Inflation and Relative Price Variability in a Low Inflation Country: Empirical Evidence for Germany*

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Abstract

The recent literature on the welfare cost of inflation emphasizes inflation’s effect on the variability of relative prices. Expected and unexpected inflation have both been proposed to increase relative price variability (RPV) and, thereby, to distort the information content of nominal prices. This paper presents new evidence on the impact of inflation on RPV in Germany. Our results indicate that the influence of expected inflation disappears if a credible monetary policy stabilizes inflationary expectations on a low level. Yet the significant impact of unexpected inflation suggests that even low inflation rates can lead to welfare losses by raising RPV above its efficient level.

Keywords: Inflation, Relative Price Variability, Welfare Cost of Inflation.

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1 Introduction

Given the general consensus about the crucial importance of price stability, it is striking that the empirical evidence on the cost of inflation is rather weak. Lucas (2000), for example, estimated the gain from reducing annual inflation from 10 percent to zero to be equivalent to an increase in real income of less than one percent for the United States. In the spirit of Friedman (1969), Lucas concentrated on the negative impact of inflation on the aggregate money demand, i.e. on the *quantity* of money. This approach tends to underestimate the cost of inflation assuming inflation not to reduce the *quality* of money. Recent emphasis has therefore been placed on inflation’s effect on the distribution of relative prices in the economy. From this perspective, inflation is costly since it increases relative price variability (RPV) and, thereby, distorts the information content of nominal prices.

Advancing on the seminal study of Parks (1978), the bulk of the literature has restricted the attention to the U.S. experience, see Aarstol (1999), Jaramillo (1999), Chang and Cheng (2000), Parsley and Popper (2002), or Banerjee et al. (2002) for recent contributions. However, as Bomberger and Makinen (1993) already emphasized, the link between U.S. inflation and RPV might be solely due to a few episodes of high inflation. The present paper therefore re-investigates the relation between inflation and RPV for Germany, the textbook example for a low inflation country.¹

Theories explaining the welfare cost of inflation through its impact on RPV are typically based on menu cost and signal extraction models. Both approaches predict that inflation increases RPV and, thus, impedes the efficient allocation of resources. While menu cost models emphasize the effects of expected inflation, the focus of signal extraction models is on the cost of unexpected inflation. The empirical analysis of the

¹ For European countries, the available empirical evidence for the inflation-RPV link is very limited. Notable exceptions are the cross-country studies of Fielding and Mizen (2000) and Silver and Ioannidis (2001). In these papers, however, the period under investigation ends already in 1993 and 1989, respectively. Ball and Mankiw (1995) and Balke and Wynne (2000) demonstrated that inflation can also increase the skewness of the distribution of relative prices.
relation between inflation and RPV therefore requires a model of inflation that generates measures of expected and unexpected inflation. To that aim, we will test the usefulness of various macroeconomic variables for forecasting German inflation. After these preliminaries, the analysis proceeds by estimating the impact of expected and unexpected inflation on RPV.

Assuming that inflation is the exogenous variable, the empirical literature on the influence of inflation on RPV has typically used simple least squares regressions. However, if RPV and inflation are affected by the same shocks on individual prices, the exogeneity assumption becomes questionable and least squares estimates are potentially biased. The role of supply shocks for the inflation-RPV link has thus been an important issue in the empirical literature. Earlier contributions tried to ameliorate this problem by excluding energy prices from the RPV measure, see e.g. Jaramillo (1999). In the current paper, we will also employ various instrumental variable estimators to account for the endogeneity of inflation and to check the robustness of our results.

The empirical results clearly indicate that — in contrast to the evidence found for the United States — expected inflation has no impact on RPV in Germany. Similar to the United States, however, there is a significant influence of unexpected inflation on RPV. Accordingly, signal extraction models seem to give a more convincing explanation for the observed inflation-RPV link in Germany than menu cost models. Overall, the evidence found for Germany strongly suggests that even low inflation leads to welfare losses by raising the variability of relative prices above its efficient level.

The remainder of the paper is organized as follows. Section 2 briefly recalls the intuition of the theoretical literature on inflation and relative price variability. Section 3 introduces the RPV measure and discusses some measurement issues. Section 4 presents the inflation forecasts for unified Germany and in Section 5 we specify the empirical model that relates RPV to expected and unexpected inflation. Section 6 focuses on the role of energy prices. Some concluding remarks are offered in Section 7.
2 Theories linking inflation and RPV

The prevalent theories generating welfare cost of inflation via its impact on RPV are (1) the menu-cost models of e.g. Sheshinski and Weiss (1977) or Rotemberg (1983) and (2) the signal-extraction models introduced by Lucas (1973), Barro (1976) and Hercowitz (1981). In both type of models, inflation increases RPV and, thus, decreases the informativeness of prices resulting in a decrease of welfare due to the misallocation of resources.

