On the European Monetary System: the Spillover Effects of German Shocks and Disinflation

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Abstract

We analyze the disinflationary experience between 1979-1993 for two traditionally inflationary countries of the European Monetary System: France and Italy. For each country, a vector autoregressive model is estimated. Shocks in the model combine domestic and foreign sources. The latter capture the world oil price shocks as well as nominal and real shocks originating in Germany. Under investigation is the hypothesis that shocks originating in Germany have a spillover disinflationary effect in France and Italy. The empirical evidence provides support to the validity of this hypothesis. Furthermore, German shocks account for an important share of the total price variance in France and Italy. These results indicate that the interaction between countries of the European Monetary System has contributed to the success of the disinflationary experiences of the eighties. The evidence sheds, therefore, some light on potential benefits that may be further realized as countries of the European Monetary System move towards their objective of achieving a single currency under a unified monetary system.

I Introduction

The monetary system of the western European countries can be broadly characterized as a system of two exchange rate regimes: the managed float relative to the U.S. dollar and the fixed, but adjustable, peg in the European Monetary System (EMS). Most of the Western European countries opted for a fixed, but adjustable, exchange rate regimes in the late 1970s. This system was motivated by the desire of these countries to guarantee the stability of the intra European trade. Subsequently, the EMS has evolved in 1979 as a set of arrangements that include the exchange rate mechanism. According to this arrangement, currencies of member countries could not fluctuate around fixed parities by more than $\pm$ 2.25%.[1] Adjustments of these pegged rates were possible with the mutual agreement of all countries involved.[2] Given this arrangement, there are two interesting aspects that pertain to the functioning of the EMS: policy coordination among the EMS member countries and its impact on the evolution of inflation in some member countries.

Inflation rates of some member countries of the European Monetary System (EMS) declined from peaks well above 10% in the early 1980s to 2-3% in the 1990s. This decline has occurred when inflation rates all over the developed world have decreased. Nonetheless, since low inflation rates have occurred within the EMS, it has led to opinions that there are features in the EMS arrangements that have facilitated, or at least contributed to the success of the disinflation process.

The European Monetary System consists of eight countries: Belgium, Denmark, France, Germany, Ireland, Italy, Luxemborg, and the Netherlands. This set of countries contains some traditionally inflationary countries, as Italy, France, and Ireland, countries with distinguished anti-inflationary bias, as Germany, and countries in between as the four countries of the core. When the EMS was founded in March 1979, there was a wide range of inflation rates among the members.

[1] For some currencies this band was wider with a maximum of $\pm$ 6 percent. If the bilateral rates diverge from the agreed parity, both central banks would begin symmetric compulsory intervention with the aim to bring the exchange rate close to its fixed parity. In July 1993, these bands were widened to $\pm$ 15 percent. Our sample period does not capture the span after the summer of 1993.

[2] From its beginning there were around a dozen of major alignments, most of them at the beginning of the functioning of the EMS.

Germany was at the low end, with an average inflation rate of 4 percent, while France and Italy were at the other extreme, with an average inflation rate of 13 percent for Italy and slightly under 10 percent for France. The rates for other members, such as Belgium, Denmark, and the Netherlands were between 6.5 percent and 8.9 percent; Collins (1990). In this paper, we concentrate on the relationship between the low inflation country, Germany, and the inflation-prone countries, France and Italy.[3]

Traditionally, when there was a possibility of realignments, or under flexible exchange rate regimes, countries like France and Italy, had relatively high inflation, despite the efforts of their governments to decrease the inflationary trend.[4] Thus, it could be argued that relatively high inflation in France and Italy before the 1980s resulted from credibility problems. In other words, both governments lacked strong commitment to monetary discipline. Indeed, Fischer (1987) described the European Monetary System as "an arrangement for France and Italy to purchase a commitment to low inflation by accepting German monetary policy."[5]

How does the working of the European Monetary System provide commitment for disinflation? Two explanations are commonly offered in the literature with respect to this question. Frattiani and von Hagen (1992) summarize them into cooperative and disciplinary interpretations of the EMS. The cooperative interpretation assumes that monetary policies were formed as a joint effort of member governments, while the disciplinary interpretation stresses the "spillover effect" stemming from the presence of the Bundesbank in the system. The first approach is expected to maximize the role of domestic policies and, subsequently, domestic shocks in determining the success of the domestic disinflation policy in the high inflation countries of the EMS. In contrast, the second

[3]Despite Ireland's historical inflationary experience, Ireland is not connected to Germany as France and Italy, which are geographically much closer. In addition, France, Germany and Italy are the cornerstone of the European integration. They have been members of the European Community since 1956. In contrast, Ireland joined the European Community much later (in the 1970s).

