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Inflation and the Skewness of the Distribution of Relative Price Changes: Empirical Evidence for Germany

by

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Abstract

The present paper uses German annual data covering the period 1969-2000 to present evidence on the link between aggregate inflation and the higher-order moments of the distribution of relative price changes. Our empirical findings confirm predictions of contributions to the theoretical literature suggesting that skewness of this distribution is an important explanatory variable for the inflation rate. Further, the skewness measure also helps to explain shifts in the Phillips curve. Moreover, a structural vector autoregression reveals that the skewness measure helps to explain the variations of real output and might, therefore, serve as a measure of supply side shocks hitting the economy.

JEL Classification: E 31

Keywords: inflation, relative prices, distribution of relative prices changes

Inflation und die Schiefe der Verteilung relativer Preisänderungen: Empirische Evidenz für Deutschland

Zusammenfassung

Der Aufsatz untersucht auf der Basis deutscher Jahresdaten für den Zeitraum 1969-2000 den Zusammenhang zwischen der gesamtwirtschaftlichen Inflationsrate und den höheren Momenten der Verteilung der relativen Preisveränderungen. Unsere empirischen Ergebnisse bestätigen die Vorhersagen theoretischer Ansätze, nach denen die Schiefe der Verteilung relativer Preisveränderungen eine erklärende Variable für die Inflationsentwicklung darstellt. Die Schiefe erlaubt es, Verschiebungen der Phillipskurve zu erklären. Empirische Untersuchungen unter Zuhilfenahme einer strukturellen Vektorautoregression zeigen, dass die Variation des Outputs teilweise mit Hilfe dieser Variablen erklärt werden kann. Die Schiefe relativer Preisveränderungen ist somit ein geeignetes Maß für Angebotsschocks, von denen die Volkswirtschaft getroffen wird.

Schlagworte: Inflation, relative Preise, Verteilung relativer Preisveränderungen

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I Introduction

A heavily discussed topic in the current macroeconomic debate is whether aggregate inflation as measured by changes in the average price level is influenced by the skewness of the distribution of relative prices. This debate has been initiated by Ball and Mankiw (1995). They have derived the theoretical link between the skewness of the distribution of relative prices and the average inflation rate within the context of a menu cost model. The core economic assumption of this model is that firms incur costs when adjusting prices. These so-called menu costs give rise to a band of inaction because firms only adjust prices when large shocks to relative prices occur. Small shocks, in contrast, drive a gap between firms' preferred and actually realized relative prices. The resulting band of inaction implies that only if relative price changes, which serve as a measure of supply side shocks, are symmetrically distributed around zero, actual upward and downward adjustments of prices cancel out each other and do not affect the aggregate price level on average. If, in contrast, the distribution of relative price changes is skewed, actual upward and downward adjustments of prices do not match on average. In this case, aggregate inflation rises (declines) whenever the distribution of relative price changes is skewed to the right (left). Thus, the sign of the supply side shocks, which require changes in relative price, is closely related to the sign of the skewness of the distribution of relative price changes: positive supply side shocks result in a left-skewed distribution of relative price changes, et vice versa.

The model suggested by Ball and Mankiw (1995) also implies that the variance of relative price movements affects inflation only insofar as it is positively related to the skewness of the distribution. A separate amplifying effect running from the second moment of the distribution of relative price changes to aggregate inflation arises only if the economic system exhibits a positive core inflation rate as in Ball and Mankiw (1994).

The points emphasized by Ball and Mankiw (1995) enrich the line of argumentation found in earlier studies, which have mainly analyzed the potential interplay between changes in the average price level and the variance of the distribution of relative price changes. In these studies, it has often been documented empirically that aggregate inflation is correlated with the cross-section variance of sectoral relative price changes (see, e.g. Fischer (1982) and Franz (1985)). Given that Ball and Mankiw (1995) have stressed that the skewness rather than the variance of the cross-section of sectoral relative prive changes is an important determinant of changes in the aggregate price level, their ideas have spawned an intensive debate among theoretical and empirical macroeconomists.

On the theoretical side, the menu cost view of the link between the third scaled moment of the distribution of relative price changes and inflation has been challenged by Balke and Wynne (2000).[1] They set up a multi-sectoral general equilibrium model with flexible prices to argue that a positive correlation between the skewness of relative price changes and the aggregate inflation rate is not a feature unique to menu cost models. The economic reasoning underlying their line of argumentation is built on the notion that in a flexible price model a positive sector-specific technology shock will entail a rise of the output of that sector and a fall of the relative price of that output. If a sufficient number of industries are buffeted by sectoral shocks of the same sign and some sectors are affected more than others, a flexible price model also allows a positive correlation between the mean and the skewness of the cross-section of the distribution of relative price changes to be generated. In other words, in such a model a positive correlation between the aggregate inflation rate and the skewness of the distribution of the cross-section of relative price changes merely arises because the underlying sectoral technology shocks are correlated. This result implies that an empirically observed positive correlation between skewness measures extracted from the cross-section of changes in relative prices and the aggregate inflation rate need not be caused by sluggishness in the adjustment of individual relative prices.

Even though the debate on the sources of a positive correlation between the skewness of the distribution of relative price changes and the aggregate inflation rate has not been settled so far, the results derived in the theoretical literature suggest that such a link should exist in real-world data. This contrasts with the

[1] In the remainder of the paper, we use the phrases ‘third scaled moment of the distribution of relative price changes’ and ‘skewness of the distribution of relative price changes’ interchangeably.

empirical evidence on this issue that is far more ambiguous. Evidence in favor of a significant positive link between the skewness of the distribution of relative price changes and the aggregate inflation rate has been reported by, e.g., Ball and Mankiw (1995) and Amano and Macklem (1997) for U.S. and Canadian producer price data sets, respectively. These results stand in contrast to the findings of, e.g., Holly (1997) who finds for Japanese producer prices that the skewness of the distribution of relative price changes does not exert a strong impact on the average inflation rate. Bryan and Cecchetti (1999) have even claimed that the link between the skewness of the distribution of relative price changes and aggregate inflation found by other researchers might be a statistical artifact. As laid out in detail below, this view has been refused by Ball and Mankiw (1999) and is still under debate.

The ambiguity of findings reported in existing empirical studies motivates the analyses contained in the present paper. We offer an additional piece of information regarding the empirical relevance of the Ball-Mankiw hypothesis and establish thereby an important stylized fact of the German business cycle. To this end, we analyze whether a positive link between the skewness of the distribution of relative price changes and the aggregate inflation rate can be established for German data. We find that the prediction of the theoretical literature that skewness has a significant impact on the inflation process cannot be rejected. Moreover, we present evidence that the skewness of the distribution of relative price changes allows shifts of the Phillips curve in Germany to be explained. In such a Phillips curve model, the skewness measure serves as a proxy for supply side disturbances hitting the economy. This result supports the notion that the instability of the Phillips curve relation in Germany can be attributed in part to supply side disturbances.

Our analysis is organized as follows. We start in Section 2 with a review of the existing empirical evidence on the link between the skewness of the distribution of relative price changes and aggregate inflation. In section 3, we present the data we use in our quantitative investigations. We analyze the properties of the cross-sectoral variance and skewness of the distribution of relative price changes and discuss whether the skewness of this distribution provides an economically reasonable measure of supply side shocks. The results of our empirical analyses are contained in Section 4. The final section offers some concluding remarks.

II Existing Empirical Evidence

A systematic empirical study of the link between the third scaled moment of the distribution of relative price changes and the aggregate inflation rate has been carried out by Ball and Mankiw (1995).[2] Utilizing annual producer price data for the United States for the second half of the twentieth century, they present regression-based evidence for a positive causal relation running from the skewness of the distribution of relative price changes to the aggregate inflation rate. Moreover, they show that the aggregate inflation rate is affected by the variance of the distribution of relative price changes only through the interaction of the latter with the skewness of the distribution of relative price changes. Ball and Mankiw (1995) further demonstrate that the importance of the third scaled moment of the distribution of relative price changes for the dynamics of aggregate inflation also obtains in the context of estimated Phillips curves. To this end, they set up Phillips curve equations that include relative prices of oil and food and the skewness of the distribution of relative price changes as competing measures of supply side shocks. In these regressions, the oil and food price coefficients are not significantly different from zero. In contrast, the coefficient comprising the influence of the third scaled moment of the distribution of relative price changes on the dynamics of the aggregate inflation rate is statistically significant and positive.