The main idea of menu cost models is that changing prices is costly. These models predict that firms respond to inflation using a so-called \((S, s)\) pricing rule: a firm holds the nominal price of its goods constant until inflation reduces the real price to the lower bound \(s\). There the price is reset such that the resulting real price equals the upper bound \(S\). As inflation increases, the difference between the optimal \(s\) and \(S\) widens. More importantly, if there are differences in fixed costs of price changes across firms or firm-specific shocks, staggered price setting will be generated and higher inflation will amplify the dispersion of relative prices. Therefore, menu cost models typically suggest a positive relationship between RPV and expected inflation.

Of course, inflation is not always anticipated correctly. The Lucas-Barro signal-extraction models derive a positive link between inflation uncertainty and RPV. The intuition behind these models can be summarized as follows. Since higher inflation uncertainty makes aggregate demand shocks more unpredictable, it becomes optimal for firms to adjust output less in response to all shocks, including idiosyncratic real demand shocks. As a consequence of this misperception, prices have to move more in each market to equate quantity demanded with the less variable quantity supplied. Since prices will be dispersed more widely fewer firms adequately respond to demand shocks with output adjustments. Hence, increased inflation uncertainty will raise RPV. In the original Lucas-Barro model, realized aggregate demand shocks have no impact on RPV because all firms react identically to any given aggregate shock. In fact, if price elasticities of
supply differ across firms, then RPV will also respond to the magnitude of unexpected inflation, see Hercowitz (1981).

Recently, Balke and Wynne (2000) showed in a general multi-sector equilibrium model that inflation can increase RPV due to correlated technology shocks. In contrast to menu cost and signal extraction models this alternative approach neither assumes price rigidities nor imperfect information. Therefore, technology shocks could provide a plausible explanation for the inflation-variability link in equity prices, see Parsley and Popper (2002).

Emphasizing the different effects of expected and unexpected inflation, these theories deliver important implications for the empirical analysis. Note, however, that they predict an increase of RPV irrespective of the sign of expected or unexpected inflation. Yet for U.S. data the evidence clearly suggests that the response of RPV to inflation is not symmetric, see Aarstol (1999), Jaramillo (1999) and Parsley and Popper (2002). In the following, we will test whether this asymmetry in the link between RPV and inflation is also prevailing in Germany.

3 The RPV measure

We compute a monthly measure of relative price variability using the categories of the German consumer price index (CPI). In Sections 4 and 5 we will focus on unified Germany (1991.01- 2003.12), where the Federal Statistical Office provides consistent data for up to 32 subcategories of the CPI. A CPI with 9 subcategories is employed in Section 6 where we check the robustness of our results for West German data (1962.01-1990.12).  

Since the influential paper of Parks (1978), the variability of relative price changes in

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2 Data for West-German CPI subcategories is not available in electronically readable form. More detailed information on the various subcategories is given in Appendix A.
period $t$ ($RPV_t$) is defined as

$$RPV_t = \sqrt{\sum_{i=1}^{k} w_i(\pi_{it} - \pi_t)^2}$$  \hspace{1cm} (1)

where $\pi_{it} = \ln(P_{it}/P_{it-1})$ and $P_{it}$ is the price index of the $i$th subcategory in period $t$. $w_i$ denotes the weight of the $i$th subcategory in the aggregate index so that $P_t = \sum_{i=1}^{k} w_i P_{it}$ gives the aggregate price level and the inflation rate $\pi_t$ is $\ln(P_t/P_{t-1})$.

Alternatively, one could use the square of $RPV$ as a measure for relative price variability but our results will not depend on this choice. Silver and Ioannidis (2001) propose the use of the coefficient of variation, with the $RPV$ measure introduced in (1) divided by the level of inflation to measure relative price variability. However, there are periods where German inflation is zero or even negative so that the coefficient of variation is simply not defined.

Using the appropriate weights in the definition of RPV is of crucial importance for the following analysis. In a typical price index, the weights of various subcategories differ drastically. While, for example, in the 32 subcategory index the weight of food is 0.09047, the prices of household textiles (0.00479) are by far less important for the economy. Neglecting these differences, as in Aarstol (1999) or Grier and Perry (1996), produces an unreliable measure of RPV.\footnote{Fielding and Mizen (2000) define RPV across several European countries as the unweighted average of national magnitudes implying that e.g. Luxembourg and Denmark have the same impact on the measure of relative price dispersion in Europe as Germany, France or the UK.}

## 4 The inflation forecast

The theories on the relation between inflation and RPV distinguish between expected and unexpected inflation. In line with the empirical literature we generate the measures for expected and unexpected inflation using a forecast equation of inflation. It is a general problem of any such decompositions that the empirical results might depend
on the accuracy of the measures of expected and unexpected inflation. In contrast to previous studies, we do not restrict our attention to models in which the inflation forecast is exclusively based on lagged inflation rates. We rather included additional macroeconomic variables into the inflation equation to improve the inflation forecast.

