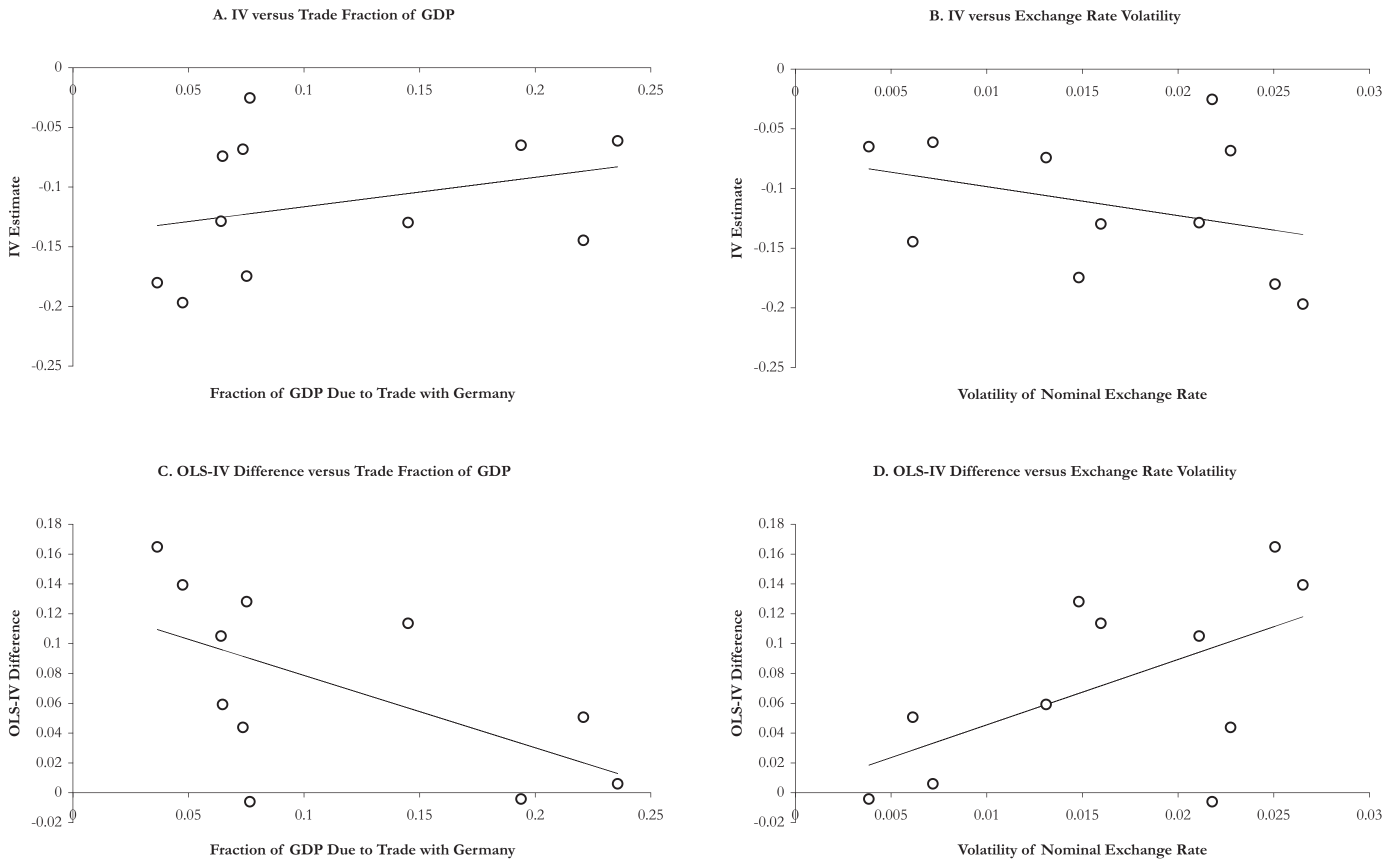
Notes: Table gives pooled OLS and IV, OLS-IV, and first-stage estimates of the effect of nominal interest rates on quarterly real economic growth for all countries except Germany. Sample is divided according to trade with Germany (panel A): low sample (ESP, FRA, GBR, ITA, NOR, PRT, SWE); high sample (AUT, BEL, CHE, NLD); and output growth correlation (panel B): variables specified under controls. Each estimate includes season indicators fully interacted with country indicators. In rows (D) and (E), lags are chosen separately for each country using significance levels. Estimates are equally weighed. The last three rows report the mean, median, and standard deviation of the coefficient estimates in rows (A) to (E). Standard errors in parentheses are Huber-Eicher-White standard errors and are robust to heteroskedasticity. Standard errors in square brackets are fourth-order Newey-West standard errors and are robust to fourth-order autocorrelation.