These results have been criticized by Bryan and Cecchetti (1999). They argue that the statistically significant positive impact of the skewness of the distribution of relative price changes on the aggregate inflation rate documented by Ball and Mankiw (1995) is due to a small sample bias. Their argument rests on the notion that even when a statistician has obtained a zero-mean sample from a zero-mean symmetric distribution, an additional single influential draw from the

[2] Early empirical evidence on the link between the moments of the distribution of relative price changes and the dynamics of aggregate inflation is reported in, e.g., Vining and Elwertowski (1976).

extreme positive tail of the distribution should make both the sample mean and the sample skewness positive and results, by construction, in a positive correlation between these statistics. To compute the exact magnitude of such a potential small sample bias, Bryan and Cecchetti (1999) resort to numerical simulation techniques. They interpret the results of the simulations as evidence suggesting that the magnitude of the small sample bias may account for the positive sample correlation of the skewness of the distribution of relative price changes and the aggregate inflation rate observed by Ball and Mankiw (1995).

Ball and Mankiw (1999) have objected to this interpretation of their results. They argue that a small sample bias of the type observed by Bryan and Cecchetti (1999) could only arise in a model in which the distribution of observed real-world relative price changes is computed by drawing a subset of price changes from an underlying true distribution of relative price changes. However, governments measure prices in *all* sectors of an economy. This leads Ball and Mankiw (1999) to conclude that the correlation of the aggregate inflation rate and the skewness of the distribution of relative price changes they documented in their empirical work is based on the full population of relative price changes across all sectors of the economy. Hence, by construction, a small *sample* bias cannot arise.

Ball and Mankiw (1999) further criticize that Bryan and Cecchetti (1999) do not derive the design of their experiments from economic theory. The experiment used by Bryan and Cecchetti (1999) a framework implies that sectoral price changes reflect random draws from a symmetric distribution that transmit directly into the aggregate price level. This constitutes a deviation from the classical model in which monetary policy determines the aggregate price level whereas relative prices are determined in the real sphere of the economy. Ball and Mankiw (1999), therefore, argue that the results of the simulation experiments "...offer a statistical "explanation" for the observed inflation-skewness correlation without any obvious interpretation." (p.98).

A simple research strategy to control for potential estimation problems caused by the small sample bias emphasized by Bryan and Cecchetti (1999) has been suggested by Amano and Macklem (1997). They argue that a small sample bias such as the one described by Bryan and Cecchetti (1999) can only arise in a regression equation in which the set of regressors consists of the higher-order oments of the distribution of sectoral relative producer price changes and the regressand is given by an aggregate inflation rate computed from the cross-section of the same producer price changes. If an aggregate inflation rate as measured by changes in the gross domestic product (GDP) deflator or in the consumer price index (CPI) is placed as the dependent variable on the left-hand side of the regression equation instead, then a small sample bias cannot arise unless the third scaled moment of the distribution of relative producer price changes is systematically correlated with the rate of change of the GDP deflator or the CPI.

Amano and Macklem (1997) have applied this research strategy to Canadian producer price data for the period 1962-1994. They find a robust positive correlation between the skewness of the distribution of relative price changes and the aggregate inflation rate. This result also holds in Phillips curve regressions that allow for the influence of other determinants of inflation dynamics to be controlled for. Furthermore, they report that not only the skewness but also the variance of the distribution of relative price changes affects inflation. This is consistent with the model of Ball and Mankiw (1994) as the inflation rate exhibited a trend component in Canada during the sample period.

In an international context, several authors have presented empirical evidence on the Ball–Mankiw hypotheses.

Empirical evidence for Japan has been reported by Holly (1997) based on quarterly producer price data covering the sample period 1976-1994. The results of his analyses are mixed. On the one hand, he estimates Ball–Mankiw style least square regressions and finds significant explanatory power of the skewness of the distribution of relative price changes for the aggregate inflation rate. On the other hand, the implementation of causality tests and the estimation of a structural model that additionally captures potentially important contemporaneous links between the aggregate inflation rate and the moments of the distribution of relative price changes reveals that inflation tends to cause fluctuations in the variance and the skewness of this distribution rather than the other way round.

Silver and Ioannidis (1996) have presented results on the extent of the skewness of the distribution of relative price changes across nine European countries including Germany. However, they are mainly concerned with the impact of the aggregate inflation rate on the third scaled moment of the distribution of relative price changes. Thus, though they analyze the potential link between the skewness of the cross-section of relative price changes and the aggregate inflation rate, they do not test the hypotheses formulated by Ball and Mankiw (1995). In addition, the data used in the paper cover the period 1982-1990 only. Hence, their sample period does not include the years of the two pronounced oil price shocks that buffeted the Western world in 1974 and in 1980. These oil price shocks are often seen as textbook examples for significant negative supply side shocks. Therefore, the analysis conducted by Silver and Ioannidis (1996) provides only limited information on the usefulness of skewness of a measure of supply side shocks.

Using monthly store-level data for Hungary for 1992-1996, Ratfai (2000) has estimated various bivariate structural vector autoregressive models containing producer price inflation and the skewness of the distribution of relative price changes. He tests whether the skewness of the distribution of relative price changes has a significant explanatory power for the dynamics of average product price inflation. His findings suggest that product-specific price shocks explain a substantial fraction of the forecast error variance of producer price inflation. In addition, he finds that increases in sectoral inflation tend to be preceded by product-specific shocks.

Bonnet et al. (1999), employing annual data for France covering the period 1960-1996, have reported estimates that point to the existence of a positive correlation between the skewness of the distribution of relative price changes and the aggregate inflation rate. They further report that the positive correlation between skewness and inflation is also found when the influence of other variables driving the French inflation process is taken into consideration. [3]

Further evidence supporting the stylized fact emphasized by Ball and Mankiw (1995) that the aggregate inflation rate and the skewness of the distribution of relative price changes are positively correlated has been reported by De Abreu Lourenco and Gruen (1995) using bi-annual data for Australia covering the sample period 1970-1992. They stress that this correlation may also depend upon the level of expected aggregate inflation. Controlling for the influence of this variable, they find that skewness has always a significant impact on the inflation rate regardless of the level of aggregate inflation. The variance of the cross-section of relative prices, in contrast, exerts only a significant influence on the aggregate inflation rate in a regime characterized by a high level of aggregate inflation.

III Data Exploration

To compute the aggregate inflation rate and the moments of the distribution of relative price changes, we use the German producer price index and its sub-indices. The number of sub-indices available to compute the higher-order moments of the distribution of relative price changes varies between 28 and 32 because the definition of the sub-indices has changed during the sample. Time series derived from a uniform definition of the sub-indices covering the entire sample period were not available. A detailed description of the data employed in the subsequent empirical analyses can be found in the data appendix at the end of the paper. The data are sampled at an annual frequency. The sample period begins in 1969 and ends in 2000.

Equipped with this data set, we can compute the sample moments of the distribution of relative price changes. Let the realization of the price index in sector i observed in period t be denoted by $p_{i,t}$. Using this notation, we define the change $\pi_{i,t}$ of the realization of sub-index i over the previous year as:

[3] Additionally, they argue that their results are in line not only with menu cost models but also with models that assume that the tails of the price distribution is fatter than the tail of a normal distribution.

[1] $$\pi_{i,t} = \ln(p_{i,t} / p_{i,t-1}) \ .$$

Following Bryan and Cecchetti (1999), the aggregate inflation rate π_t in period t can then be computed as the weighted average of the changes of the sub-indices:

[2] $$\pi_t = \sum_{i=1}^{N} \omega_i \pi_{i,t} \ ,$$

where N denotes the number of sectors in the sample and the symbol ω_i represents the weight attached to the price change in sector i. The calculation of the weights is described in the appendix at the end of the paper.