[4]Among the incentives for the government to inflate are usually domestic political distortions, strong trade unions, and expansionary government policies. "This happens because the public understands such incentives, anticipates the actions of monetary authorities and effectively neutralizes them, with the result that higher inflation is generated....."(Giovannini, 1992, p. 803).

[5]The quotation is based on Giavazzi and Giovannini (1989, p. 85).

approach is expected to maximize the role of the German monetary policy, and subsequently German monetary shocks in determining the success of the domestic disinflation policy in the high inflation countries of the EMS.

There are strong points that argue for the validity of both approaches. The establishment of the EMS was the result of efforts of a number of countries that aimed at establishing joint monetary policy as the outcome of a cooperative decision-making process of all member countries. Indeed, some authors (see, e.g., Melitz (1988) and Fratianni and von Hagen (1992)) are of the opinion that the EMS was characterized by cooperation in terms of monetary policies. However, advocates of the disciplinary view recall that previous fixed exchange rate arrangements were characterized by a hegemonic solution in which one country, usually the largest one, independently determined monetary policy for the system as a whole. Similar to the position of the Great Britain in the classical gold-standard period, or the United States in the Bretton Woods period, Germany in the European Monetary System is seen as the hegemonic country, the inflationary anchor of the system. In words of Giavazzi and Giovannini (1989) “Germany is the center country and runs monetary policy for the whole system."

From a practical standpoint, it is possible for the German dominance to arise within the European Monetary System. If there is a fixed exchange rate arrangement among n countries, there are only n-1 exchange rates to be fixed. Thus, assuming that n-1 countries exhibit monetary policy aimed at maintaining the pegged exchange rate, then the nth country could pursue an independent monetary policy. Why Germany is the proper candidate to play hegemonic role in the formulation of European monetary policies? Among the arguments for German dominance, usually the size of the German economy and the high credibility of the Bundesbank are mentioned. The credibility argument rests usually on Barro and Gordon (1983) who showed that the credibility of the monetary authority is crucial in determining the inflationary trend in a country. It is, therefore, plausible to perceive a scenario in which countries with traditionally higher inflation, by joining the EMS, actually delegated part of their monetary autonomy to the Bundesbank and achieved lower inflation in return. As suggested by Thygesen (1988, p. 5-6): “By pegging to the less inflationary currencies over long intervals, with the prospect that they cannot fully devalue in accordance with their excess inflation on the infrequent occasions of a realignment, the authorities of the weaker currencies gain credibility for their disinflationary stance."

Giavazzi and Giovannini (1989) argue that Germany was the only member country in the EMS whose domestic monetary conditions were relatively unanffected by the events originating from other member countries. They showed that the difference between onshore and offshore Deutschemark interest rates were considerably less affected in times of realignments than the French or the Italian interest rates. They concluded that German monetary policy was more independent of international conditions than was the French or the Italian policy. Cohen and Wyplosz (1989) use a vector autoregressive model of interest rates and monetary base to test the German dominance hypothesis. Their conclusion was that German monetary variables affect nominal variables in the other EMS countries. Fratianni and von Hagen (1992) use a vector autoregressive model to estimate cross-country monetary base growth rates. They refuted the hypothesis of German dominance, “these results do not imply dominance by the DM area, but they tell us that there is indeed a large degree of conformity of EMS policies with regard to outsiders."

As to the success of the disinflation policy within countries of the EMS, Collins (1988) agreed that there is no overwhelming evidence that the EMS members performed substantially better than non-EMS members in decreasing inflation rates in the period under consideration. Frankel and Phillips (1991) did not find support for the idea that imported credibility has helped to decrease inflation rates in France and Italy. They emphasized the change in domestic government policies with respect to their support of the EMS as the factor behind disinflation. This argument rests on the fact that the EMS was initially quite soft as a disinflationary mechanism. Major alignments occurred between 1979 and 1983. During that time, the French and Italian governments did not keep the nominal exchange rates pegged for too long as they were prepared for adjustments if the pressure on the exchange rates was considered excessive. In March 1983, the French government decided to hold firmly on the pegged rates. The Italian government decided in a similar spirit somewhat earlier.

In an effort to reconcile the previous differences, the present investigation will aim at developing direct evidence on the effect of German monetary dominance in determining disinflation in other EMS member countries. Specifically, we focus on the relative importance of domestic cooperative policies versus German spillover policies in determining disinflation in the two traditionally inflationary countries of the EMS: France and Italy. Towards this objective, we devise a model that includes five shocks:[6] shocks originating from the world oil market, real and monetary shocks originating from Germany, and real and nominal shocks originating from the home country. The investigation will then focus on analyzing the effects of these shocks on price inflation in France and Italy.

