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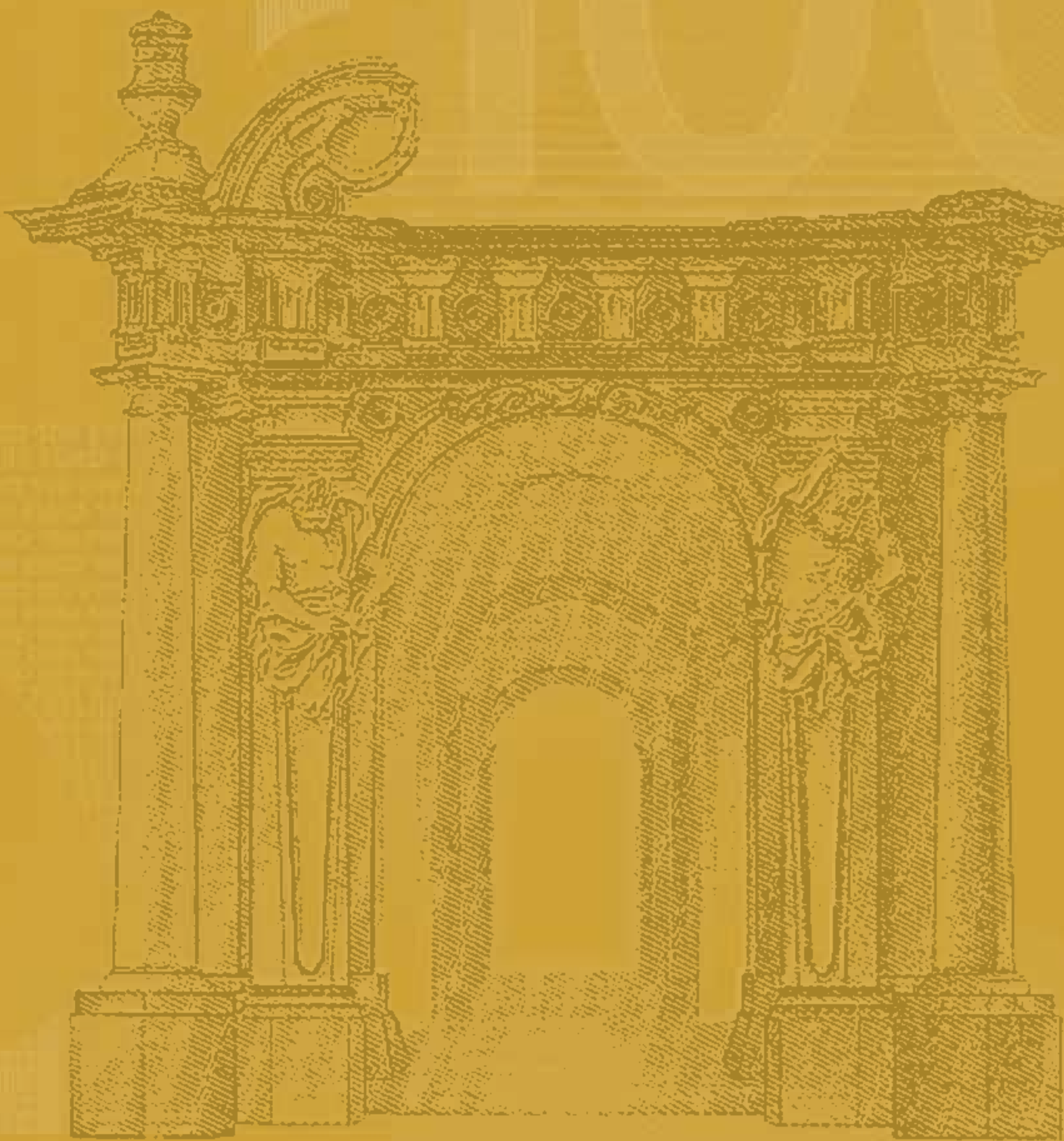