the results (available in di Giovanni et al., 2005) confirm those of table 1; with or without lags of growth and inflation IV estimates are systematically more negative than OLS estimates, and more so for larger countries. The differences between the EMS period and the full sample are small but as expected. Most countries experience a small increase in the magnitude of OLS coefficients. Similarly, most countries see a slight reduction in the size of IV estimates.

As noted in section II, estimates for the static model of equations (2) and (5) are a reduced-form parameter for the stance of monetary policy over the recent past. Specifically, if there are lagged effects of nominal interest rates on output growth, the results in tables 1, 2, and 3 can be interpreted as the weighted sum of the impact of current and lagged interest rates (see equations [13] and [14]). The differences in the point estimates across countries could thus be partly explained by the accumulation of differential effects over time and differences in the persistence of interest rates.

As a check on our results, we also ran a dynamic specification including lagged home-country interest rates,

FIGURE 1.—IV AND OLS-IV AND COUNTRY CHARACTERISTICS



assuming that lagged interest rates are predetermined. This assumption is tenuous, and would be violated if the central bank were able to accurately estimate output growth more than one period ahead. The results suggest that lagged interest rates may be endogenous as well, consistent with monetary policy actions with a horizon of several quarters.[31] Unfortunately, as suggested in section II, the lags of German interest rates are too persistent to provide separate instruments for lags of followers' interest rates, and thus we cannot move beyond this point.

As an alternative, to obtain insights on the dynamic effect of German interest rates we have also estimated a series of simple VARs. A VAR allows us to trace the full dynamic effect of changes in German interest rates on domestic output growth. This yields an estimate of the "reduced form" (RF) effect of the instrument on output growth going through all channels of the system. We compare the impulse response functions from this RF model to that of a typical domestic VAR. If there is a strong forward-looking bias, we would expect the immediate and persistent effect of the German interest rate to be stronger than that of the domestic interest rate.[32]

We specify our VARs as four-variable models including output growth, inflation, the growth rate of the German mark exchange rate, and interest rates. As commonly done in the literature, the innovations to the interest rate equations are identified by imposing a certain ordering of the variables in the model. We also include the growth rate of the commodity price index as an exogenous variable.[33] In the RF model the German interest rate is ordered first, and can thus have contemporaneous effects on domestic output and inflation. Following the literature, in the domestic VAR the interest rate follows growth and inflation such that there is no contemporaneous feedback from the interest rate to these variables. We use one lag for all variables of the system (which corresponds to three lags if we were to use the level of output), as well as the contemporaneous commodity price growth rate. Adding more lags significantly reduced precision, but overall results are similar. The standard errors are obtained from a bootstrap with replacement, and are generally quite high.