II Empirical Framework

The disciplinary interpretation of the EMS stresses that the decreased inflation rate in France and Italy hinges on the effects of the shocks originating in Germany. This explanation implies a large weight of German disturbances in explaining domestic prices in France and Italy. In contrast, the cooperative interpretation of the EMS suggests that the decreased inflation rate in the 1980s is primarily attributable to the change in domestic economic policies. That is, German disturbances are likely to have a low weight in explaining domestic prices in France and Italy.

To shed some light on these alternative hypotheses, we examine a multivariate system that includes the real world oil price (WOP), the real gross domestic product in 1990 constant prices of

[6]Traditionally, models of exchange rates, as for example, Fleming and Mundel model, would suggest that the importance of foreign shocks on the inflation rate depends on the exchange rate regime. Fixed exchange rate regimes, as in the EMS, are supposed to lead to the transmission of inflation from one country to another, while flexible exchange rates would insulate the economy from the outside disturbances. The theoretical work of Turnovsky (1981) and Marston (1985) in the traditional Mundell-Fleming setting with rational expectations, and of Stockman and Svensson (1987) and Svensson and van Wijinbergen (1989) in intertemporal optimization setting showed that insulation of the home economy from foreign disturbances, even under flexible exchange rates, applies only for some special cases. For a more elaborate discussion of this topic, see, for example, Genberg, Salemi and Swoboda (1987). The views also widely differ in assessing the role of the exchange rate regime in a disinflation process.

Germany (GRGDP), seasonally adjusted narrow money of Germany (GNM), real gross domestic products in 1990 constant prices for France (FRGDP) or Italy (IRGDP), and the consumer price index for France (FCPI) or Italy (ICPI). Our objective is to examine the sources of disturbances that affected the inflation rate in Italy and France. The data are quarterly from 1979.I to 1993.II. All data are taken from *International Financial Statistics* available from the International Monetary Fund.

The world real oil price represents changes in prices of basic commodities which could have an impact on the domestic price level as supply components or through their impact on the prices of domestic substitutes. The EMS began to function at the beginning of 1979 during the second world oil price shock. However, during most of the 1980s, the real oil price was more or less stable and somewhat declining in real terms. Real GDP in Germany approximates real (supply-side) shocks. Real output growth keeps prices from rising in Germany which is likely to have a spillover disinflationary effect on prices in other member countries of the EMS according to the disciplinary explanation of disinflation. Shocks to the narrow money of Germany are included to test for the spillover disinflationary effect of German monetary policy. In addition to foreign shocks, inflation is likely to vary in response to real and nominal domestic shocks. Shocks to real GDP approximate real domestic supply shocks. Output expansion is likely to moderate domestic price inflation. In addition shocks to the consumer price index approximate domestic nominal shocks. These shocks capture the effects of domestic spending as well as demand-side policies that aim at controlling price inflation.

In order to approximate relations among domestic and foreign shocks in the model, we rely on recently developed techniques for imposing long-run constraints. First, we estimate a vector autoregressive model and impose the constraints implied by structural relations among the model's variables. Secondly, we estimate a vector autoregressive model and impose the constraints implied by the long-run cointegration results among the model's variables. Having imposed the relevant constraints, we approximate the effects of the model's shocks on price inflation using the impulse response function. In addition, we rely on the variance decomposition results to measure the share of the model's shocks to total price variability.

III Methodology and Empirical Results

The variables of interest are denoted in the empirical model as follows: the world real oil price (O^f), real output in Germany (Y^f), the money supply in Germany (M^f), domestic real output in Italy or France (Y^d), and the domestic price level in Italy or France (P^d) where the superscript f denotes foreign variables and the superscript d denotes domestic variables. Let $X = (O^f, Y^f, M^f, Y^d, P^d)'$ and $u = (u^{Of}, u^{Yf}, u^{Mf}, u^{Yd}, u^{Pd})'$ is the vector of five disturbances. We assume that these five disturbances are uncorrelated and, therefore, their covariance matrix is diagonal, i.e., $E(u_t u_t') = \Sigma$, which is assumed to be diagnonal and $E u_t u_s' = 0$ for $s \neq t$. Alternatively, shocks are normalized such that $E(u_t u_t') = I$.

The variables in X contain unit roots. However, ΔX_t is stationary (see Table 1 for details) and we assume that it follows a stationary process of the form:

$$\Delta X_t = A(L) u_t = \sum_{j=0}^{\infty} A_j u_{t-j}, \tag{1}$$

where A_j are matrices in the lag operator L. In detail, (1) is written as:

$$\begin{aligned}
\Delta O_t^f &= a_{11}(L) u_t^{Of} + a_{12}(L) u_t^{Yf} + a_{13}(L) u_t^{Mf} + a_{14}(L) u_t^{Yd} + a_{15}(L) u_t^{Pd} && (2)\\
\Delta Y_t^f &= a_{21}(L) u_t^{Of} + a_{22}(L) u_t^{Yf} + a_{23}(L) u_t^{Mf} + a_{24}(L) u_t^{Yd} + a_{25}(L) u_t^{Pd} && (3)\\
\Delta M_t^f &= a_{31}(L) u_t^{Of} + a_{32}(L) u_t^{Yf} + a_{33}(L) u_t^{Mf} + a_{34}(L) u_t^{Yd} + a_{35}(L) u_t^{Pd} && (4)\\
\Delta Y_t^d &= a_{41}(L) u_t^{Of} + a_{42}(L) u_t^{Yf} + a_{43}(L) u_t^{Mf} + a_{44}(L) u_t^{Yd} + a_{45}(L) u_t^{Pd} && (5)\\
\Delta P_t^d &= a_{51}(L) u_t^{Of} + a_{52}(L) u_t^{Yf} + a_{53}(L) u_t^{Mf} + a_{54}(L) u_t^{Yd} + a_{55}(L) u_t^{Pd} && (6)
\end{aligned}$$

where $a_{ij}(L) = \sum_{k=0}^{\infty} a_{ij,k} L^k$. Furthermore, we also adopt the notation that $a_{ij}(1)$ is the sum of all coefficients and gives the effect of u_{jt} on variable i over time.

In order to decompose disturbances in the variables into the five sources specified in the model, restrictions on the multivariate dynamic system in (1) are necessary. To assure the exogeneity of the shocks in the VAR model, the relationship between variables will be analyzed under two sets of constraints: (i) theoretical constraints, as implied by the assumptions of structural relationships among the model's variables, and (ii) atheoretical constraints, as implied by the results of cointegration tests among the model's variables.

III-A Imposing Structural Constraints

In model (1), we assume the following constraints:

- u_t^{Of} is exogenous to the remaining four variables in a sense that in the long-run the remaining four macro variables have no impact on the world real oil price. This assumption yields four restrictions.

- Real output and the monetary aggregate in Germany (the hypothesized more dominant country in the EMS) are not affected in the long-run by movements of variables from less dominant countries. These assumptions yield four restrictions.

- The aggregate supply curve is vertical in the long-run for Germany. This implies that in the long-run, the effect of monetary shocks on real output vanishes. This assumption yields one restriction.

- The aggregate supply curve is vertical in the long-run for the home country. This implies that in the long-run, the effect of domestic nominal shocks on real output vanishes. This assumption yields one restriction.

In order to identify the model, we can estimate a finite order VAR, i.e.,:

$$\Delta X_t = \sum_{i=1}^{n} B_i \Delta X_{t-i} + e_t \tag{7}$$

The estimation of the VAR model in (7) requires, however, that we determine the appropriate lag length of the variables in the model where the maximum lag length n is chosen such that the residuals e_t are white noise. We use the likelihood ratio test, as outlined in Hamilton (1994), p. 296-8.[7] Table 2 presents the results of the likelihood ratio tests for lag determination. The null hypothesis that a set of variables is generated from a VAR system with n lags is tested against the alternative specification of n_1 lags where $n < n_1$. Based on the Chi-square significance level, there is a clear support for the null hypothesis of four lags for France. For Italy, the evidence appears less conclusive. The probability of accepting the null hypothesis of four lags is the largest. Nonetheless, the difference in probability of accepting the null hypothesis of four lags is not largely pronounced compared to the probability of accepting the null hypothesis of two lags. We determined the optimal lag length to be four lags, however. Choosing a lag span shorter than one year would probably not provide for significant dynamics of the system and more lags will take away too many degrees of freedom.[8] Since the elements of ΔX_t are stationary, the system in (7) can be inverted to obtain an infinite order MA representation, i.e.,:

$$\Delta X_t = e_t + C_1 e_{t-1} + C_2 e_{t-2} + ... \tag{8}$$

Comparing (1) and (8), we observe that u, the vector of original (structural) disturbances and e, the vector of innovations are related, i.e.,:

$$e_t = A_0 u_t, \tag{9}$$

and $A_j = C_j A_0$ for all j. Thus, knowing A_0, we can recover u from e, and also obtain A_j from C_j. Furthermore, from (9) we also note that:

$$E(e_t e_t') = \Omega = A_0 E(u_t u_t') A_0' = A_0 A_0' \tag{10}$$

[7] In this test, we use a modification of the likelihood ratio test to take into account the small sample bias, as suggested by Sims (1980, p. 17).

[8] We do not allow for different lag length since it is common to use the same lag lengths for all equations in order to preserve the symmetry of the system. For previous applications of this approach, see, .e.g., Bayoumi and Eichengreen (1992) and Blanchard and Quah (1989).