In light of the findings of Stock and Watson (1999) for the United States, we tested the explanatory power of the German unemployment rate ($U$), industrial production ($y$), sales in the manufacturing sector ($ms$), the DM/Dollar exchange rate ($e$), the monetary aggregates M1 and M3, and of a representative short- ($i_S$) and long-term interest rate ($i_L$) for predicting inflation. All variables are in logs and stationarity requires that they enter the equations in first differences.\footnote{For brevity, results of unit root tests are not presented. Notice further that the levels of the relevant variables are not cointegrated. The data is seasonally adjusted using the Census X12-method. See Appendix B for a detailed data description.}
Table 1: Forecasting inflation in unified Germany

$$\pi_t = \sum_{i=1}^{12} \alpha_i \pi_{t-i} + \sum_{i=1}^{12} \beta_i x_{t-i} + \varepsilon_t$$

<table>
<thead>
<tr>
<th>x</th>
<th>$\Delta u$</th>
<th>$\Delta m1$</th>
<th>$\Delta m3$</th>
<th>$\Delta i_S$</th>
<th>$\Delta i_L$</th>
<th>$\Delta e$</th>
<th>$\Delta y$</th>
<th>$\Delta ms$</th>
</tr>
</thead>
<tbody>
<tr>
<td>F-statistic ($H_0: \forall i \beta_i = 0$)</td>
<td>1.18</td>
<td>0.93</td>
<td>0.73</td>
<td>0.27</td>
<td>0.49</td>
<td>0.72</td>
<td>0.78</td>
<td>1.97</td>
</tr>
<tr>
<td>Q(12)</td>
<td>0.68</td>
<td>3.61</td>
<td>1.69</td>
<td>0.98</td>
<td>0.85</td>
<td>1.25</td>
<td>1.21</td>
<td>2.43</td>
</tr>
<tr>
<td>ARCH(12)</td>
<td>0.48</td>
<td>0.90</td>
<td>1.13</td>
<td>0.89</td>
<td>0.97</td>
<td>1.24</td>
<td>1.07</td>
<td>0.82</td>
</tr>
</tbody>
</table>

Notes: \(x\) does not improve the univariate inflation forecast whenever the F-statistic is not significant. \(Q(12)\) denotes the Ljung-Box statistic testing for serial correlation in the residuals. \(ARCH(12)\) indicates the LM-statistic for ARCH effects. p-values are given in brackets. Sample: 1992.02-2003.12.

Figure 1 shows the development of monthly inflation in Germany from February 1991 to December 2003. As a consequence of the Bundesbank’s strict anti-inflation policy in the aftermath of German unification, inflation rates in the 1990s have been rather low and stable. Since the start of the European Monetary Union, the ECB was equally successful in keeping inflation low in Germany. The three largest observations in monthly inflation can be attributed to increases in the rentals in East-Germany (1991.10 and 1993.01) and due to increases of the petroleum tax (1991.07) and the value-added tax (1993.01). It is important to note that these price increases were all preannounced and, thus, fully anticipated by rational economic agents. Therefore, the inflation rate is regressed on the impulse dummy variables \(d91.07, d91.10\) and \(d93.01\) to avoid the distorting effect of those expected price changes on the inflation forecast.

The estimation results of predicting inflation are displayed in Table 1. We start by testing the forecasting ability of each variable (\(x\)) separately. Specifically, we augmented
an AR(12) model of inflation by lags of $x$ and tested for their significance. A variable cannot predict inflation if the corresponding F-statistic indicates that it may be omitted from the inflation equation. Table 1 shows that according to this criterion only the growth rates of sales in the manufacturing sector ($\Delta ms$) improve the inflation forecast. In the following, expected and unexpected inflation are therefore generated by the inflation forecast equation including lags of inflation and $\Delta ms$.

According to Table 1 there is no evidence for ARCH effects in German inflation, i.e. its conditional variance is constant over time. As a consequence, the inflation forecast equation cannot be used to generate a time-varying measure of ex ante inflation uncertainty, see Engle (1983). In this framework, Grier and Perry (1996) demonstrated that inflation uncertainty drives relative price variability in the United States. The negligible impact of inflation uncertainty in Germany might be a consequence of the high reputation of the Bundesbank and its low inflation record.

5 The empirical link between inflation and RPV

Figure 2 shows the RPV measure based on 32 categories of the German seasonally adjusted CPI covering the sample period from 1991.02 to 2003.12. Apparently, the two largest values of RPV are caused by the anticipated price increases in 1991, see Section 4. However, preannounced relative price changes that are known to be solely induced from the supply side increase RPV but are likely to be less distorting for the information content of prices. For that reason, the impulse dummy variables ($d_{91.07}, d_{91.10}, d_{93.01}$) defined above account for the impact of this known relative price changes.