INFLATION AND RELATIVE PRICE ASYMMETRY

by Attila Rátfai

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CONTENTS

Abstract

By placing store-level price data into bivariate Structural VAR models of inflation and relative price asymmetry, this study evaluates the quantitative importance of idiosyncratic pricing shocks in short-run aggregate price change dynamics. Robustly to alternative definitions of the relative price, identification schemes dictated by two-sided (S,s) pricing theory and measures of asymmetry in the relative price distribution, idiosyncratic shocks explain about 25 to 30 percent of the forecast error variance in inflation at the 12-month horizon. While the contemporaneous correlation between inflation and relative price asymmetry is positive, idiosyncratic shocks lead to a substantial build-up in inflation only after two to five months following the initial disturbance.

Key words: (S,s) Pricing, Relative Price, Inflation, Structural VAR

JEL Classification: E31

Non-technical summary

This study examines the dynamic interaction between inflation and relative price asymmetry, in a novel, store-level sample of a selected group of good and service prices. To improve on our understanding of short-run inflation dynamics, the main focus is on characterizing the contribution of idiosyncratic pricing shocks to aggregate price changes.

The empirical approach adopted is specific in three respects, all of them being related to central themes in models of optimal two-sided (S,s) pricing rules and their aggregation. First, the empirical specification is focusing on the inflation-asymmetry, instead of the more traditional inflation-dispersion relationship in studying the relationship between inflation and relative price variation. Second, to control for their dynamic and simultaneous determination, the paper employs bivariate, Structural VAR (SVAR) models of inflation and relative price asymmetry. Third, to bring measurement closer to the theory motivating the empirical work, the data analysis utilizes a unique sample of highly disaggregated, store-level price quotations recorded in Hungary. The sample consists of a series of large cross-sections of monthly frequency price observations of twenty-seven homogenous consumption items, mostly specific food products and a few services, available over a period of fifty-five months.

In selecting identifying restrictions in the SVAR analysis, two straightforward aggregate implications of two-sided (S,s) models are utilized. First, pure aggregate shocks have no contemporaneous impact on the shape of the relative price distribution. Second, pure idiosyncratic shocks have a mean reverting impact on the aggregate price level; that is, any unit root in the price level is driven solely by aggregate shocks.

To detect the central tendency in the data, the relevant product-specific information is merged in three distinct ways. The first approach combines data. Under the assumption that all store level prices are drawn from the same underlying distribution, a bivariate SVAR model of pooled relative price asymmetry and aggregate inflation is estimated. The other two approaches combine information from product-specific bivariate SVAR systems by reporting on the quartiles, primarily the median of the product-specific estimates. The second empirical approach does so without restrictions on the reduced form VAR model, while the third approach makes the system a panel SVAR by constraining the reduced form dynamics to be the same across products.

Robustly to alternative definitions of the relative price, identification schemes dictated by two-sided (S,s) pricing theory, measures of asymmetry in the relative price distribution and the Bryan and Cecchetti (1999) criticism, the main findings are the following. The results in general provide support for theories emphasizing the aggregate importance of idiosyncratic elements in lumpy and heterogeneous microeconomic behavior. More specifically, idiosyncratic shocks explain a significant portion, 25 to 30 percent of the forecast error variance in inflation at the 12-month horizon. While the contemporaneous correlation between inflation and relative price asymmetry is positive in monthly frequency data, idiosyncratic shocks lead to a substantial build-up in inflation after two to five months following the initial disturbance. The impact multiplier is ambiguous in sign. Overall, the findings suggest that a theory of short-term inflation dynamics needs to account for both aggregate and idiosyncratic pricing shocks when explaining the joint determination of inflation and relative price asymmetry.

1 INTRODUCTION

This study examines the dynamic interaction between inflation and relative price asymmetry, using a novel, store-level price data set for a selected group of goods and services. To improve on our understanding of short-run inflation dynamics, still a largely puzzling issue in macroeconomics[1] the main focus is on characterizing the contribution of idiosyncratic pricing shocks to aggregate price changes.