Since Ω is a symmetric matrix with known elements, it imposes 15 restrictions on the matrix of contemporaneous effects, A_0, which has 25 elements. Additional ten restrictions are needed to identify A_0, so that the orthogonal shocks u_{it} can be recoverd using equation (9). The traditional method is to pick A_0 as the Choleski factorization of Ω which has been criticized on the grounds that it imposes an arbitrary structure on the orthogonal u_{it} sequences. Blanchard and Quah (1989) propose an interesting way of circumventing the problem of arbitrary identification. This can be seen from the relationship between the matrices of long-term effects. If we evaluate the polynomials embedded in equations (1) and (8) at $L = 1$, and note the relationship that $A_1 = C_1 A_0$, where C_1 contains known elements. Once A_0 is identified, one can recover the orthogonal shocks using equation (9). In order to identify the shocks, we impose the following ten restrictions on the long-run matrix $A(1)$, which were discussed earlier. Note that these restrictions transform $A(L)$ at $L = 1$ into:

$$\begin{bmatrix} a_{11}(1) & 0 & 0 & 0 & 0 \\ a_{21}(1) & a_{22}(1) & 0 & 0 & 0 \\ a_{31}(1) & a_{32}(1) & a_{33}(1) & 0 & 0 \\ a_{41}(1) & a_{42}(1) & a_{43}(1) & a_{44}(1) & 0 \\ a_{51}(1) & a_{52}(1) & a_{53}(1) & a_{54}(1) & a_{55}(1) \end{bmatrix} \tag{11}$$

Furthermore, the above restrictions on the long-run sum of impulse responses imply nothing about the effect of shocks on prices. Instead, the model allows the estimation of the weight of domestic and foreign factors in determining inflation in France and Italy.

Once the VAR is estimated and the reduced-form shocks are obtained, we proceed with the second step, we recover the structural shocks. This identification procedure guarantees that the shocks are not correlated among each other by imposing structural constraints on the long-run relation between the reduced-form shocks. Figures 1 and 2 give the impulse response functions of price inflation in Italy and France to innovations in all five variables. The impulse response functions appear similar for all five variables in both countries. Two shocks are driving the domestic price level upward: domestic nominal shocks and real world oil price shocks. In addition, output growth is disinflationary in both countries, as evident by the negative response of price to the shocks to the growth of domestic real output. More importantly, the results using structural VAR provide support to the spillover disinflationary effect of German shocks. The response of price inflation to monetary and real shocks originating from Germany is negative in France and Italy. That is, output growth in Germany is an important factor in controlling price inflation that has a strong spillover effect on disinflation in France and Italy. In addition, the design of monetary policy in Germany that aims at controlling price inflation has had a strong spillover effect on disinflation in France and Italy.

In the next step, we calculate the percentage of the expected n-period ahead squared forecast error (variance decomposition) of price in Italy and France, as produced by the innovation in the structural model of (9). The results are represented in Table 3. The weights of real and nominal shocks originating from Germany account for a large share of price variability. Specifically, the proportion of price variance in France associated with German shocks ranged from 62.78 percent after two quarters to 67.99 percent after 24 quarters. In Italy, the proportion of price variance associated with German shocks ranged from 28.53 percent after two quarters to 41.72 percent after 24 quarters. In addition, real oil price shocks are an important determinant of price variability which is particularly evident for Italy. The share of world real oil price of price variability ranges from 10.56 percent after two quarters to 6.01 percent after 24 quarters in France. In contrast, this share ranges from 26.44 percent after two quarters in Italy to 25.15 percent after 24 quarters in Italy. The proportion of price variance in France associated with domestic real and nominal shocks ranges from 26.65 percent after two quarters to 26.00 percent after 24 quarters. In Italy, the proportion of price variance associated with domestic real and nominal shocks ranges from 45.02 percent after two quarters to 33.13 percent after 24 quarters.

Overall, the evidence imposing structural constraints provides strong support to the disinflationary share of German shocks in determining price variability in France and Italy. Specifically, imposing structural constraints isolates exogenous shocks originating in Germany, which appear important in explaining a large share of price variability in France and Italy over time. In addition, the structural constraints account for the dependency of domestic shocks on German shocks in the long-run. By accounting for this dependency, the exogenous component of domestic shocks appears less important in explaining price variability in France and Italy over time.

In summary, the impulse response functions from the structural VAR model indicate that innovations from domestic nominal variables and from world real oil price exhibit an upward effect on price inflation in France and Italy. In contrast, the effect of nominal and real shocks originating from Germany is negative on price inflation in France and Italy. In addition, the results of variance decomposition indicate that the exogenous shocks originating from Germany, not only drive the price level downward, but also explain quite a large part of the total variation of price.

III-B Imposing Cointegration Constraints

Since variables in log levels are non-stationary, we need to consider the long-run constraint implied by cointegration, i.e., whether the nonstationarity in level is due to a smaller number of common stochastic trends.