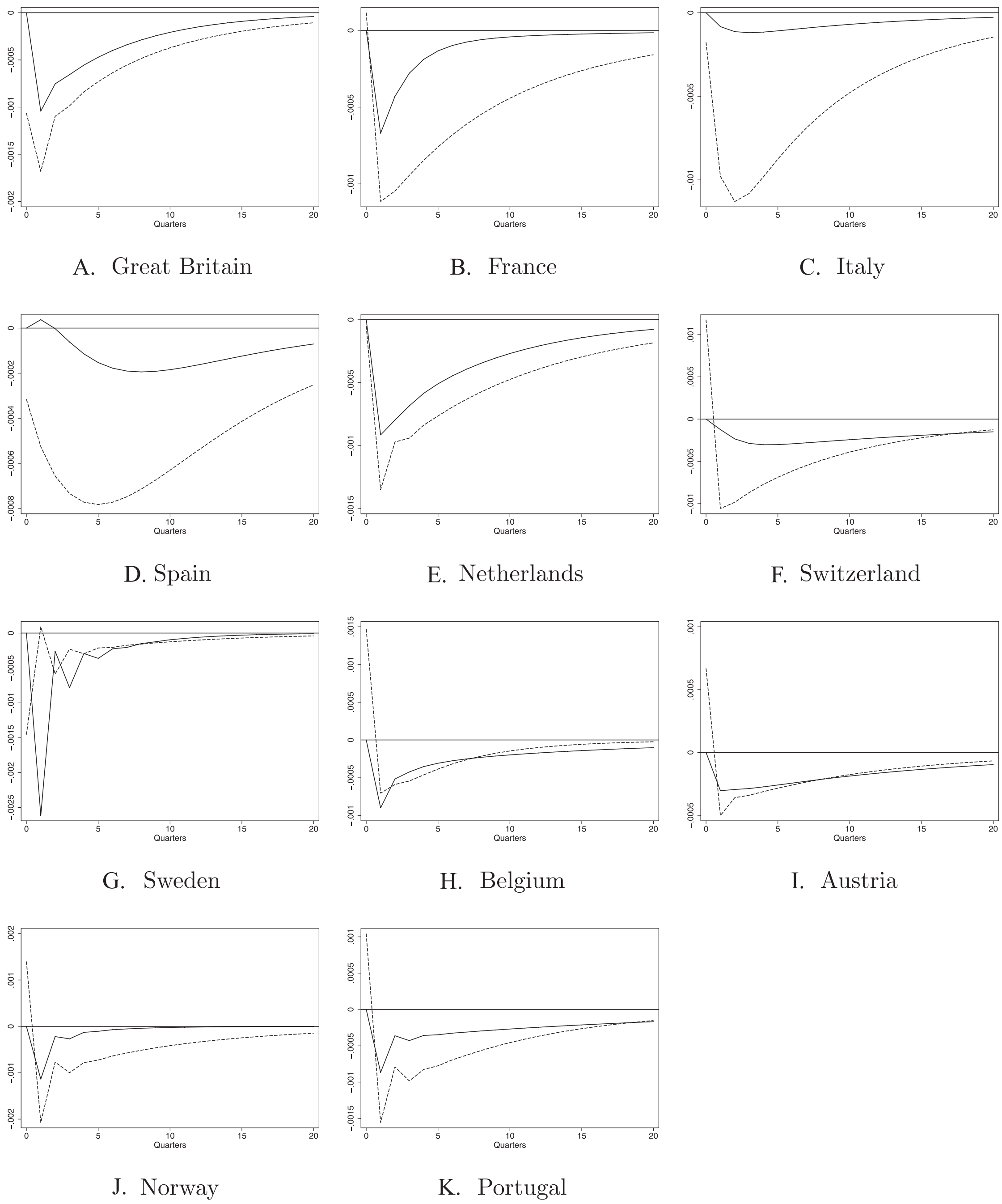
Figure 2 and Table 4 show the impulse response function (IRF) and the cumulated IRF for both models separately by

[31] See di Giovanni et al. (2005) for further discussion.

[32] The straightforward calculation of the omitted variable bias does not apply, since the impulse response function captures effects going through all other lags of the system. However, if lags of domestic interest rates also correlate with future shocks to inflation and output, then we should see a dampening of the entire propagation mechanism. The intuition is clear in a simple model of two lags with central-bank forecasts reaching two periods ahead.

[33] All variables except for the exchange rate are seasonally detrended.