With the aggregate inflation rate π_t at hand, it is now possible to compute the higher-order central moments of the distribution of relative price changes. The cross-sectional variance csv_t of this distribution is given by:

[3] $$csv_t = \sum_{i=1}^{N} \omega_i (\pi_{i,t} - \pi_t)^2 \ .$$

The cross-sectional skewness css_t is defined as the third scaled moment of the distribution of relative price changes:

[4] $$css_t = \left[\sum_{i=1}^{N} \omega_i (\pi_{i,t} - \pi_t)^3 \right] \times csv_t^{-\frac{3}{2}} \ .$$

Coding up these recipes, we computed the aggregate producer price inflation rate depicted in Figure 1 and the skewness and the variance of the distribution of relative price changes graphed in Figure 2 and 3, respectively. We calculated the aggregate inflation rate (p_t) and higher order central moments (csv_t and css_t) both with equal weighting of all prices (so that $\omega_i = 1/N$ for all i) and with the sectoral weights described in the data appendix. This makes it possible to shed light on the influence of the weighting factors g_i on the aggregate inflation rate and on higher-order central moments of the distribution of relative price changes.

Figure 1 — Producer Price Inflation in Germany (1969-2000)

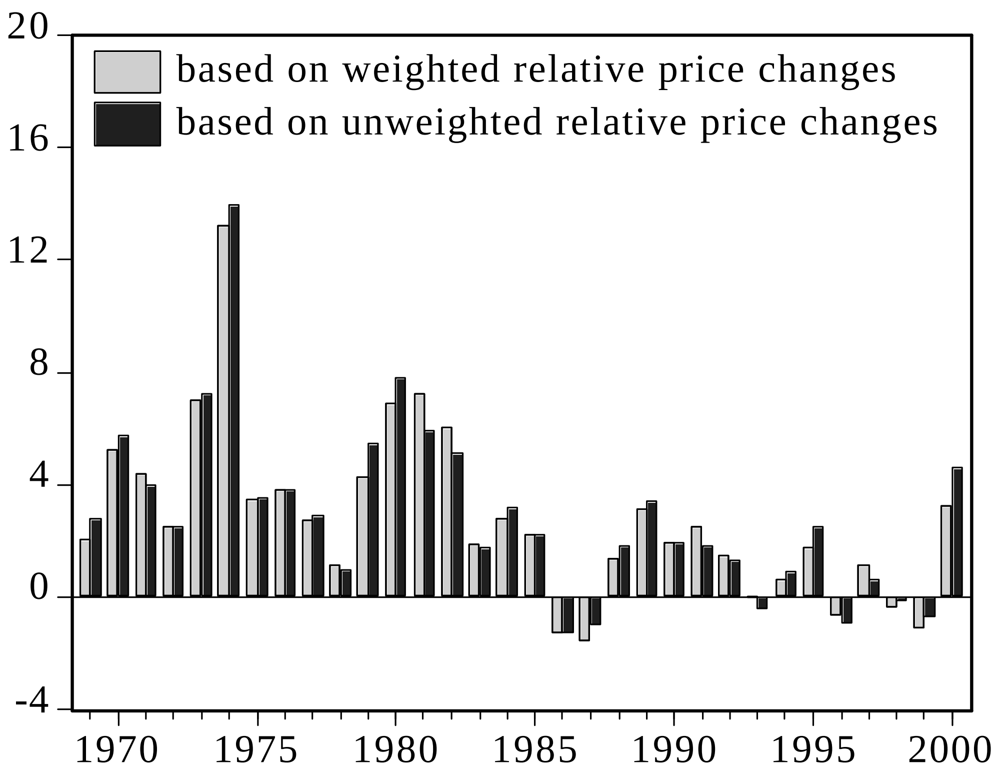
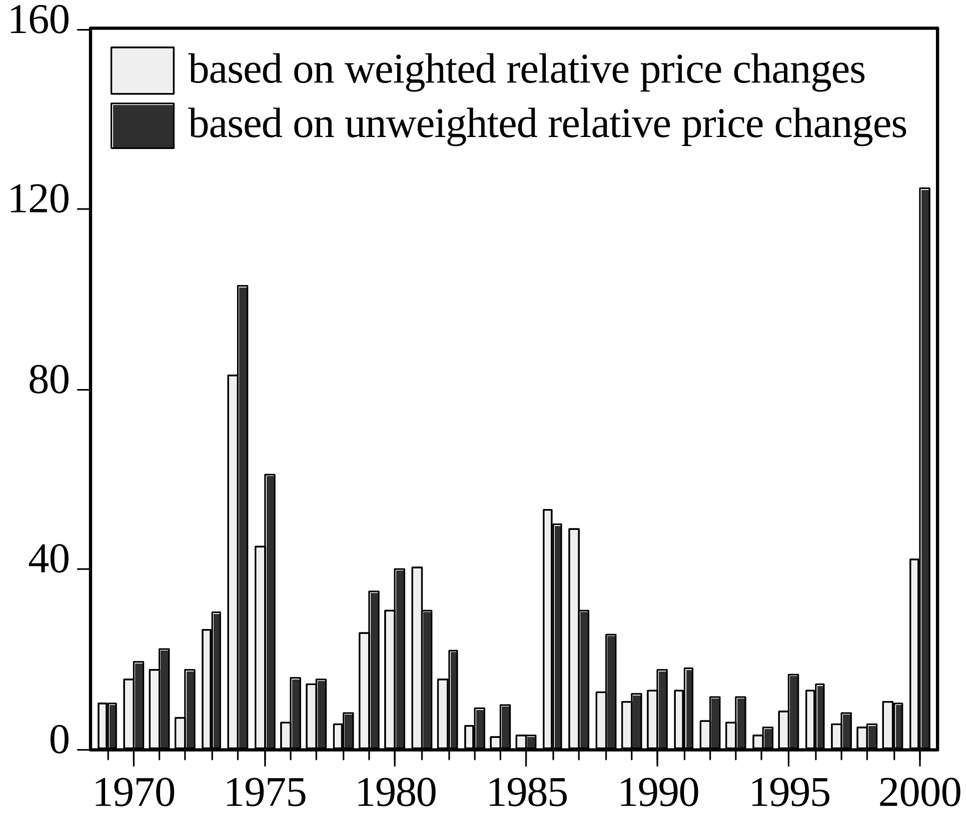


Figure 1 shows that for the data set used in this study the assumption underlying the analysis of Ball and Mankiw (1995) that the core inflation rate is approximately equal to zero can hardly be justified empirically. Instead, the average producer price inflation was remarkably high in the mid-seventies and in the early eighties. These periods were characterized by the two oil price shocks. Relatively modest negative producer price inflation rates were realized in 1986 and 1987 and in a period of time beginning in the mid-nineties. Overall, however, the evidence presented in Figure 1 suggests that the core inflation rate in Germany was positive during the sample period analyzed. This, in turn, implies that the variance of the distribution of relative price changes can be expected to influence the aggregate inflation rate not only through its interaction with the skewness of the distribution of relative price changes. Instead, following the analysis of Ball and Mankiw (1994), a direct amplifying effect running from the variance of the distribution of relative price changes to the aggregate inflation rate may exist.