In a univariate context, the postulated correlation between various measures of cross-sectional relative price variation and aggregate inflation is an old and extensively studied issue in macroeconomics; its history goes back at least to the seminal work of Mills (1927). Initiating a voluminous literature, one of the first related studies in the modern era is Vining and Elwertowski (1976). By examining various forms of regression equations with some measure of cross-sectional variability in sector-specific inflation rates on the left and inflation on the right hand side, along with many subsequent investigations, they find that cross-sector price variability is positively related to inflation.[2] This result is often interpreted as being indicative of the welfare costs of inflation.

There exist however several hitherto overlooked aspects of the comovement between inflation and cross-sectional relative price variation. The present paper seeks to make advance along three dimensions, all of them related to implications of modeling microeconomic pricing decisions in a two-sided (S,s) framework. First, the new "supply side" theory of inflation contends that the asymmetry in the distribution of idiosyncratic pricing shocks impacts on the overall price level.[3] The theory cast in a two-sided (S,s) pricing framework also predicts that the dispersion in relative prices has no independent impact on inflation, only in interaction with

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[1] See, for instance, Atkeson and Ohanian (2001).
[2] For dissenting views, see Reinsdorf (1994) and Silver and Ioannidis (2001). Weiss (1993) gives a comprehensive survey of the early literature.

asymmetry. Existing evidence seems to corroborate this theory; inflation and measures of relative price asymmetry are positively associated in the data.[4] Driven by these considerations, the empirical specification adopted in this paper is focusing on the inflation-asymmetry, instead of the more traditional inflation-dispersion relationship.

Second, while the traditional literature tends to emphasize the direction of causality running from inflation to relative price variation, (S,s) pricing theories motivating the current analysis do not rule out this traditional channel; they instead point to the possibility of the reverse direction. To control for the dynamic and simultaneous determination of inflation and relative price variation, this study employs bivariate, Structural VAR (SVAR) models of inflation and relative price asymmetry. The main virtue of the SVAR approach is that it is able to isolate the underlying disturbances, without imposing strong constraints on the joint dynamics of the endogenous variables involved.

Finally, in order to bring measurement closer to the theory motivating the study of relative price variation, the empirical analysis utilizes a sample of highly disaggregated, store-level price quotations. Using store level price data in this context constitutes a major departure from much of the literature on relative price variation and inflation. Probably caused by the paucity of appropriate data, most related studies focus on the cross-sectional variation in sector- or city-level price indices and neglected relative price measures based on establishment level observations.[5]

To preview the main results, idiosyncratic pricing shocks account for a substantial share of fluctuations in inflation. Still, while the partial correlation between inflation and relative price

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[3] See Ball and Mankiw (1995).
[4] Findings in Amano and Macklem (1997), Ball and Mankiw (1995) and Suvanto and Hukkinen (2002) suggest that asymmetry is a much more important determinant of inflation than dispersion in relative prices.

asymmetry is positive in monthly frequency data, idiosyncratic shocks generating variation in relative price asymmetry do not result in a contemporaneous surge in inflation: the impact multiplier is ambiguous in sign. Idiosyncratic shocks however lead to a significant response in inflation with a lag of two to five months. Overall, the findings suggest that a theory of short-term inflation dynamics needs to account for both aggregate and idiosyncratic pricing shocks in explaining the joint determination of inflation and relative price asymmetry.

The rest of the paper proceeds as follows. To motivate the identification strategy in the VAR model, Section 2 explains some implications of two-sided (S,s) pricing models. Section 3 takes up measurement issues. The microeconomic price data are described in Section 4. The VAR model together with specification tests is discussed in Section 5. The baseline results are presented in Section 6, while the VAR results are covered in Section 7. Section 8 provides a more detailed discussion of a study with a close bearing on the present research. Conclusions are offered in Section 9.