FIGURE 2.—EFFECTS OF DOMESTIC AND GERMANY MONETARY POLICY SHOCKS ON DOMESTIC OUTPUT GROWTH: A COMPARISON



Notes: The figures present orthogonalized impulse response functions (IRFs) for domestic monetary policy (i_t) and Germany monetary policy (i_t^*) shocks. The IRF for the shock to domestic monetary policy is estimated from model 1A (see appendix A), and is represented by solid lines (—). The IRF for the shock to German monetary policy is estimated from model 1B, and is represented by dashed lines (--). Note that each country's figure is on a separate scale.

country, respectively. These IRFs refer to the effect of a standard deviation increase in the innovation to the interest rate equation obtained by a Choleski factorization. One can see clearly that the immediate and cumulated shocks of German interest rates are higher on average than that of domestic interest rates, and more so for larger countries, consistent with results from the previous sections.

Note that albeit the results appear qualitatively similar to those of our main analysis, the numerical magnitudes are hard to compare. First, the IRF captures the effect of an innovation running through the entire system, summing direct and indirect effects of exchange rates. The IV estimates summarize the dynamic effect of interest rates only. Second, the IRF is based on an innovation to the underlying interest rate equation, whereas the IV estimate relates to an increase of the interest rate itself. To nevertheless compare the order of magnitude taking the average coefficient from table 2 of 0.13, the predicted effect of a 1 standard deviation

TABLE 4.—THE EFFECT OF DOMESTIC AND GERMAN INTEREST RATES IN VECTOR AUTOREGRESSION FRAMEWORK

Panel A: Cumulate Impulse Response Function; Recursive VAR Domestic Interest Rate Shock (Model A)

Quarter	DEU	AUT	BEL	CHE	ESP	FRA	GBR	ITA	NLD	NOR	PRT	SWE
1	−0.0012	−0.0005	−0.0009	−0.0001	0.0000	−0.0007	−0.0010	−0.0001	−0.0009	−0.0011	−0.0009	−0.0026
	(0.0013)	(0.0007)	(0.0012)	(0.0007)	(0.0016)	(0.0010)	(0.0008)	(0.0012)	(0.0013)	(0.0011)	(0.0010)	(0.0013)
5	−0.0046	−0.0020	−0.0025	−0.0010	−0.0003	−0.0017	−0.0034	−0.0005	−0.0035	−0.0018	−0.0024	−0.0044
	(0.0043)	(0.0023)	(0.0034)	(0.0028)	(0.0090)	(0.0027)	(0.0023)	(0.0031)	(0.0035)	(0.0023)	(0.0020)	(0.0024)
10	−0.0074	−0.0034	−0.0036	−0.0021	−0.0012	−0.0020	−0.0049	−0.0009	−0.0052	−0.0020	−0.0039	−0.0052
	(0.0067)	(0.0038)	(0.0050)	(0.0048)	(0.0151)	(0.0032)	(0.0035)	(0.0033)	(0.0054)	(0.0027)	(0.0030)	(0.0029)
15	−0.0090	−0.0045	−0.0044	−0.0029	−0.0020	−0.0022	−0.0056	−0.0011	−0.0062	−0.0021	−0.0051	−0.0054
	(0.0083)	(0.0050)	(0.0061)	(0.0062)	(0.0184)	(0.0034)	(0.0042)	(0.0034)	(0.0070)	(0.0029)	(0.0038)	(0.0032)
20	−0.0100	−0.0052	−0.0049	−0.0035	−0.0025	−0.0023	−0.0059	−0.0013	−0.0066	−0.0021	−0.0060	−0.0055
	(0.0094)	(0.0060)	(0.0069)	(0.0072)	(0.0202)	(0.0035)	(0.0047)	(0.0035)	(0.0081)	(0.0029)	(0.0044)	(0.0033)

Panel B: Cumulate Impulse Response Function; Recursive VAR German Interest Rate Shock (Model B)