Following Parks (1978), our analysis starts with regressing RPV on the absolute value

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5 The Ljung-Box Q-statistic shown in Table 1 indicates that the seasonal lag order sufficiently accounts for the serial correlation in the inflation model.
Figure 2: Relative price variability in Germany

Notes: Relative price variability (percentage points) is defined in (1) and is based on 32 subcategories of the seasonally adjusted Consumer Price Index, 1991.02–2003.12.

of contemporaneous inflation:\(^6\)

\[ RPV_t = \alpha_0 + \alpha_1 |\pi_t| + \varepsilon_t \]  \hspace{1cm} (2)

As the first row of Table 2 shows, the coefficient of inflation estimated with ordinary least squares (OLS) is plausibly signed and highly significant. In line with the results typically found for U.S. data, this result strongly indicates a positive link between the aggregate inflation and RPV.

Estimating Equation (2) with ordinary least squares assumes that RPV is the endogenous variable and that the aggregate inflation rate is exogenous. However, as a referee pointed out, if both variables are jointly determined by the same shocks in individual prices, there is correlation between inflation and the error term of Equation (2) leading

\(^6\) Using the absolute value instead of actual inflation is in line with the implications of the theoretical models, see Section 2. In the same vein, Parks (1978) regressed \(R PV^2\) on \(\pi^2\).
Table 2: The inflation-RPV link for unified Germany

\[ R_{PV_t} = \alpha_0 + \alpha_1|\pi_t| + \varepsilon_t \]

<table>
<thead>
<tr>
<th>Method</th>
<th>( \alpha_1 )</th>
<th>( \bar{R}^2 )</th>
<th>( ARCH(6) )</th>
<th>Over-identifying restrictions test</th>
</tr>
</thead>
<tbody>
<tr>
<td>OLS</td>
<td>0.80**</td>
<td>0.31</td>
<td>0.80 [0.57]</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(2.97)</td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>TSLS</td>
<td>1.86**</td>
<td>0.23</td>
<td>1.12 [0.35]</td>
<td>( \chi^2(7) = 9.83 )</td>
</tr>
<tr>
<td></td>
<td>(3.02)</td>
<td></td>
<td></td>
<td>( [0.20] )</td>
</tr>
<tr>
<td>GMM</td>
<td>2.10**</td>
<td>0.18</td>
<td></td>
<td>( \chi^2(7) = 8.98 )</td>
</tr>
<tr>
<td></td>
<td>(2.99)</td>
<td></td>
<td></td>
<td>( [0.25] )</td>
</tr>
</tbody>
</table>

Notes: *, ** indicate significance at the 5% and 1% level. t-statistics are given in parentheses, p-values in brackets. The regression includes lagged endogenous variables to capture the serial correlation pattern. \( ARCH(6) \) indicates the LM-statistic for ARCH effects. The \( \chi^2 \)-statistic tests for the validity of the following instruments (lags in parentheses): \( \Delta m1(0,2) \), \( \Delta m3(1,3) \), \( \Delta m1(1,3) \), \( \Delta u(0) \), \( \Delta e(4) \). Sample: 1992.02-2003.12.

to biased parameter estimates. So far, the empirical literature tried to ameliorate this problem by eliminating energy prices from the RPV measure since those prices are particularly prone to supply side shocks, see e.g. Jaramillo (1999) and Section 6. In the following, we suggest an alternative solution by employing instrumental variable estimators which should account for the possible endogeneity of inflation. Our first choice is the GMM estimator which has excellent asymptotic properties but may perform poorly in small samples, see e.g. Hansen, Heaton and Yaron (1996). Therefore, we additionally applied two stage least squares (TSLS) estimators which typically have more convincing small sample properties.

Valid instruments must be uncorrelated with the residuals of the RPV equation (2) and correlated with the endogenous explanatory variable (i.e. inflation). Instruments that are only weakly correlated with the instrumented variable can produce biased estimators and test statistics. Therefore, we identified the strongest instruments for each explanatory variable applying the model selection algorithm PcGets developed
by Hendry and Krolzig (1999). PcGets starts with a general model containing contemporaneous and lagged values of all potential instruments and removes redundant instruments successively. Re-estimating Equation (2) with GMM and TSLS yields the results presented in row 2 and 3 of Table 2. The set of potential instruments for inflation contained all variables employed in the inflation forecast exercise (unemployment rate, industrial production, manufacturing sales, DM(Euro)/Dollar exchange rate, monetary aggregates M1 and M3, short- and long term interest rates), compare Section 4. Note that the test for over-identifying restrictions (see e.g. Davidson and MacKinnon (1993)) indicates for both estimators the validity of the underlying instruments. The GMM and TSLS estimators confirm the significant influence of inflation on RPV suggested by the OLS estimation. Yet the downward bias of the OLS estimate is substantial.