The test for cointegration follows the suggestions of Johansen (1988). Consider the unrestricted VAR from equation (7):

$$X_t = \sum_{i=1}^{n} B_i X_{t-i} + e_t$$

This model can be represented in the form:

$$\Delta X_t = \sum_{i=1}^{n-1} \Pi_i \Delta X_{t-i} + \Pi X_{t-n} + e_t \tag{12}$$

where

$$\begin{aligned} \Pi_i &= -I + B_1 + ... + B_i, \\ &= -(I - B_1 .... - B_i), \end{aligned}$$

where (I is a unit matrix) and n= maximum lag length. The rank of matrix Π can at most be five, the number of variables in the X_t vector. If the rank of matrix $\Pi = 5$, the vector X is integrated of order zero, i.e., stationary. If the rank of matrix $\Pi < 5$, the rank determines the number of cointegrating vectors in the VAR model explaining X_t. Hence, in our VAR model explaining five variables, there are at most four cointegrating vectors.

Table 5 presents the likelihood ratio results for the null hypothesis that the number of cointegrating vectors=0,1,2,3,4. The cointegration analysis reported in Table 4 shows that the joint process of five variables in both models is cointegrated at the five percent level of significance. These results indicate that this five variable system can be characterized by a lesser number of common stochastic trends which drive the system.[9] For the model of France, we detected two cointegrating vectors. For the model of Italy, we detected three cointegrating vectors. Accordingly, the residuals from the cointegrating regressions with the dependent and all independent variables expressed in levels are included in the estimation of the VAR system in (12). This vector error correction model imposes the constraint implied by the cointegration relationship among the variables in the system. Thus, we rely on the restricted vector autoregressive system with the appropriate number of error correction terms.

Having estimated the vector autoregression model, we then focus on the effects of innovations in the system, as represented by the impulse-response function and the variance decomposition. Figures 3 and 4 report the response of price to one standard deviation innovation in all five variables of the vector autoregressive model for Italy and France.[10] The impulse response functions show that price inflation, in France and Italy, accelerates in response to a rise in the world real oil price. In addition, nominal domestic shocks have long-lasting effects on price inflation in both countries. The inflationary effect of the domestic nominal shocks appears larger, however, than that of the real world oil price shocks in France. More importantly, it is interesting to note the disinflationary

[9] All pairwise combinations of real world oil price with the rest of the variables, as well as all pairwise combinations of the narrow money of Germany with the rest of the variables are not cointegrated at the five percent significance level. Thus, the cointegrating relationship rests on the strength of the relationship of real output of Germany with real output of France and Italy. Evidently, a common component of the long-run movement of real output is due to technological shocks which spread across borders.

[10] The order of variables is WOP, GRGDP, GNM, FRGDP, FCPI in the model for France and WOP, GRGDP, GNM, IRGDP, ICPI in the model for Italy.

effects of German shocks in France and Italy. Shocks to real output growth in Germany decrease price inflation in France and Italy. That is, output growth in Germany is an important factor in controlling price inflation that has a strong spillover effect on disinflation in France and Italy. In addition, German monetary shocks have negative effects on price inflation which appear particularly evident in France. That is, the design of monetary policy in Germany, which aims at controlling price inflation, has had a strong spillover effect on disinflation in France, and to a lesser extent in Italy.

We complement the observations of the change in price inflation with the effects of the shocks on real output growth. In Figures 3 and 4, we see that real oil price shocks cause a permenant decline in real output growth. This is consistent with the inflationary effect of oil price shocks on price inflation in France and Italy. The growth of output in Germany has a positive expansionary effect on output growth in France. This is consistent with the disinflationary effect of real shocks in Germany on price inflation in France and Italy. This is in contrast to the design of monetary policy in Germany which aims at controlling the inflationary effects of the increased demand. Consequently, monetary growth in Germany has a negative contractionary effect on output growth in France and Italy. Consistently, monetary shocks in Germany have a spillover disinflationary effect on prices, particularly in France. Domestic shocks to output growth have larger expansionary effect on output growth over time in Italy compared to France. Consistently, domestic real shocks have disinflationary effects on price in Italy. In contrast, nominal domestic shocks interfere with output growth in Italy and accelerate price inflation. Consistently, the inflationary effect of nominal domestic shocks appears more moderate in France compared to Italy.

The results of the forecast error variance decomposition are reported in Table 5. The results shed some light on the decomposition of the variance of the domestic price into foreign disturbances (WOP, GRGDP, GNM) as opposed to domestic disturbances (FRGDP, FCPI) or (IRGDP, ICPI). According to the forecast error variance decomposition, real and nominal shocks originating from Germany account for a large share of price variability in France and Italy. Specifically, the proportion of price variance in France associated with German shocks ranged from 4.13 percent after two quarters to 18.84 percent after 24 quarters. In Italy, the proportion of price variance associated with German shocks ranged from 0.69 percent after two quarters to 18.45 percent after 24 quarters.