Quarter	AUT	BEL	CHE	ESP	FRA	GBR	ITA	NLD	NOR	PRT	SWE
1	−0.0001	0.0008	0.0003	−0.0008	−0.0010	−0.0027	−0.0014	−0.0014	−0.0006	−0.0003	−0.0014
	(0.0014)	(0.0019)	(0.0021)	(0.0033)	(0.0017)	(0.0013)	(0.0024)	(0.0020)	(0.0013)	(0.0015)	(0.0014)
5	−0.0016	−0.0011	−0.0029	−0.0038	−0.0045	−0.0063	−0.0057	−0.0049	−0.0038	−0.0036	−0.0027
	(0.0025)	(0.0032)	(0.0031)	(0.0096)	(0.0033)	(0.0023)	(0.0039)	(0.0033)	(0.0025)	(0.0026)	(0.0023)
10	−0.0028	−0.0021	−0.0053	−0.0073	−0.0071	−0.0087	−0.0088	−0.0078	−0.0063	−0.0063	−0.0034
	(0.0035)	(0.0043)	(0.0039)	(0.0158)	(0.0047)	(0.0032)	(0.0046)	(0.0045)	(0.0037)	(0.0035)	(0.0030)
15	−0.0036	−0.0025	−0.0067	−0.0098	−0.0087	−0.0100	−0.0104	−0.0096	−0.0077	−0.0079	−0.0038
	(0.0041)	(0.0048)	(0.0044)	(0.0192)	(0.0055)	(0.0039)	(0.0050)	(0.0053)	(0.0045)	(0.0042)	(0.0033)
20	−0.0040	−0.0026	−0.0074	−0.0113	−0.0096	−0.0107	−0.0112	−0.0107	−0.0085	−0.0087	−0.0040
	(0.0045)	(0.0051)	(0.0047)	(0.0208)	(0.0061)	(0.0044)	(0.0052)	(0.0058)	(0.0050)	(0.0048)	(0.0036)

Notes: This table gives accumulated impulse response functions for output growth given a one standard deviation shock to the interest rate (obtained from a Choleski factorization). Panel A presents results for model A, which includes domestic output growth, domestic inflation, the German mark exchange rate growth, and the domestic interest rate (in this order). Panel B presents results for model B, which includes the German interest rate, domestic output growth, domestic inflation, and the German mark exchange rate growth (in this order). Both models include one lag of the endogenous variables, as well as the contemporaneous and one lag of commodity price growth as an exogenous variable. All output and price data are seasonally adjusted. Bootstrapped standard errors are presented in parentheses.

increase in German output is about 0.003.[34] If we take the initial value of the cumulated IRF not affected by feedback through the system as the lower bound, and the cumulated effect of later years as the upper bound, we see that IV has a similar order of magnitude. As expected, the cumulated IRF shocks quickly become larger.

We also tried a variety of alternative specifications, which are summarized in a Web appendix available from the authors' Web sites. As is well known, results are not very robust to the ordering of variables, in particular for the position of the interest rate. On the other hand, the inclusion of endogenous variables such as German output growth and inflation did not change much. As is not uncommon for VAR application, standard errors are typically high, especially for impulse response functions. Overall, not surprisingly, some specification changes make substantial differences to our results. We stick to a standard VAR specification here, and refer to a recent paper analyzing European data (Mojon & Peersman, 2003) and to classic discussions in the literature with in-depth treatment of these points (for example, Christiano, Eichenbaum, & Evans, 1999; and Stock & Watson, 2001).[35]

IV. Conclusion

We have presented a sequence of simple estimates of the effect of monetary policy on real output growth, ranging from least squares contrasts to instrumental variables estimates. The identification strategy we have pursued attempts to exploit the fact that monetary policymakers may sometimes have competing goals. In particular, since the breakdown of the Bretton Woods system, many European central banks have followed the leadership of the Bundesbank in setting monetary policy to stabilize their exchange and inflation rates. Using quarterly German nominal interest rates as an instrument for other European countries' nominal interest rates, we estimate that the causal effect of a 5 percentage point increase in nominal interest rates is a contraction in annual real growth of 2 to 3 percentage points. This is in contrast to naïve OLS estimates, which suggest a more modest contraction of 0.5 to 1 percentage point.