Equation (2) is misspecified if there is a different impact of expected and unexpected inflation on RPV. In this case, the following regression is more appropriate:

\[
RPV_t = \beta_0 + \beta_1 EI_t + \beta_2 UI_t + \epsilon_t
\]

where \( EI \) and \( UI \) denote expected and unexpected inflation, respectively, generated by the inflation model selected in the previous section. Similar to the results obtained by Aarstol (1999) for the United States, Table (3) shows that – irrespective of the estimation strategy – unexpected inflation is much more important for RPV in Germany than expected inflation. In particular, testing for the equality of the coefficients via a standard F-Test yields rejection at the 5% level in all cases. In contrast to the U.S. evidence, expected inflation has no significant influence on RPV in Germany.\(^8\)

\(^7\) PcGets searches all possible paths of the testing-down process and reports the most parsimonious model that does not violate a reduction test. For a further application of PcGets, see e.g. Hayo and Hofmann (2003).

\(^8\) Equation (3) involves generated regressors such that the validity of standard t-statistics is not obvious. Pagan (1984) has shown that OLS estimation is consistent and does not necessarily lead to efficiency losses if generated forecasts \((EI)\) as well as forecast errors \((UI)\) enter the equation. The only problem concerns the OLS-generated t-statistic of the coefficient of \(EI\) which tends to be overstated. Yet this is not a big problem, since the acceptance of the relevant null hypothesis (no
Table 3: The impact of expected and unexpected inflation on RPV

\[ RPV_t = \beta_0 + \beta_1 |EI_t| + \beta_2 |UI_t| + \varepsilon_t \]

<table>
<thead>
<tr>
<th></th>
<th>( \beta_1 )</th>
<th>( \beta_2 )</th>
<th>( F(\beta_1 = \beta_2) )</th>
<th>( R^2 )</th>
<th>ARCH(6)</th>
<th>Over-identifying restrictions test</th>
</tr>
</thead>
<tbody>
<tr>
<td>OLS</td>
<td>-0.63</td>
<td>1.38**</td>
<td>12.65</td>
<td>0.32</td>
<td>0.68</td>
<td>( \chi^2(4) = 2.19 )</td>
</tr>
<tr>
<td></td>
<td>(-1.59)</td>
<td>(3.19)</td>
<td>[0.00]</td>
<td></td>
<td>[0.67]</td>
<td></td>
</tr>
<tr>
<td>TSLS</td>
<td>-0.60</td>
<td>1.79</td>
<td>5.27</td>
<td>0.32</td>
<td>0.73</td>
<td>( \chi^2(4) = 2.49 )</td>
</tr>
<tr>
<td></td>
<td>(-1.52)</td>
<td>(1.74)</td>
<td>[0.02]</td>
<td></td>
<td>[0.63]</td>
<td></td>
</tr>
<tr>
<td>GMM</td>
<td>-0.48</td>
<td>1.55*</td>
<td>4.86</td>
<td>0.32</td>
<td>0.65</td>
<td></td>
</tr>
<tr>
<td></td>
<td>(-1.41)</td>
<td>(2.01)</td>
<td>[0.03]</td>
<td></td>
<td></td>
<td></td>
</tr>
</tbody>
</table>

Notes: *,** indicate significance at the 5% and 1% level. t-statistics are given in parentheses, p-values in brackets. \( F(\beta_1 = \beta_2) \) indicates the F-statistic testing \( H_0: \beta_1 = \beta_2 \). The \( \chi^2 \)-statistic tests for the validity of the following instruments for \( |UI| \) (lags in parentheses): \( \Delta m1(1,4), \Delta m3(3), \Delta y(1), \Delta u(0) \). For more information, see Tab. 2. Sample: 1992.02-2003.12.

In the United States, RPV has a more pronounced response to inflation when inflation is unexpectedly high, i.e. when \( UI > 0 \), see Aarstol (1999). In order to check whether this result can be confirmed for Germany, we run the following RPV regression allowing for an asymmetric response of RPV to unexpected inflation:

\[ RPV_t = \gamma_0 + \gamma_1 |EI_t| + \gamma_2 UIP_t + \gamma_3 UIN_t + \varepsilon_t \]

where \( UIP = UI \) if \( UI \geq 0 \) and \( UIN = |UI| \) otherwise.

Table (4) confirms that the link between inflation and RPV tends to be particularly strong in Germany when unexpected inflation is positive. For both instrumental variable estimators, the hypothesis that the coefficients of positive and negative unexpected inflation are equal can be rejected at the 10% level.