Real oil price shocks and domestic price shocks appear to be important determinants of price variability in France and Italy. Specifically, the proportion of price variance in France associated with domestic real and nominal shocks ranged from 83.46 percent after two quarters to 66.26 percent after 24 quarters. In Italy, the proportion of price variance associated with domestic real and nominal shocks ranged from 52.54 percent after two quarters to 19.84 percent after 24 quarters. In addition, real oil price shocks explain a share of price variance that ranges from 12.42 percent after two quarters to 14.89 percent after 24 quarters in France. The inflationary share of an increase in the world oil price is even more pronounced of the price variance in Italy. Specifically, this share ranges from 46.78 percent after two quarters to 61.71 percent after 24 quarters. Overall, it is interesting to observe the increased share of foreign shocks of price variability in France and Italy over time. That is, the spillover disinflationary effect of German shocks appears long-lasting on price in France and Italy. In addition, the reduction in the world oil price appears to have long-lasting effect on disinflation in Italy and, to a lesser extent, France.[11]

IV Conclusion

In this paper, we use a vector autoregressive model to evaluate the importance of different shocks on the inflation rate in France and Italy between 1979 and 1993. We classify the shocks as real world oil price shocks, real and monetary shocks originating from Germany, and real and nominal domestic shocks. We use two approaches to impose long-run constraints in order to guarantee the exogeneity of the shocks in the model. The first approach imposes theoretical structural constraints

[11]The results are robust with respect to a change in the order of variables in the VAR model. These results are available upon request.

that govern the long-run relations between shocks in the model. The second approach is atheoretical based on the results of cointegration among the trends of the variables in the model. The results of both models indicate that the effect of the shocks originating from Germany had a decreasing impact on price inflation in France and Italy. In addition, the weight of German shocks in explaining the variance of the inflation rate is large in both countries. This weight appears significantly larger in the structural VAR model which isolates domestic shocks in France and Italy from their correlation with German shocks to measure the contribution of the exogenous component of domestic shocks to disinflation.

Our results support, therefore, the view which stresses the importance of the spillover effect from Germany in decreasing inflation in traditionally high inflation countries of the European Monetary System. Nonetheless, domestic policies that aim at expanding output growth and curbing nominal shocks are also important components of the success of disinflation policies in France and Italy. Coordination between policies is evident by the change in the contribution of domestic and German shocks to price variability upon accounting for structural correlations between the shocks in the VAR model. The evidence, therefore, supports the spillover effects implied by the German dominance in the European Monetary System. Nonetheless, the coordination of domestic policies among countries of the European Monetary System remains necessary towards achieving a fast reduction in price inflation. Finally, lower world oil prices cannot be overlooked in evaluating the success of disinflation policies in France and Italy during the eighties and early nineties.

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Table 1: Unit Root Test Results for First-Differences

Variable	$\tau(.)$	Φ_3	$Z(\Phi_3)$	$Z(t\alpha^{\sim})$	$Z(\alpha^{\sim})$
WOP	-4.8660*	25.6756*	27.4537*	-7.3788*	-40.8456*
GRGDP	-5.2103*	33.0988*	32.8402*	-8.0951*	-63.4733*
GNM	-4.0629*	23.1935*	22.7404*	-6.7431*	-47.0488*
FRGDP	-4.5313*	17.9067*	19.1016*	-6.1810*	-52.2344*
FCPI	-5.4380*	5.3367[a]	4.9292	-3.1348[a]	-17.5578*
IRGDP	-5.4107*	18.4182*	13.0693*	-5.3061*	-36.2141*
ICPI	-3.8993*	7.9683*	7.9310*	-3.9826*	-25.9711*

Notes:

- $\tau(.)$ is the t-value according to the Augmented Dickey-Fuller Test for the null hypothesis of nonstationarity. The model specification for the test employs a constant, a time trend variable, and four lags of the dependent variable.

- We also use the Perron, and Phillips Perron tests. For these tests, we use a model with a constant and a time trend, i.e.,

$$X_t = \mu^{\sim} + \beta^{\sim}(t - n/2) + \alpha^{\sim} X_{t-1} + \epsilon_t, \quad t = 1, 2, ..., n.$$

  n is the number of observations. In this model, the null hypothesis, $H_0 : \beta^{\sim} = 0, \alpha^{\sim} = 1$ is tested by using Φ_3 and $Z(\Phi_3)$, and $H_0 : \alpha^{\sim} = 1$ is tested by using the test statistic $Z(t\alpha^{\sim})$ and $Z(\alpha^{\sim})$. The critical values for Φ_3 and $Z(\Phi_3)$ are given in Dickey and Fuller (1981, p. 1063) and the critical values for $Z(t\alpha^{\sim})$ and $Z(\alpha^{\sim})$ are given in Fuller (1976, p. 37 and 373).