2 IDENTIFICATION

Microeconomic evidence suggests that pricing decisions exhibit elements of both lumpiness and heterogeneity: stores tend to keep their prices unaltered for extended periods of time and when they change them, they do so in a non-uniform manner a by non-trivial amounts.[6] The pricing model best matching this description of microeconomic behavior is in turn developed in two-sided (S,s) models. The underlying idea is that stores operating in monopolistically competitive markets and facing fixed cost to price adjustment ("menu cost"), optimally trade off the benefit from adjusting their nominal price against the corresponding cost of adjustment. The central

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[5] Notable exceptions include Danziger (1987), Fengler and Winter (2000), Lach and Tsiddon (1992), Reinsdorf (1994) and Tommasi (1993).

object in the model is the relative price, the potentially non-zero deviation between the actual and the target log price. Aggregate and idiosyncratic shocks continuously alter the relative price through impacting on the target price. If pricing shocks are small, the optimal policy implies that the nominal price is temporarily held constant, while the relative price keeps gradually moving in between the optimally determined adjustment thresholds, *S* and *s*. Actual pricing action is only triggered when the relative price gets sufficiently eroded to surpass one of these thresholds.

In selecting identifying restrictions in the SVAR analysis, two straightforward aggregate implications of two-sided (S,s) models are of particular interest. First, the separation of aggregate and idiosyncratic pricing shocks to the target price is central to the modeling of optimal pricing decisions in the presence of fixed adjustment costs. Indeed, as illustrated in Figure Ia, a pure aggregate shock common to all price-setters has no contemporaneous impact on the shape of the pre-adjustment relative price distribution. The aggregate shock displaces all target prices, thus relative prices identically in the state space, leaving the shape of the distribution unaltered, while moving some relative prices outside the adjustment threshold, prompting for pricing action and relative price adjustment in these instances.[7] After repricing has taken place following an aggregate inflationary shock, assuming only symmetric idiosyncratic shocks in the rest, prices keep rising and thus inflation is positive but decelerates. The readjustment process is shown in Figure Ib. As demonstrated in Ball and Mankiw (1994) and Tsiddon (1993), this argument extends to a non-zero drift in the target price. If a positive trend in the target process continuously erodes the relative price, making the non-adjustment band asymmetric with a more heavily populated downward portion, even symmetric idiosyncratic shocks make the relative price distribution right skewed, thus breeding more frequent nominal price increases than price

[6] See, for instance, Bils and Klenow (2002), Lach and Tsiddon (1992) and Rátfai (2004). For a recent survey, see Wolman (2000).
[7] See also Caballero, Engel and Haltiwanger (1995).

decreases, validating positive inflation.[8] All in all, the main implication of this reasoning for identification in the VAR analyses to come is simply that a pure aggregate shock has no contemporaneous impact on the shape of the relative price distribution.

Second, as exemplified by Ball and Mankiw (1995) in a model with fixed price adjustment cost and no trend in the target price, the average price level is determined by the distribution of shocks to the target price. If for example the distribution of shocks is of mean zero in population but happens to be right-skewed in realization, more firms are likely to raise the nominal price, making the aggregate price level to rise. A similar argument applies to left-skewed distributions and the decline in the aggregate price level. Using numerical simulations, Ball and Mankiw show that the reasoning on the positive relation between the asymmetry in the distribution and aggregate price changes extends to the shape of relative price distribution itself.

While it in general remains silent about the temporal relationship between inflation and relative price asymmetry, the fundamentally static analysis in Ball and Mankiw (1995) still points to an important dynamic corollary: non-symmetric realizations of purely idiosyncratic shocks can have no long run impact on the aggregate price *level*. To see why this is the case, consider a situation where trend inflation is zero and there are only idiosyncratic shocks. Also assume that the population distribution of shocks and thus the inaction band and the distribution of relative prices are symmetric. The realization of idiosyncratic shocks however is not always symmetric. For instance, in a period dominated by a few large inflationary shocks together with many smaller deflationary ones, the number of stores passing the lower adjustment boundary exceeds the number of stores passing the upper boundary, making the aggregate price level rise. After actual price changes have been completed, however, relative prices tend to get bunched closer to the

[8] A multi-sector real business cycle model with an asymmetric input-output structure delivers a similar result, with relative prices being defined as sector-specific inflation rates. See Balke and Wynne (2000).