Figure 2 shows the dynamics of the second central moment of the distribution of relative price changes. Not surprisingly, this cross-sectional variance of relative producer price changes shows peaks in the years of the two sharp oil price increases in 1973/74 and in 1979/80 as well in the years 1986/87 which were characterized by a pronounced drop of oil prices. A further result shown in the graph is that the variance of relative producer price changes does not correspond to the flat and even negative aggregate producer price inflation rates

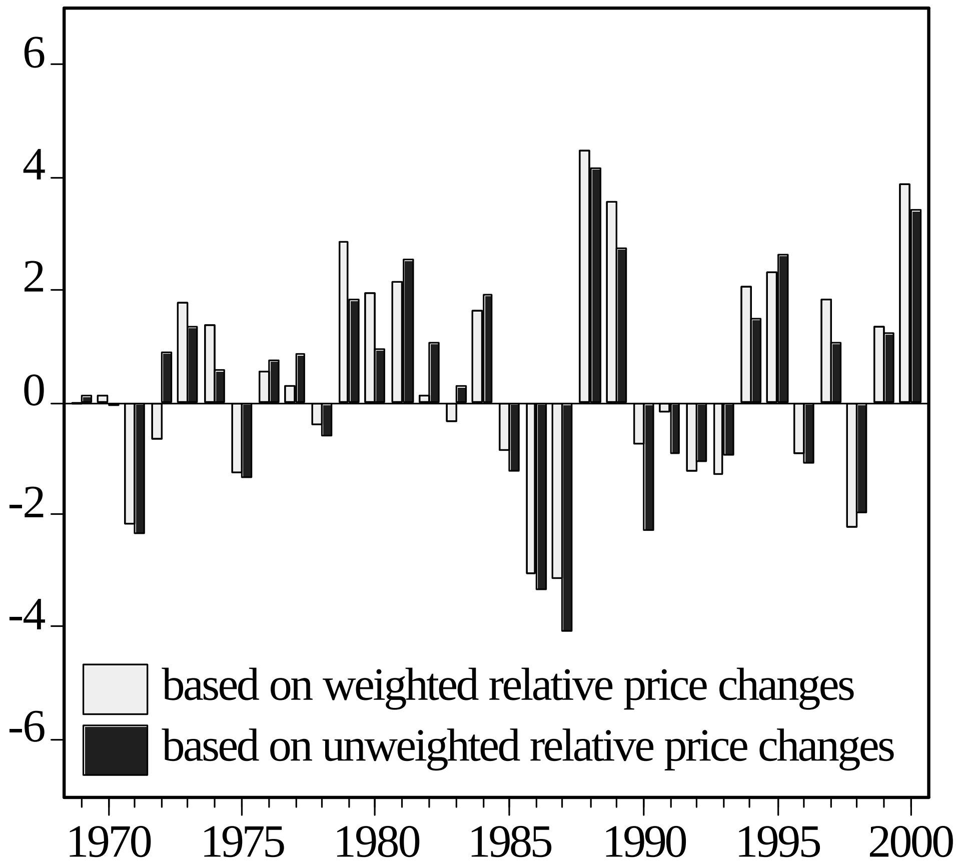
Figure 2 — Variance of the Distribution of Relative Price Changes in Germany (1969-2000)



realized in the nineties. This suggests that the dynamic evolution of relative producer prices in the mid-nineties was relatively similar across sectors.

The third scaled moment of the distribution of relative price changes is given in Figure 3. The graph shows that this skewness was positive in the mid-seventies. Yet, given the pronounced increases in oil prices that took place in these years, the positive skewness depicted in the figure is rather small. This contrasts with the widespread belief that the aggregate inflation rate in the early seventies was mainly driven only by sharp increases in oil prices. In fact, a closer look at the cross-section of relative price changes revealed that the inflationary process during this period of time was fostered by price increases in almost all sectors of the German economy. This contrasts with the economic situation at the beginning of the eighties and in 1989 and in 1999/2000. These periods of time time witnessed isolated rises of energy prices that led to a pronounced positively skewed distribution of relative price changes.

Figure 3 — Skewness of the Distribution of Relative Price Changes in Germany (1969-2000)



To gain further insights into the dynamics of the distribution of relative price changes, we ran kernel regressions to obtain (for each year in the sample) a kernel density estimate of the distribution of relative price changes. A kernel density estimate can be interpreted as a smoothed relative frequency histogram of a time series. A conventional step-wise linear histogram is computed by figuring out the relative frequencies of equally weighted observations found within intervals of arbitrary width equidistantly spaced between the minimum and the maximum of an analyzed data set. In a kernel regression, in contrast, a so-called kernel or smoothing function is utilized to transform the discontinuous histogram into a continuous real-valued function.

To elucidate the idea behind this estimation technique in more detail, let the mapping $\pi_{i,t} \mapsto G_{RF}$ relating relative price changes to the corresponding relative frequencies G_{RF} with which these price changes are observed be defined by the unknown non-linear function $g_K(\pi_{i,t})$. In a kernel regression, an estimate $\hat{g}_K(\pi_{i,t})$ of this function for any arbitrary realization $\pi_{i,t}$ of the cross-section of relative price changes observed in period t is calculated as the weighted arithmetic average of the local intensities of observations found in a small neighborhood around $\pi_{i,i}$ (see, e.g., Härdle 1990: 19):

[5] $$\hat{g}_K\left(\pi_{i,t}\right) = \frac{1}{T}\sum_{i=1}^{N} \varsigma\left(i, N, \pi_{i,t}\right) G_{RF} \quad ,$$

where the magnitude of the weighting factors $\varsigma\left(i, N, \pi_{i,t}\right)$ declines as the distance between $\pi_{i,t}$ and the relative price changes found in a band of width w spaced symmetrically around this reference value increases.

In a kernel regression, the concrete functional form of the weighting factors $\varsigma\left(i, N, \pi_{i,t}\right)$ is determined by specifying a continuous kernel density function. The kernel density function is a non-negative, continuous, bounded, symmetric real valued function that integrates to one. In the subsequent analyses, the so-called Epanechnikov kernel function was used (see, e.g., Härdle 1990: 25). The kernel function is used to re-express the weighting factors. The resulting kernel weights are plugged into equation [5] so that the Nadayra–Watson kernel density estimator obtains (see Härdle 1990: 25).

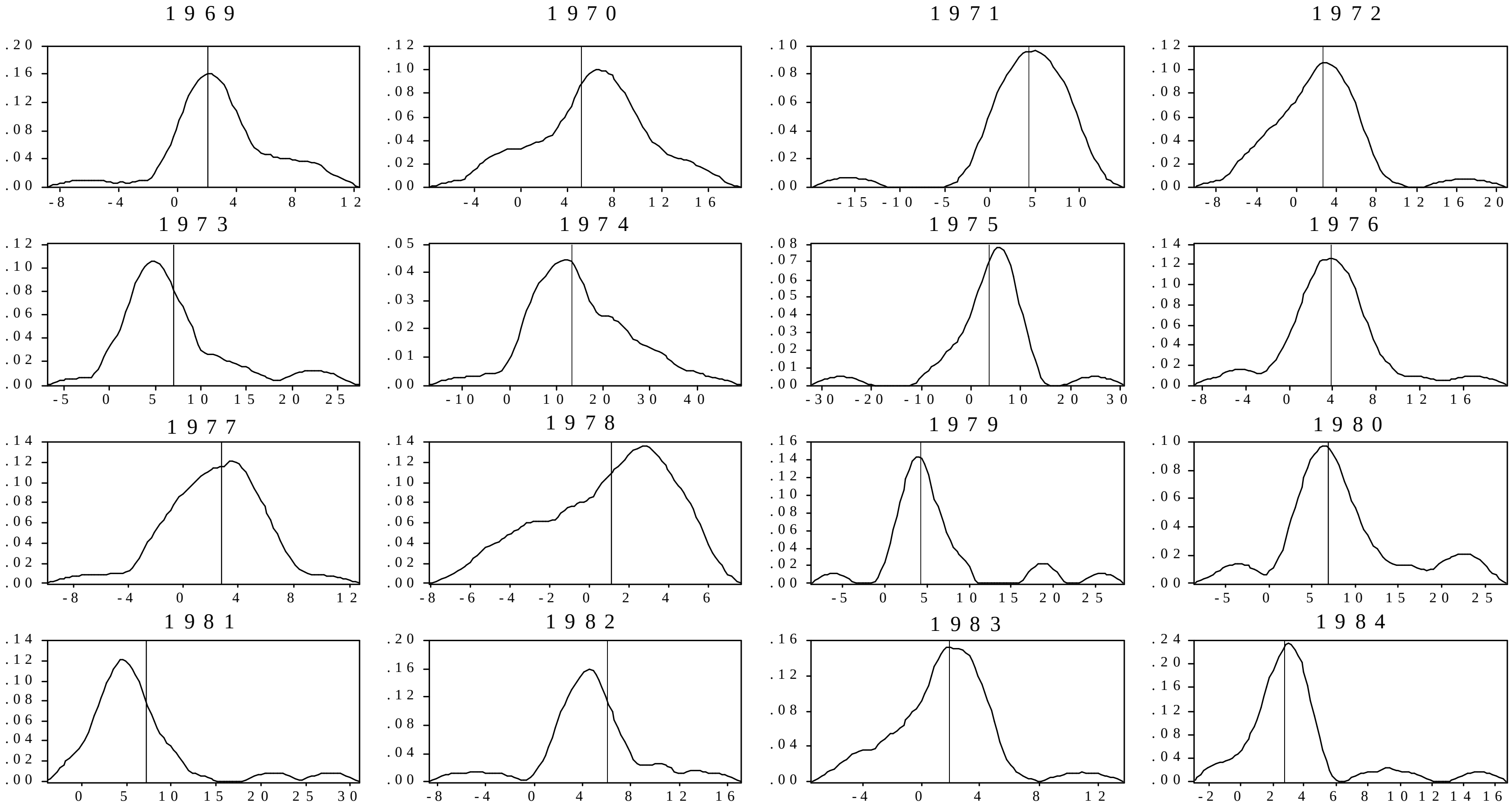
The smoothness of the estimated kernel density rises as the numerical value ascribed to the bandwidth parameter w is raised. Thus, choosing the parameter w too large can result in an excessively smooth kernel density estimate overfitting the data and washing out the inherent non-linearity of the function $g_K(\pi_{i,t})$. To tackle this problem, several approaches allowing to select an optimal bandwidth parameter for a given kernel function have been suggested in the literature. We employed the approach advocated by Silverman (1986: 47-48) to specify the bandwidth parameter.