Overall, our results indicate that the relationship between inflation and RPV in Germany is due to unexpected rather than expected inflation. Apparently, menu costs play
Table 4: The inflation-RPV link allowing for asymmetry

\[RPVi = \gamma_0 + \gamma_1|EIi| + \gamma_2UIPi + \gamma_3UINi + \varepsilon_i\]

<table>
<thead>
<tr>
<th></th>
<th>(\gamma_1)</th>
<th>(\gamma_2)</th>
<th>(\gamma_3)</th>
<th>(F(\gamma_2 = \gamma_3))</th>
<th>(R^2)</th>
<th>(ARCH(6))</th>
<th>Over-identifying restrictions test</th>
</tr>
</thead>
<tbody>
<tr>
<td>OLS</td>
<td>-0.64</td>
<td>1.63**</td>
<td>0.98</td>
<td>1.67 [0.20]</td>
<td>0.33</td>
<td>0.58</td>
<td>(\chi^2(13) = 18.97) [0.12]</td>
</tr>
<tr>
<td></td>
<td>(-1.63)</td>
<td>(3.44)</td>
<td>(1.83)</td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>TSLS</td>
<td>-0.67</td>
<td>3.44**</td>
<td>1.18</td>
<td>2.73 [0.10]</td>
<td>0.20</td>
<td>1.51</td>
<td>(\chi^2(13) = 19.49) [0.11]</td>
</tr>
<tr>
<td></td>
<td>(-1.54)</td>
<td>(3.41)</td>
<td>(0.92)</td>
<td></td>
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</tr>
<tr>
<td>GMM</td>
<td>-0.51</td>
<td>3.11**</td>
<td>1.02</td>
<td>3.16 [0.08]</td>
<td>0.19</td>
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<td></td>
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<tr>
<td></td>
<td>(-1.56)</td>
<td>(3.24)</td>
<td>(0.97)</td>
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</tbody>
</table>

Notes: * ** indicate significance at the 5% and 1% level. t-statistics are given in parentheses, p-values in brackets. \(F(\gamma_2 = \gamma_3)\) indicates the F-statistic testing \(H_0: \gamma_2 = \gamma_3\). The \(\chi^2\)-statistic tests for the validity of the following instruments for \(UIP\) and \(UIN\) (lags in parentheses): \(\Delta m1(0,1,4), \Delta m3(0,1), \Delta u(0,1,2,4), \Delta e(0,4), \Delta ms(3), \Delta i_L(1,2)\). For more information, see Tab. 2.


no significant role in Germany, where inflation has been low for a long time. There is more evidence in favor of the signal extraction approach for explaining the cost of inflation. Similar to the U.S. evidence, the effect of inflation on the variability of relative prices seems to be stronger in Germany if inflation has been unexpectedly high.

6 The role of supply shocks: results for West Germany

In contrast to anticipated price changes, like the preannounced tax increases in Germany discussed above, the role of supply shocks for the inflation-RPV link is still under debate. Corroborating the arguments of Fischer (1981) and Taylor (1981), Bomberger and Makinen (1993) demonstrated that the relation between inflation and RPV established by Parks (1978) for the United States vanishes if the oil shock years are omitted from the sample, or if energy prices are excluded from the RPV measure.\(^9\)

\(^9\) Jaramillo (1999) reinstated Parks’ results by allowing for an asymmetrical response of relative prices to episodes of positive and negative inflation.
In this section, we examine the effects of supply side shocks on the German inflation-RPV relation. Compared with the crucial importance of the oil price shocks in the 1970s, supply shocks played no dominant role for the German economy in the 1990s. The influence of supply shocks on the inflation-RPV link is therefore investigated using West German data from 1962 to 1990.

Following Bomberger and Makinen (1993) we control for the effect of oil price shocks by computing an adjusted measure of RPV that excludes the categories dominated by energy prices. For the West German CPI based on 9 categories, these categories are energy and transport and communication. RPV based on the complete price index and the energy-adjusted RPV are shown in Figure 3. As for U.S. data, RPV peaked in Germany as a result of the two oil price shocks in 1974 and 1980. Apparently, both outliers are well captured by the energy-adjusted RPV measure.
Table 5: Supply shocks and the inflation-RPV link for West Germany

\[
RPV_t = \alpha_0 + \alpha_1|\pi_t| + \varepsilon_t \quad \quad RPV_t = \gamma_0 + \gamma_1|EI_t| + \gamma_2UIP_t + \gamma_3UIN_t + \varepsilon_t
\]