upper, deflationary end of the distribution than to the lower, inflationary one. Then, assuming the symmetry of shocks in the upcoming periods, fewer price increases than price cuts are expected to take place, implying a fall in price level. The adjustment process with alternating periods of rising and falling prices continues until the symmetric steady state distribution of relative prices is restored and the aggregate price level returns to its original level. Assuming idiosyncratic shocks with a direct impact on inflation as an example, this dynamic adjustment process is illustrated in the four panels of Figures Ic and Id. In sum, the reasoning suggests that pure idiosyncratic shocks can have a mean reverting impact on the aggregate price level; that is, any unit root in the price level is driven solely by aggregate shocks.

3 MEASUREMENT

One of the missions of this study is to bring measurement closer to theory motivating the empirical analysis of relative prices and their variation. Two related observations in this regard bear on the data analysis to follow. First, studies on relative price variability tend to bury the diverse outcomes of microeconomic pricing decisions into measures of *inter*-sector or *inter*-city variability of sector level inflation rates. However, unless stores' pricing policies are fully synchronized within sectors or cities, merely averaging microeconomic price observations resulting in sector or city level index numbers could mask the pervasive within-sector differences in price setting decisions. In addition, as they genuinely reflect microeconomic heterogeneity, empirical work drawing on (S,s) pricing models should utilize highly disaggregated price data reflecting regularities in heterogeneous microeconomic behavior.

Second, different models of microeconomic price heterogeneity do not necessarily have observationally equivalent implications for the measurement of cross-sectional price variation. Indeed, it is useful to distinguish three distinct concepts in this regard: variation in price *changes* (inflation variability), in price *levels* (price dispersion) and in the log *deviation* between the actual and the target price (relative price variation). The vast majority of empirical papers examine the cross-sector standard deviation of the *change* in sector level price indices. As opposed to the notion of cross-sector price dispersion that would just compare apples to oranges, the price change measure has clearly the benefit of capturing a meaningful object. At the same time, it does not seem to correspond to any of the theoretical concepts motivating the study of the correlation between inflation and price variation. Indeed, from the perspective of (S,s) models, it is the notion of relative price variation that is relevant for the purposes of empirical work.[9]

In light of these considerations, the present study employs a definition of relative price variation, which is not only feasible to measure but also consistent with the theory motivating the study of the relationship between inflation dynamics and relative price variability. In particular, the relative price in store i of product j at time t, before any potential adjustment at t is defined as the log deviation of the lagged actual individual price level from the product-specific mean price, $z_{ijt} = p_{ij,t-1} - p_{jt}$. That is, the mean nominal price serves as a proxy to capture the target price in the data. While this approach to measure the unobserved target is somewhat *ad hoc*, on the one hand, it still conforms to standard models of monopolistic competition, in which the optimal individual price is proportional to the aggregate of prices. On the other hand, making the target equal to the mean is dictated by the data constraint that the sample used in the current analysis does not allow for tracking individual prices over time, as not being longitudinal.[10] Finally, taking the mean of prices (or price changes) as a reference point in defining relative prices (or price differentials) is *the* standard approach in the related empirical literature.[11]

[9] Search models tend to have mostly static implications for the product-specific dispersion in price levels.
[10] If a true panel were available, more structural procedures to uncover the target price could be devised. See Rátfai (2003).
[11] See, for example, Ball and Mankiw (1995), Lach and Tsiddon (1992), Silver and Ioannidis (2001) and Weiss (1993).

Given this reasoning, first, the product-specific mean price is defined as an equally weighted index of nominal prices of product *j* at time *t*, computed as $P_{jt} = \frac{1}{n_{jt}} \sum_{i=1}^{n_{jt}} P_{ijt}$, where P_{ijt} is the nominal price in store *i* of product *j* at time *t* and n_{jt} is the number of stores for product *j* in month *t*. The log mean price is then obtained as $p_{jt} = \ln(P_{jt})$, and the inflation rate for product *j* in month *t* is computed as

$$\Pi_{jt} = p_{jt} - p_{j,t-1}\,.$$

The skewness statistic capturing the asymmetry in the relative price distribution in product *j* is the defined as

$$S_{jt} = \frac{n_{jt}}{(n_{jt}-1)(n_{jt}-2)} \sum_{i=1}^{n_{jt}} \left(\frac{z_{ijt} - z_{jt}}{V_{jt}^{1/2}} \right)^3 ,$$

where n_{jt} is the number of price observations, z_{jt} is the mean and V_{jt} is the variance of relative prices for product *j*, in month *t*.