The estimated kernel density functions are plotted in Figure 4. The figure shows that the shape of the kernel density estimates of the distribution of relative price changes varied substantially over the years. Moreover, it can be seen that pronounced right-skewed kernel densities correspond to years in which significant negative supply shocks buffeted the German economy. For example, consider the shape of the kernel density estimate of the cross-section of relative price changes in the years 1973/74 and 1979-81 in which the first and second oil price shocks occurred, respectively. The kernel density estimated for the year 2000 represents a further example for a right-skewed distribution shaped by a sharp increase in energy prices. Comparing the skewness of this function with the skewness of the kernel densities realized during earlier oil price shocks suggests that a pronounced negative supply shock hit the German economy at the end of the millennium.

Examples for left-skewed kernel densities shaped by positive supply side shocks, can be found in the figures describing the situation in Germany in 1986/87 and, to a lesser extent, in 1998.

All in all, the above discussion suggests that the skewness of the distribution of relative price changes is a reliable measure of shocks originating on the supply side of the economy. Therefore, further investigations of the interplay between important macroeconomic aggregates and the moments of the distribution of the cross-section of relative price are worthwhile.

Figure 4 — Kernel Estimates of the Distribution of Relative Prices in Germany (1969 – 2000)



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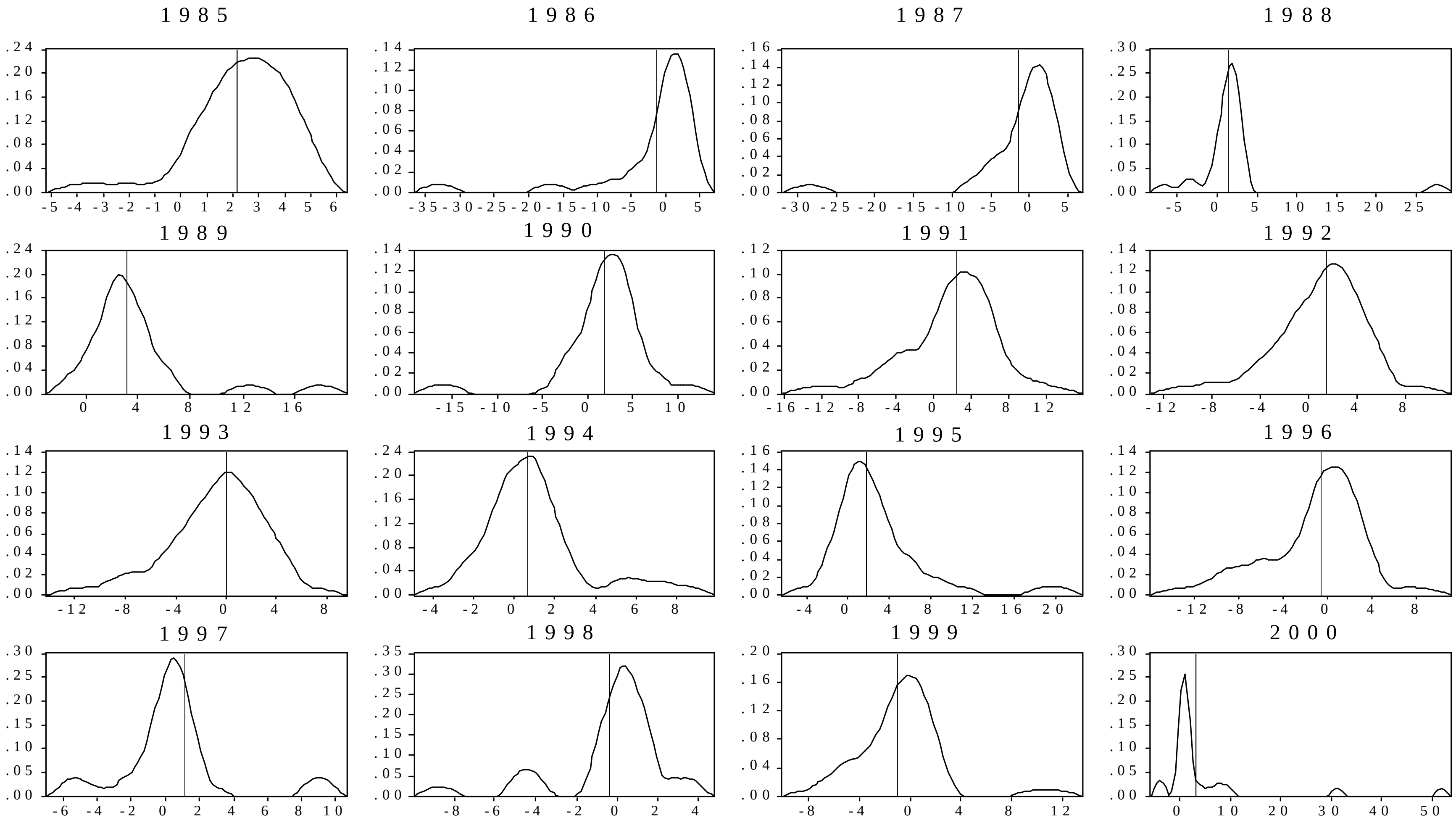
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IV Empirical Results

We now turn to more formal analyses of the impact of the higher-order moments of the distribution of relative price changes on the average inflation rate. For this purpose, we employ several regression-based framework as advocated by Ball and Mankiw (1995). Additionally we estimate Phillips curve equations and structural vector autoregressions (VARs) to examine further whether the skewness of the distribution of relative price changes can be used as a measure of supply side shocks. The Phillips curve equations allow the link between the aggregate inflation rate, the unemployment gap and css_t to be analyzed. The structural VAR contains the change in real GDP and css_t as endogenous variables. We supplement the results obtained from implementing this approach with the findings from a structural vector autoregressive model.

4.1 Regression Based Evidence on the Link Between Inflation and Skewness

We estimated regression equations in the tradition of Ball and Mankiw (1995). Equation (6) presents a formal representation of the econometric framework:

[6] $$\pi_t = \beta_0 + \beta_1 \pi_{t-1} + \beta_2 csv_t + \beta_3 css_t + \varepsilon_t \ ,$$

where β_i, $i = 0,1,2,3$ are regression coefficients and ε_t denotes a serially uncorrelated normally distributed disturbance term. The coefficient β_3, which captures the influence of the skewness variable css_t on the aggregate inflation rate π_t, should be significantly different from zero and positive if the theoretical work of Ball and Mankiw (1995) and Balke and Wynne (2000) provides an appropriate analytical tool for studying the sources of the German business cycle. The study of Ball and Mankiw (1994) further suggests that the positive aggregate core inflation rate we detected in our data should result in a significant and positive coefficient β_2.