The primary threat to the econometric validity of our IV estimates is the potential for economic shocks common to the European community. This is less of a concern for large countries with low trade with Germany, and indeed we find much larger differences between OLS and IV for these countries. However, for smaller countries in particular, our IV estimates will be too conservative, in the sense of being

[34] The standard deviation of German interest rates is 0.023; see appendix table B1.

[35] As suggested by a referee, we also included both domestic and German interest rates into a single VAR. The results from these models can be interpreted only under an extension of the assumption in equation (3); under the assumption in equation (4) maintained in this paper, the IRF of neither the German nor the domestic interest rate can be interpreted. We experimented with three models, each differing by whether further German variables are included as either exogenous or endogenous variables. Again, the results are very sensitive to the particular ordering chosen. If the German interest rate is ordered before the domestic variable, its effect is bigger than that of the domestic interest rate, and sometimes even reverses the sign of the latter's effect.

biased toward OLS. Since our IV estimates are in fact larger than our OLS estimates on average, we view our estimates as consistent with fairly decisive impacts of monetary policy on real output.

The difference between OLS and IV estimates may be interpreted as a measure of the endogeneity component of monetary authority actions. We report two key findings regarding the OLS-IV difference. First, we find that the difference is unaffected by the inclusion of lagged values of GDP growth or inflation as control variables. This finding implies that traditional controls for the history of the system are not rich enough to capture the information available to central bankers. This suggest the relevance of the recent extensions of the VAR approach to control variables of high dimension (see, for example, Bernanke, Boivin, & Eliasz, 2005). Second, we use the OLS-IV difference to directly test for the presence of bias in simple estimates, by relating the size of the bias to economic conditions affecting monetary policy. We show that the difference is decreasing with respect to the economic closeness between a country and Germany, as measured by physical distance and trade with Germany. We also show that the difference is increasing with exchange rate volatility vis-à-vis the German mark.

APPENDIX A

Description of VAR Models

This appendix describes the recursive VAR models that we estimate to compare with our core results. We follow Mojon and Peersman (2003) as closely as possible, but also make adjustments in order to remain consistent with the main OLS and IV regressions. The core recursive VAR structural representation can be written as

$$\mathbf{A_0}\begin{bmatrix}\mathbf{Y}_t^*\\ \mathbf{Y}_t\end{bmatrix} = \mathbf{A_1}(L)\begin{bmatrix}\mathbf{Y}_{t-1}^*\\ \mathbf{Y}_{t-1}\end{bmatrix} + \mathbf{B}(L)\mathbf{X}_t + \begin{bmatrix}\varepsilon_t^*\\ \varepsilon_t\end{bmatrix}, \tag{A1}$$

where $\mathbf{Y}_t$ is the block of domestic variables, $\mathbf{Y}_t^*$ is the block of German variables, $\mathbf{X}_t$ is a block of exogenous variables, ε_t are the structural shocks, and the matrix $\mathbf{A_0}$ is an upper-triangular matrix that defines the recursive structure of the system.

The domestic endogenous variables include seasonally adjusted output growth and inflation, the bilateral German-domestic exchange rate growth,[36] and the nominal interest rate. The German block is similar, but does not include any exchange rate. Finally, we include seasonally adjusted commodity price growth in $\mathbf{X}_t$. All specifications have one lag of endogenous variables, and the contemporaneous and one lag of the growth rate of commodity prices as exogenous variables.[37]

We examine two main VAR specifications. The first ignores the German block of endogenous variables, $\mathbf{Y}_t^*$, which is akin to running a standard domestic VAR similar to those estimated in the U.S. monetary policy literature (though we also include the exchange rate). The second specification examines the impact of German interest rate shocks in a system where we do not include the domestic interest rate. This allows us to compare the direct impact of a German monetary policy shock with that of a domestic shock of the previous model.