<table>
<thead>
<tr>
<th>RPV measure</th>
<th>(\alpha_1)</th>
<th>(\gamma_1)</th>
<th>(\gamma_2)</th>
<th>(\gamma_3)</th>
<th>(F(\gamma_2 = \gamma_3))</th>
</tr>
</thead>
<tbody>
<tr>
<td>complete CPI</td>
<td>0.34**</td>
<td>0.12</td>
<td>0.88**</td>
<td>0.53**</td>
<td>4.80</td>
</tr>
<tr>
<td></td>
<td>(3.60)</td>
<td>(1.25)</td>
<td>(4.62)</td>
<td>(4.44)</td>
<td>[0.03]</td>
</tr>
<tr>
<td>energy-adjusted CPI</td>
<td>0.17**</td>
<td>0.09</td>
<td>0.45**</td>
<td>0.37**</td>
<td>0.83</td>
</tr>
<tr>
<td></td>
<td>(3.69)</td>
<td>(1.66)</td>
<td>(5.57)</td>
<td>(5.32)</td>
<td>[0.35]</td>
</tr>
</tbody>
</table>

Notes: Each regression contains seasonal dummies and lagged endogenous variables to capture the serial correlation pattern. White t-statistics are given in parentheses, p-values in brackets. \(F(\gamma_2 = \gamma_3)\) indicates the F-statistic testing \(H_0: \gamma_2 = \gamma_3\). Sample: 1962.02-1990.12.

Table 5 presents the results for the inflation-RPV link for both RPV measures. In the first column of Table 5 we adopted Parks’ (1978) approach by estimating the impact of the absolute value of inflation on RPV. The estimated coefficient of inflation is significant for both RPV measures but the exclusion of the energy-sectors seems to dampen the effects of inflation on RPV.

In the second column of Table 5, we employ the specification introduced in Equation (4) allowing RPV to depend on expected and unexpected inflation.\(^{10}\) For both RPV measures, the absence of an effect of expected inflation is confirmed for Germany. In line with the evidence found for unified Germany, West German RPV reacts more strongly to \(UIP\) than to \(UIN\). There are only two notable differences: (i) the coefficient of \(UIN\) is significant in the West German data set and (ii) the difference between the impact of \(UIP\) and \(UIN\) is not significant for the energy-adjusted RPV measure. Generally, the inflation-RPV link is weaker if the energy sectors are excluded from the RPV measure. Yet our results do not indicate that the relation between inflation and

\(^{10}\) To derive the inflation forecast equation for West Germany we applied the same procedure and tested the same set of macroeconomic variables as in Section 4. For West Germany, the specified forecast model is essentially univariate and the results do not depend on additional macroeconomic variables. For brevity, the inflation forecast equation for West Germany is not presented but results are available on request.
RPV is exclusively due to supply shocks.

7 Concluding remarks

While there is no doubt that an economy will suffer with inflation is persistently high, economists have often difficulty giving a convincing account for the cost of low or moderate inflation rates. Attempts to estimate the welfare gains from reducing inflation via its influence on aggregate money demand tend to fail since they assume that inflation leaves the functions of money unaffected.\[^{11}\] This paper deals with the distorting effect of inflation on the information content of nominal prices. Expected inflation, realized unexpected inflation, and inflation uncertainty have all been proposed to increase relative price variability (RPV) and to impede the efficient allocation of resources.

So far the literature focused on the U.S. experience. However, it is often argued that the observed link between U.S. inflation and RPV might be solely due to supply shocks and a few episodes of high inflation. The present paper therefore investigates the impact of inflation on RPV in Germany, the textbook example for a low inflation country. It is worth noting that the evidence found for Germany can also be viewed as exemplary for the European Monetary Union (EMU) because the European Central Bank apparently mimics the strong anti-inflation attitude of the Bundesbank.

The main results of the paper can be summarized as follows. In accordance with the evidence found by Aarstol (1999) for the United States, unexpected inflation significantly increases RPV in Germany. This is also confirmed for West German data where we paid particular attention to the influence of oil price shocks. However, in contrast to the United States there is no effect of expected inflation on German RPV.

\[^{11}\] Recently, Simonsen and Cysne (2001) extended the traditional ”money demand” approach to measure the welfare cost of inflation by including interest-bearing deposits into the analysis. Along these lines, they show that inflation leads to excess financial intermediation. Another strand of the literature explores the relationship between inflation and inflation uncertainty. Following Friedman (1977), higher inflation invokes the inflation variability incurring higher welfare cost. However, in a recent study Hwang (2001) supports Engle’s (1983) famous result that a high rate of inflation does not necessarily increase inflation uncertainty.
Our empirical findings suggest that the influence of expected or average inflation disappears if a credible monetary policy stabilizes inflationary expectations on a low level. Yet the significant impact of unexpected inflation on German RPV indicates that even low inflation can reduce the information content of nominal prices. Advancing on previous work of e.g. Blejer and Leiderman (1980) or Neumann and von Hagen (1991), future research on the welfare implications of inflation should therefore account for the consequences of inflation-induced increases of relative price variability.