We estimated equation (6) using ordinary least squares. The results of the estimations are given in Tables 1 and 2. The results given in Table 1 were computed using unweighted measures of the moments of the distribution of relative price changes. Results based on the corresponding weighted measures are depicted in Table 2. As can be seen in the tables Equation (6) was estimated in various specifications. Column (1) of the tables gives the results for a benchmark regression with a constant and the lagged aggregate inflation rate as regressors. To generate the results plotted in Column (2), we added to the set of regressors used in the benchmark regression the variance of the distribution of relative price changes. Column (3) shows the estimation results for a regression including a constant, the lagged aggregate inflation rate, and the skewness of the distribution of relative price changes. Finally, Column (4) gives results for a quantitative framework including both the variance and the skewness of the distribution of relative price changes as regressors.

Table 1 — Inflation and the Distribution of Price Changes in Germany (1969-2000)

Regressors: Unweighted Measures of Higher-Order Moments				
	(1)	**(2)**	**(3)**	**(4)**
constant	1.5576	-1.2522	1.1628	-1.1536
	(2.5340)***	(-1.3994)	(2.3694)**	(-1.2644)
π_{t-1}	0.4752	0.3322	0.5423	0.4113
	(3.6299***)	(1.6244)	(4.3361)***	(2.2461)**
csv_t		0.6947		0.5874
		(2.2001)**		(1.9444)*
css_t			0.7356	(0.6251)
			(4.4826)***	(4.4320)***
R^2	0.2236	0.4221	0.4267	0.5641
BG	0.1731	0.0147	0.0538	0.5745
JB	10.06***	0.8160	24.4541***	1.7622

Notes: The figures in brackets under the coefficients are *t*-ratios which were calculated using robust standard errors. Robust standard errors were computed upon implementing the method of Newey and West (1987). *BG* denotes a Breusch-Godfrey test on autocorrelation of first order (*F*-value). *JB* denotes a Jarque-Berra test on normality of the residuals. *** (**, *) denotes rejection of the null hypothesis at the 1 (5, 10) percent level.

Table 2 — Inflation and the Distribution of Price Changes in Germany (1969-2000)

Regressors: Weighted Measures of Higher-Order Moments	**(1)**	**(2)**	**(3)**	**(4)**
constant	1.3483	-0.5310	0.5908	-1.0271
	(2.3151)**	(-0.4451)	(1.2888)	(-1.4417)
π_{t-1}	0.5253	0.4187	0.6648	0.5649
	(3.8122)*	(1.9210)*	(4.7729)***	(3.0070)***
csv_t		0.5516		0.4848
		(1.1577)		(1.8207)*
css_t			0.8335	0.7964
			(4.3642)***	(5.2751)***
R^2	0.2762	0.3825	0.5422	0.6238
BG	0.2580	0.0387	0.0507	0.2212
JB	6.9751***	1.3749	13.6083***	2.0040

Notes: The figures in brackets under the coefficients are *t*-ratios which were calculated using robust standard errors. Robust standard errors were computed upon implementing the method of Newey and West (1987). *BG* denotes a Breusch-Godfrey test on autocorrelation of first order (*F*-value). *JB* denotes a Jarque-Berra test on normality of the residuals. *** (**, *) denotes rejection of the null hypothesis at the 1 (5, 10) percent level.

As predicted by economic theory, the skewness of the distribution of relative price changes turns out to exert a significantly positive effect on the average inflation rate. The respective *t*-ratios show that the coefficients are significant at the one percent level. Furthermore, the fit of the equation is satisfactory and is improved by adding the skewness variable to the set of regressors. The empirical results given in Table 1 and 2, thus confirms the predictions of the theoretical frameworks developed by Ball and Mankiw (1995) and Balke and Wynne (2000).

An additional piece of information provided by the estimation results reported in the tables is that the second moment of the distribution of relative price changes also affects the aggregate inflation process. The coefficients exhibit always the sign predicted by economic theory. Yet, as compared to the skewness, the impact of the variance on the aggregate inflation rate is less strong. Nevertheless, the results support the finding of Ball and Mankiw (1994) that the variance of the distribution of relative prices influences the dynamics of the aggregate inflation rate directly as long as the core inflation rate is positive. For Germany, this latter assumption is clearly justified for the sample period under investigation.[4]

A potential problem is that in the equation that contains the skewness measure as the only additional explanatory variable the residuals are not normally distributed. Visual inspection of the residual series suggested that the deviation from normality is mainly due to one influential outlier in 1974. We therefore reestimated the equation using a reduced sample beginning in 1975. We found this a change in the sample period does not change the results presented in the third columns of Table 1 and 2 qualitatively.

A further problem that might arise in the interpretation of the results documented in Tables 1 and 2 is that our findings may be distorted due to the small sample bias mentioned by Bryan and Cecchetti (1999). To take this criticism into account, we adapted the estimation strategy suggested by Amano and Macklem (1997). They argue that the small sample bias mentioned by Bryan and Cecchetti (1996) can only arise in a regression equation in which the aggregate inflation rate calculated from the cross-section of producer price changes is regressed on the moments of the distribution of sectoral relative producer price changes. If, in contrast, the aggregate inflation rate is measured in terms of an alternative deflator not constructed from the distribution of relative producer price changes then a small sample bias is unlikely to beleaguer the estimation results.

In contrast to Amano and Macklem (1997), we did not use the GDP deflator to replace the aggregate producer price inflation rate as the dependent variable of the regression equation (6). The economic reasoning motivating our decision is that the GDP deflator also contains the deflator of imports. Thus, the GDP deflator shows a tendency to decline whenever negative supply side shocks caused, for example, by a rise in oil prices buffet the economy. The GDP deflator is therefore often blamed to be a misleading indicator of inflationary pres-

[4] One might wonder whether the data provide support for the prediction of the model of Ball and Mankiw (1995) that the variance term may affect the aggregate inflation rate through its interaction with the skewness of the distribution of relative price changes. In their seminal paper, Ball and Mankiw (1995) have used the product of the standard deviation and the skewness of the distribution of relative price changes to capture such a potential interaction between the second and third scaled moments. We do not report estimation results for regression equations including this regressor because it turned out that such regressions would be beleaguered by severe collinearity problems.

sures. We therefore decided to reestimate regression equation (6) with the deflator of private consumption as dependend variable. The results of this exercise are given in Table 3. For the sake of brevity, only results for a specification with the moments of the distribution of weighted relative price changes as regressors are presented.

Table 3 — Inflation and the Distribution of Price Changes in Germany Based on the Deflator of Private Consumption (1969-2000)

Regressors: Weighted Measures of Higher-Order Moments				
	(1)	**(2)**	**(3)**	**(4)**
constant	0.6694	0.1083	0.4739	-0.0830
	(1.4641)	(0.1743)	(1.0777)	(-0.1409)
π_{t-1}	0.7935	0.7767	0.8196	0.8029
	(6.7488)***	(6.6487)***	(7.3619)***	(7.2881)***
csv_t		0.1567		0.1557
		(1.3137)		(1.3920)
css_t			0.2397	0.2391
			(2.1730)**	(2.2033)**
R^2	0.6110	0.6336	0.6671	0.6894
BG	2.8895	1.9337	2.6025	1.9229
JB	1.2295	2.5320	0.9810	0.7424

Notes: The figures in brackets under the coefficients are t-ratios which were calculated using robust standard errors. Robust standard errors were computed upon implementing the method of Newey and West (1987). *BG* denotes a Breusch-Godfrey test on autocorrelation of first order (F-value). *JB* denotes a Jarque-Berra test on normality of the residuals. *** (**, *) denotes rejection of the null hypothesis at the 1 (5, 10) percent level.

The results presented in Table 3 confirm that for the sample period under investigation the skewness of the distribution of relative price changes was an important determinant of the inflation process in Germany. Also note that the Jarque-Bera test statistic indicates that it is now possible not to reject the hypothesis of normally distributed residuals in the variant of the estimated regression equation including only a constant, the lagged aggregate inflation rate, and the skewness measure as regressors.