1. *Model 1A: Domestic VAR*

The endogenous variables of the first VAR setup using domestic variables can be expressed as follows:

$$\mathbf{Y}_t' = [y_t, \pi_t, x_t, i_t], \tag{A2}$$

where y_t is domestic output growth, π_t is inflation, x_t is exchange rate growth, and i_t is the domestic interest rate. This ordering implies that the domestic interest rate responds contemporaneously to the other endogenous variables, the exchange rate to all but the interest rate, and so forth. Ignoring the exchange rate, this is a standard setup in the U.S. domestic monetary policy literature (for example, see Stock & Watson, 2001). The recursive structure then implies that orthogonalized monetary policy shocks (feeding through via i_t) will not have a contemporaneous effect on output growth.

2. *Model 1B: German Interest Rate in Domestic VAR*

The endogenous variables of the second VAR setup using domestic variables and the German interest rate can be expressed as follows:

$$[\mathbf{Y}_t^*\mathbf{Y}_t]' = [i_t^*, y_t, \pi_t, x_t], \tag{A3}$$

where y_t is domestic output growth, π_t is inflation, x_t is exchange rate growth, and i_t^* is the German interest rate. In this setup, shocks to domestic variables do not feed through to the German interest rate contemporaneously, whereas the shocks to German monetary policy do have a contemporaneous impact on the domestic block of variables.

[36] We also experimented with the real effective exchange rate, but results were similar, and we lost observations given data coverage.

[37] Standard tests suggested the use of one lag, though results do not vary greatly if we include two or three lags. One should note that Mojon and Peersman (2003) use two or three lags of *levels,* so the use of one lag of growth fits with this.

APPENDIX B

TABLE B1.—COUNTRY SUMMARY STATISTICS

Country	GDP	Trade/GDP	sd(NER)	GDP Growth	Interest Rate	Inflation
	251.46	0.194	0.004	0.005	0.065	0.038
Austria	—	—	—	(0.081)	(0.023)	(0.018)
	302.22	0.236	0.007	0.005	0.070	0.029
Belgium	—	—	—	(0.081)	(0.023)	(0.018)
	1,747.97	0.065	0.013	0.005	0.096	0.059
France	—	—	—	(0.007)	(0.043)	(0.040)
	2,400.66	—	—	0.005	0.060	0.032
Germany	—	—	—	(0.010)	(0.023)	(0.015)
	1,794.86	0.047	0.027	0.005	0.090	0.076
Great Britain	—	—	—	(0.010)	(0.037)	(0.055)
	1,465.90	0.064	0.021	0.005	0.124	0.096
Italy	—	—	—	(0.008)	(0.042)	(0.054)
	511.56	0.221	0.006	0.006	0.066	0.021
Netherlands	—	—	—	(0.009)	(0.025)	(0.016)
	221.58	0.075	0.015	0.010	0.101	0.050
Norway	—	—	—	(0.041)	(0.040)	(0.038)
	149.45	0.074	0.023	0.007	0.147	0.130
Portugal	—	—	—	(0.019)	(0.053)	(0.078)
	836.10	0.037	0.025	0.006	0.126	0.095
Spain	—	—	—	(0.005)	(0.048)	(0.054)
	300.80	0.077	0.022	0.007	0.103	0.046
Sweden	—	—	—	(0.116)	(0.038)	(0.038)
	309.47	0.145	0.016	0.003	0.027	0.030
Switzerland	—	—	—	(0.012)	(0.025)	(0.022)

Notes: All other variables are averaged over 1974–1998. Standard deviations are in parentheses. Trade is the total value of bilateral trade between the country and Germany. The exchange rate volatility measure, sd(NER), is calculated by taking the standard deviation of the change of end-of-month log nominal exchange rate viz. Germany over 1974Q1–1998Q4. Output growth is the quarterly rate. It is calculated by first deseasonalizing output growth for each country. The interest rate is a quarterly average of the domestic call/money market rate. The inflation rate is calculate from the annual average of the quarterly GDP price deflator.

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