- * indicates that the null hypothesis of non-stationarity is rejected at the ten percent level.

- [a] denotes statistical significance at approximately 12%.

- WOP is the real world oil price; GRGDP is the real gross domestic product in 1990 constant prices of Germany; GNM is the seasonally adjusted narrow money of Germany; FRGDP and IGRDP are real gross domestic products in 1990 constant prices for France and Italy; and FCPI and ICPI are the consumer price index for France and Italy.

Table 2: Test Results for the Determination of the Lag Length in the VAR Model

Null Hypothesis	Alternative Hypothesis	Acceptance France	Probability Italy
4 lags	8 lags	0.999	0.999
4 lags	6 lags	0.658	0.860
2 lags	4 lags	0.003	0.699
3 lags	4 lags	0.007	0.334

Notes:

- Acceptance probability is based on the Chi-square distribution for the likelihood ratio test.
- Following the suggestions of Sims (1980, p. 17), we take into account small sample bias by correcting the likelihood ratio statistic by the number of parameters estimated per equation. Thus, the likelihood ratio test=$T - C\{log]\Sigma_0] - log]\Sigma_1]\}$, where Σ_0 and Σ_1 are the variance covariance matrices of the residuals estimated from a VAR model with a constant and the number of lags under the null and alternative hypotheses, respectively. T is the number of used observations and C is the number of variables in the unrestricted equations.
- The degree of freedom for the Chi-square test equal the number of restrictions implied by variation in the lag length.

Table 3: Test Results for the Number of Cointegrating Vectors in the VAR Model

Null Hypothesis # of Cointegrating Vectors	Likelihood France	Ratio Italy	5% Critical Value
none	87.3387*	104.1572*	68.52
at most 1	50.6362*	57.2226*	47.21
at most 2	24.2709	30.3616*	29.68
at most 3	6.8793	10.5783	15.41
at most 4	0.7599	1.7744	3.76

Notes:

- * denotes significance at the five percent level.

Table 4: Variance Decomposition of Price Based on VAR Estimation with Cointegration Constraints

Period Ahead	The WOP	VAR GRGDP	Model GNM	for FRGDP	France FCPI
1	9.643	0.572	1.034	6.134	82.613
2	12.416	2.138	1.990	8.305	75.154
5	20.209	9.252	4.589	9.847	56.101
10	22.884	13.824	4.415	15.870	43.004
24	14.894	13.077	5.762	28.378	37.886

Notes:

- Numbers represent the percentage of the variance of the nth-period ahead forecast error for price inflation in France and Italy that is explained by the variables in the VAR model.
- WOP is the real world oil price; GRGDP is the real gross domestic product in 1990 constant prices of Germany. GNM is the seasonally adjusted narrow money of Germany; FRGDP and IRGDP are real gross domestic products in 1990 constant prices for France and Italy; and FCPI and ICPI are the consumer price index for France and Italy.

Table 5: Variance Decomposition of Price Based on VAR Estimation with Cointegration Constraints

Period Ahead	The WOP	VAR GRGDP	Model GNM	for IRGDP	Italy ICPI
1	16.7643	2.8713	1.2690	0.0628	79.0325
2	34.3782	2.2957	1.2477	2.9172	59.1610
5	41.3261	4.6314	0.2598	7.9962	44.0891
10	35.1188	6.4542	0.2555	9.7694	48.4018
24	23.9132	12.9750	0.2739	10.0350	52.8026

Notes:

- Numbers represent the percentage of the variance of the nth-period ahead forecast error for price inflation in France and Italy that is explained by the variables in the VAR model.
- WOP is the real world oil price; GRGDP is th real gross domestic product in 1990 constant prices of Germany. GNM is the seasonally adjusted narrow money of Germany; FRGDP and IRGDP are real gross domestic products in 1990 constant prices for France and Italy; and FCPI and ICPI are the consumer price index for France and Italy.

Table 6: Variance Decomposition of Price Based on VAR Estimation with Structural Constraints

Period Ahead	The WOP	VAR GRGDP	Model GNM	for FRGDP	France FCPI
1	12.184	8.119	55.946	2.029	21.720
2	10.561	9.759	53.022	2.252	24.403
5	8.128	19.943	42.707	3.894	25.326
10	7.016	26.712	38.611	4.037	23.620
24	6.009	33.868	34.119	4.780	21.222

Table 7: Variance Decomposition of Price Based on VAR Estimation with Structural Constraints

Period Ahead	The WOP	VAR GRGDP	Model GNM	for IRGDP	Italy ICPI
1	9.695	14.652	11.271	10.961	53.419
2	26.443	22.734	5.797	21.535	23.488
5	28.135	32.455	1.847	12.640	24.921
10	28.462	32.281	2.441	10.036	26.778
24	25.152	38.740	2.977	5.928	27.199