The significance of the skewness measure in the above equations may also be important for a better understanding of other prominent macroeconomic phenomena. For example, Ball and Mankiw (1995) stress that the skewness of the distribution of relative price changes as a measure of supply side shocks may help to explain the shifts of the Phillips curve often observed in the empirical literature.[5] To address this point, we estimated the following Phillips curve equation:

$$[7] \qquad \pi_t = \beta_0 + \beta_1 \pi_{t-1} + \beta_2 csv_t + \beta_3 css_t + \beta_4 u_t + \varepsilon_t \ .$$

where u_t represents the unemployment rate gap measured as the deviation of the actual German unemployment rate from its smoothed component extracted by implementing the filter technique developed by Hodrick and Prescott (1997). Applying the industrial standard, the smoothing parameter was set equal to $\lambda = 100$. The other symbols in equation (7) are as in equation (6).

The estimation results are depicted in Tables 4 and 5. Whereas the results reported in Table 4 are for equations with the aggregate inflation rate computed from the cross-section of relative price changes as dependent variable, the findings

Table 4 — Phillips Curve Regression Results for Germany (1969-2000)

Regressors: Weighted Measures of Higher-Order Moments				
	(1)	**(2)**	**(3)**	**(4)**
constant	1.2112	-1.1375	0.8817	-1.0518
	(3.0367)***	(-1.0334)	(2.2994)**	(-0.9697)
π_{t-1}	0.5087	0.3783	0.5674	0.4504
	(5.5517)***	(2.6582)**	(5.7322)***	(3.6368)***
csv_t		0.5989		0.5035
		(1.8640)*		(1.6601)
css_t			0.6830	0.5981
			(4.7190)***	(5.3110)***
u_t	-5.4093	-4.2582	-4.8301	-3.9345
	(-3.2412)***	(-2.6718)**	(-4.2232)***	(-3.7544)***
R^2	0.3645	0.5056	0.5380	0.6351
BG	1.1047	0.8177	2.1152	3.3344*
JB	18.9788***	0.6825	58.8362***	9.0954**

Notes: The figures in brackets under the coefficients are *t*-ratios which were calculated using robust standard errors. Robust standard errors were computed upon implementing the method of Newey and West (1987). *BG* denotes a Breusch-Godfrey test on autocorrelation of first order (*F*-value). *JB* denotes a Jarque-Berra test on normality of the residuals. *** (**, *) denotes rejection of the null hypothesis at the 1 (5, 10) percent level.

Table 5 — Phillips Curve Regression Results for Germany Based on the Deflator of Private Consumption (1969-2000)

[5] For recent empirical work documenting this stylized fact for German data, see, e.g., Franz (2001).

Regressors: Weighted Measures of Higher-Order Moments				
	(1)	(2)	(3)	(4)
constant	0.5636	0.1329	0.5387	0.1720
	(1.5257)	(0.2551)	(1.5084)	(0.3395)
π_{t-1}	0.7777	0.7668	0.7741	0.7650
	(8.2007)***	(8.0979)***	(8.4484)***	(8.3163)***
csv_t		0.1026		0.0877
		(1.1652)		(1.0197)
css_t			0.1591	0.1491
			(1.7341)*	(1.6163)
u_t	-3.4288	-3.2638	-3.2779	-3.1464
	(-4.0801)***	(-3.8534)***	(-4.0150)***	(-3.8093)***
R^2	0.7560	0.7677	0.7805	0.7889
BG	0.7454	0.9777	0.4982	0.7517
JB	0.7538	2.6418	0.1355	0.8878

Notes: The figures in brackets under the coefficients are *t*-ratios which were calculated using robust standard errors. Robust standard errors were computed upon implementing the method of Newey and West (1987). *BG* denotes a Breusch-Godfrey test on autocorrelation of first order (*F*-value). *JB* denotes a Jarque-Berra test on normality of the residuals. *** (**, *) denotes rejection of the null hypothesis at the 1 (5, 10) percent level.

documented in Table 5 refer to equations with the deflator of private consumption as the dependent variable. In all equations, the coefficient capturing the impact of the unemployment rate gap on the aggregate inflation rate is significantly different from zero and exhibits the expected negative sign.

The coefficients of the skewness variable css_t given in row four of Table 4 underpin the importance of the asymmetry of the distribution of the relative price changes for explaining shifts in the Phillips curve. Since the coefficients are positive it is possible to infer from the estimation results that a negative supply side shock, which results in a right-skewed distribution of the cross-section of relative prices, leads to an outward shift of the Phillips curve in the inflation-unemployment plane.

A similar result can be inferred from the estimation results reported in Table 5. As compared to the results depicted in Table 4 the residuals of the equations given Table 5 are normally distributed. The coefficient capturing the impact of the skewness variable css_t on the inflation process is significant at the 10 percent level in Column (3) and almost significant at a marginal significance level of 12 percent level in Column (4).

4.2 Skewness and Output Fluctuations

The theoretical arguments put forward by Ball and Mankiw (1995) and Balke and Wynne (2000) suggest that the skewness of the distribution of relative price changes can be used as a measure of supply side shocks. Such shocks have been found in the existing empirical literature to explain a substantial portion of the variation of real GDP at least in the long run.[6] This, in turn, suggests that the response of changes in real GDP to variations in this skewness of the distribution of relative price changes can be used to assess whether skewness measures supply side shocks properly. One way to study the dynamic interplay between skewness and fluctuations in real GDP is to compute impulse response functions obtained from a structural vector autoregression (SVAR).[7]

The framework to be estimated contains the skewness (css_t) of the distribution of relative price changes and changes in real GDP (Δy_t) as endogenous variables. The variable y_t denotes the natural logarithm of real GDP and Δ is the first–difference operator. Let the vector of endogenous variables be defined by $X_t \equiv (css_t \quad \Delta y_t)'$. Let the reduced form representation of this bivariate system be given as below:

$$[8] \qquad X_t = A_0 + \sum_{j=1}^{p} A_j X_{t-j} + e_t \ ,$$

where A_0 is a (2×1) vector of constants, A_i are (2×2) matrices of coefficients, and e_t represents a (2×1) disturbance vector. Using ordinary least squares, consistent and asymptotically efficient estimates of the coefficients of this bivariate system obtain. The lag length p of this system was determined by minimizing the Schwartz Bayesian Criterion. According to this criterion, one lag of the endogenous series was included in the VAR. A Portmanteau test indicated that using one lag of the vector of endogenous variables suffices to reject the hypothesis of remaining joint autocorrelation in the residuals of the VAR.

[6] For evidence for German data, see, for example, Funke (1997).
[7] For an application of SVAR models to the analysis of the interplay between average inflation and the skewness of the distribution of relative price changes, see Ratfai (2000).

As long as the roots of the characteristic equation of the bivariate system formalized in equation (8) can be found inside the unit circle, the unrestricted bivariate vector autoregression can be represented in its infinite vector moving average representation as:

$$[9] \qquad X_t = \overline{A}_0 + \sum_{j=0}^{\infty} A_1 L^j e_t \ ,$$

where we used the fact that $p = 1$. In equation (9), $\overline{A}_0$ denotes a (2×1) vector of coefficients and L symbolizes the lag operator.

Represent the moving average representation of the underlying structural model by:

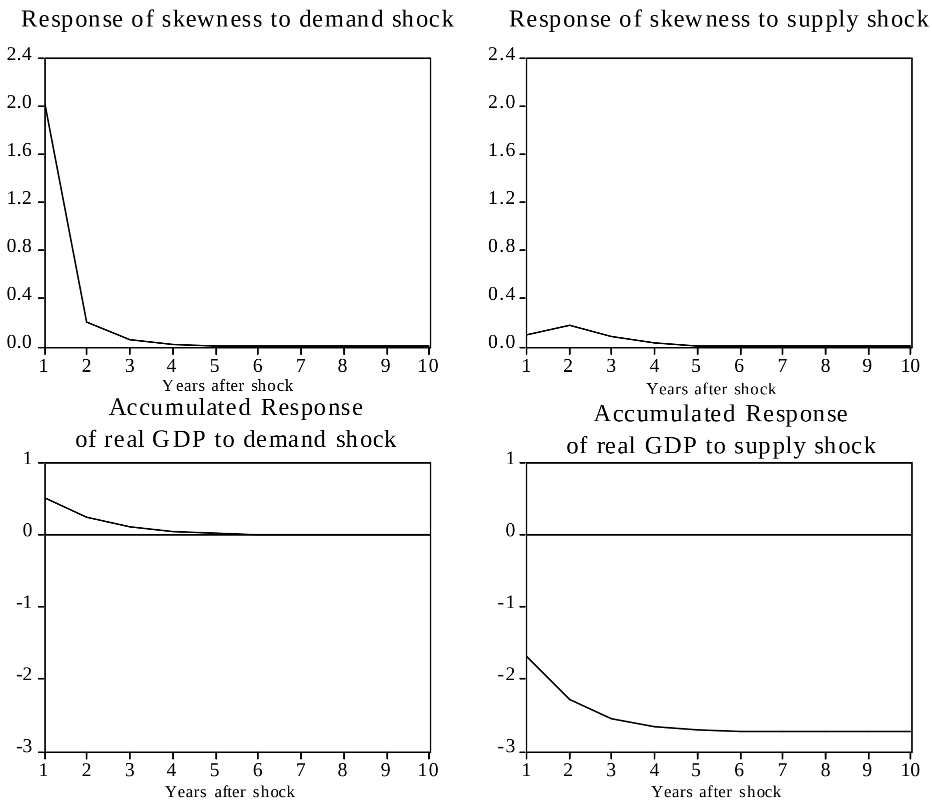
$$[10] \qquad X_t = \overline{A}_0 + \sum_{j=0}^{\infty} C_j L^j e_i \ ,$$