References


Appendix

A The categories of the German Consumer Price Index

Weights of categories are presented in parentheses. Data source for CPI data is the Federal Statistical Office Germany.

The 32 categories of the CPI for unified Germany

<table>
<thead>
<tr>
<th>Category</th>
<th>Weight</th>
</tr>
</thead>
<tbody>
<tr>
<td>Food (90.47)</td>
<td></td>
</tr>
<tr>
<td>Non-alcoholic beverages (12.88)</td>
<td></td>
</tr>
<tr>
<td>Newspaper, books and stationery (19.08)</td>
<td></td>
</tr>
<tr>
<td>Glassware, tableware and household utensils (4.91)</td>
<td></td>
</tr>
<tr>
<td>Tools and equipment for house and garden (5.73)</td>
<td></td>
</tr>
<tr>
<td>Other major durables for recreation and culture (1.63)</td>
<td></td>
</tr>
<tr>
<td>Package holidays (19.80)</td>
<td></td>
</tr>
<tr>
<td>Medical products, appliances and equipment (16.33)</td>
<td></td>
</tr>
<tr>
<td>Clothing (44.92)</td>
<td></td>
</tr>
<tr>
<td>Footwear, including repairs (10.17)</td>
<td></td>
</tr>
<tr>
<td>Hospital services (5.63)</td>
<td></td>
</tr>
<tr>
<td>Actual rentals (including imputed rentals for housing), Water supply and miscellaneous services relating to the dwelling (244.15)</td>
<td></td>
</tr>
<tr>
<td>Operation of personal transport equipment (82.22)</td>
<td></td>
</tr>
<tr>
<td>Electricity, gas and other fuels (47.02)</td>
<td></td>
</tr>
<tr>
<td>Education (6.66)</td>
<td></td>
</tr>
<tr>
<td>Accommodation services (9.71)</td>
<td></td>
</tr>
<tr>
<td>Alcoholic beverages (16.86)</td>
<td></td>
</tr>
<tr>
<td>Tobacco (19.87)</td>
<td></td>
</tr>
<tr>
<td>Household textiles (4.79)</td>
<td></td>
</tr>
<tr>
<td>Other recreational items and equipment, gardens and pets (18.89)</td>
<td></td>
</tr>
<tr>
<td>Recreational and cultural services (28.79)</td>
<td></td>
</tr>
<tr>
<td>Furniture and furnishings, carpets and other floor coverings (33.61)</td>
<td></td>
</tr>
<tr>
<td>Purchase of vehicles (37.26)</td>
<td></td>
</tr>
<tr>
<td>Audio-visual, photographic and information processing equipment (22.66)</td>
<td></td>
</tr>
<tr>
<td>Out-patient services (13.50)</td>
<td></td>
</tr>
<tr>
<td>Communication (25.21)</td>
<td></td>
</tr>
<tr>
<td>Household appliances (11.28)</td>
<td></td>
</tr>
<tr>
<td>Goods and services for routine household maintenance (8.22)</td>
<td></td>
</tr>
<tr>
<td>Maintenance and repair of the dwelling (11.49)</td>
<td></td>
</tr>
<tr>
<td>Transport services (19.17)</td>
<td></td>
</tr>
<tr>
<td>Catering services (36.86)</td>
<td></td>
</tr>
<tr>
<td>Other goods and services (70.23)</td>
<td></td>
</tr>
</tbody>
</table>
The 9 categories of the West German CPI (1962-1990)

- Food, Beverages, Tobacco (224.9)
- Housing (191.93)
- Education, Culture, Recreation (91.66)
- Furnishings and Household Equipment (76.99)
- Personal Care, Accommodation Services and Miscellaneous Goods and Services (65.9)
- Clothing and Footwear (73.83)
- Electricity, Gas and Other Fuels (53.41)
- Health (53.53)
- Transport and Communication (167.85)

B Inflation forecast variables

- $P$ Consumer Price Index; not seasonally adjusted, 2000=100, Federal Statistical Office Germany;
- $U$ Unemployment rate (% of civil labor force), not seasonally adjusted, OECD-Database;
- $M1$ Monetary aggregate M1 (since 1999 German contribution to M1 of the Euro-area), not seasonally adjusted, Deutsche Bundesbank;
- $M3$ Monetary aggregate M3 (since 1999 German contribution to M3 of the Euro-area), not seasonally adjusted, Deutsche Bundesbank;
- $E$ Exchange rate (monthly average), German Mark to US $, Datastream;
- $Y$ Industrial production, Volume Index, not seasonally adjusted, 1995=100, OECD-Database;
- $MS$ Manufacturing sales, Volume Index, not seasonally adjusted, 1995=100, OECD-Database;
- $IS$ Overnight rate (monthly average), IMF International Financial Statistics;
- $IL$ Government bond yield (Umlaufsrendite, monthly average), IMF International Financial Statistics;