where the (2×2) matrices C_i represent the impulse response functions of the model and $\varepsilon_t \equiv (\varepsilon_{css,t} \quad \varepsilon_{\Delta y,t})$ is a vector of orthogonal serially uncorrelated structural shocks.

To recover these structural shocks from the sequence of the residuals e_t, first note that the relation between the vector autoregression and its vector moving average representation implies that $e_0 = C_0 \varepsilon_0$. Once the four elements of the matrix C_0 are identified all other matrices C_j, $j > 0$ can be computed and the response of the endogenous variables to the structural shocks can be traced out. The identification of the four elements of the matrix C_0 requires the imposition of a set of four restrictions on the system. The first three restrictions where obtained by normalizing the variance–covariance matrix of the underlying structural shocks to be given by an identity matrix. To derive the remaining restriction, we interpret the components of the vector ε_t of the structural shocks as structural supply ($\varepsilon_{css,t}$) and structural demand ($\varepsilon_{\Delta y,t}$) shocks, respectively. The impact of demand side shocks on real GDP are assumed to die out in the long-run (see also Blanchard and Quah 1989).

Resorting to this identification strategy allows the impulse response functions graphed in Figure 5 to be traced out. The responses of the endogenous variables to one-time, one standard deviation innovation in the real GDP equation are given in the first column of the figure. The dynamic implications of a one-time, one standard deviation innovation in the equation describing the skewness process are graphed in the second column of the figure. In the case of real GDP, we plot accumulated impulse response functions.

Figure 5 — Impulse Response Functions for the SVAR



A positive demand shock leads to a temporary increase in real GDP. Real GDP converges to its pre-shock value after approximately five years. The skewness of the distribution of relative price changes, in contrast, first increases in response to an expansionary demand shock and then decreases and to its long-run value. A one-time positive innovation in the skewness equation, that is, a negative supply side shock, results in a permanent pronounced decline in real GDP. As regards the impact of the supply side innovation on the skewness measure, the graph shows that the distribution of relative price changes becomes right-skewed in the aftermath of the shock. In the long run, the impact of this shock on the skewness measure dies out.

In a nutshell, the impulse response functions depicted in Figure 5 show a qualitative pattern typically found in business cycle analyses. We therefore conclude that the skewness of the distribution of relative price changes, which has been suggested by Ball and Mankiw (1995) as a measure of supply side shocks, is a useful tool to describe and to understand the sources of the German business cycle.

A further point to note is that an analysis of the relative magnitude of the responses of the endogenous variables to structural shocks indicates that resorting to the relatively low-dimensioned model outlined in equation (8) comes at a cost. In particular, the quantitative importance of demand side shocks for the dynamics of the skewness of relative price changes is not in line with the predictions of economic theory. Still, given the results of the above analyses, we are confident that the skewness of the distribution of relative price changes should also turn out to be a useful measure of supply side shocks in more complex VAR models. Constructing such models is left for future research.

V Conclusion

The paper presented evidence for German data regarding the interaction between aggregate inflation rate and the third scaled moment of the distribution of relative price changes. Motivated by the contributions of Ball and Mankiw (1995) and Balke and Wynne (2000), we tested whether the third scaled moment of the distribution of relative price changes influences the overall inflation rate significantly. We also estimated Phillips curve equations and a structural vectorautoregression to study whether the third scaled moment of the distribution of relative price changes can be used a a measure of supply side shocks.

Our main findings confirm the prediction of the theoretical literature that skewness should have explanatory power for the aggregate inflation process. These results turned out to be robust with respect to the specification of the quantitative model used in the empirical analysis. In particular, we found that the significant explanatory power of the skewness measure does not depend upon whether the aggregate inflation rate is measured in terms of the average of the cross-section of relative price changes or in terms of the deflator of private consumption.

Moreover, our results support the notion that shifts of the Phillips curve in Germany can be modeled by using the skewness of the distribution of relative price changes as a measure of supply side shocks. This result lends support to the argument that the instability of the Phillips curve relation in Germany can be attributed, at least to a certain extent, to supply side disturbances.

Finally, our results suggest that the skewness of the distribution of relative price changes may be a useful technical tool for analyzing and forecasting the evolution of the macroeconomic environment in Germany. In particular, the instrument may help to identify supply and demand side disturbances buffeting the economy. This is not only of interest for theoretical reasons but for practical purposes as well. For instance, policy makers and business cycle analysts, who are expected to arrive at decisions which depend on the nature of the shocks driving the stance of the business cycle, may find our results useful as well because they suggest that the skewness of the distribution of relative changes is a relatively accurate and timely measure of inflationary pressure.

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Appendix: Data Sources

The skewness of price changes was calculated on the basis of the producer price index for Germany. The sub-indices refer to Western Germany from 1969 to 1997, and to the unified Germany from 1998 onwards.

The definition of the sub-indices has changed during the sample. Time series derived from a uniform definition covering the complete sample period were not available. We therefore proceed as follows. We calculated the moments of price changes for the years 1969 to 1978 from 28 sub-indices of the "Index der Erzeugerpreise industrieller Produkte (1970=100) : insg. , nach Investitions-, Verbrauchsgütern, Hauptgruppen, ausgewählten Warengruppen, -Klassen und-Arten des WI, Ausgabe 1970 (früheres Bundesgebiet)". This is the „Segment 233" of the data base provided by the Federal Statistical Office Germany. For the time period 1979 to 1998, we used 32 sub-indices of the "Index der Erzeugerpreise gewerblicher Produkte (1991=100) : insg. , nach Gütergruppen, -Zweigen, -Klassen und -Arten des GP, Ausgabe 1989 (früheres Bundesgebiet)". This is the „Segment 3364" of the data base mentioned above. For 1999 and 2000, we used 27 sub-indices of the "Index der Erzeugerpreise gewerblicher Produkte(1995=100) : insgesamt, nach Gütergruppen, -Zweigen, -Klassen und-Arten des GP, Ausgabe 1995". This is the „Segment 3783" of the same data base.

The weights used for computing the weighted moments of the distribution of relative price changes were taken from the Statistical Yearbook for Germany. In particular, the weights for 1969 to 1978 were taken from the yearbook for 1976, the weights for the time period from 1979 to 1998 were taken from the yearbook for 1991, and the weights for 1999 and 2000 were taken from the yearbook for 2000.

As regards the deflator of private consumption, we used the time series for the national accounts for Western Germany up to 1994. From 1995 onwards, we employed EVSG-data for the unified Germany.

The unemployment rate used in the Phillips curve regressions is the standardized unemployment rate for Western Germany provided by the OECD (Main Economic Indicators).

Real gross domestic product (GDP) for Western Germany for the period 1968-1991 was taken from the CD-Rom "50 Jahre Deutsche Mark" edited by the Deutsche Bundesbank. From 1992 onwards, the data refer to the unified Germany and are taken from the publication "Fachserie 18, Volkswirtschaftliche Gesamtrechnung, Hauptbericht" edited by the Federal Statistical Office Germany.