The corresponding definitions at the aggregate level readily follow. Aggregate inflation in the sample is computed as the weighted mean of product-specific inflation rates, $\Pi_t = \frac{1}{J} \sum_{j=1}^{J} w_j \left(p_{jt} - p_{j,t-1} \right)$, where the w_j terms are expenditure weights summing up to one.[12] The skewness statistic in pooled relative prices is defined as

$$S_t = \frac{N_t}{(N_t - 1)(N_t - 2)} \sum_{j=1}^{J} \sum_{i=1}^{n_{jt}} \left(\frac{z_{ijt} - z_t}{V_t^{1/2}} \right)^3 ,$$

where $N_t = \sum_{j=1}^{J} n_{jt}$ is the total number of price observations, z_t is the mean and V_t is the variance of all product-level relative prices pooled together in month t.

4 DATA

The empirical analysis utilizes a unique sample of store level consumer prices recorded in Hungary. The data set consists of a series of cross-sections of monthly frequency price observations of twenty-seven homogenous consumption items, mostly specific food products and a few services. It is drawn from the larger sample of consumer prices collected for the monthly computation of the CPI by the Central Statistical Office, Hungary. Products are selected from the full CPI database with an eye to obtaining very narrowly defined (according to size, branding, type and flavor), continuously available items with negligible variation in non-price characteristics. Table I lists all the products together with the expenditure weight attached to them in computing the aggregate CPI and their relative expenditure weight in the current sample. The table also reports on the mean and standard deviation of product-level and aggregate monthly inflation rates.

The data are available from 1992:1 until 1996:7 at the monthly frequency. In each month, there are 100-150 price observations (on average about 125) for each product. The number of stores is stable with an average standard deviation of less than 3 over time. The data are recorded in geographically dispersed locations including all the 19 counties in Hungary and the capital

[12] Using the unweighted inflation measure produces virtually identical results.

city, Budapest. Although stores in the sample are identified only by their location and are not longitudinally matched, the data collectors are formally instructed by the CSO to keep the composition of stores as stable as possible.

Despite the turbulent economic environment during the early years of economic transition, inflation was relatively moderate in Hungary. The year-to-year changes in the monthly aggregate CPI and its food component are plotted in Figure II. The graph shows that after an initial burst, inflation somewhat decelerated by the beginning 1994. After reaching a minimum of about 15 percent, it turned on an increasing path eventually reaching about 30 percent at its peak in early 1995. Starting with the second quarter of 1995, shortly after the announcement and implementation of a policy adjustment package in March 1995, a disinflationary trend took effect again.

5 THE BASELINE CASE

Before proceeding to the dynamic analysis, it is first useful to document the static relationship between inflation and relative price asymmetry. Motivated by their two-sided (S,s) theory, Ball and Mankiw (1995) use a simple univariate regression framework to test the impact the asymmetry in the distribution of sector-specific inflation rates exerts on aggregate inflation in U.S. data. Robustly to alternative definitions of asymmetry, they find that asymmetry has a sizeable and statistically significant contemporaneous effect on inflation.

To replicate their main result in the current sample, consider first the skewness in the distribution of all relative prices pooled together as a measure of relative price asymmetry. The OLS estimation in the preferred specification results in the estimated equation of

$$\begin{array}{lllll} \Pi_t = & 1.118 & + \; 0.538\,\Pi_{t-1} & + \; 1.410\,S_t, & \bar{R}^2 = 0.291 \\ & (0.427) & \quad (0.117) & \quad (0.931) & \end{array},$$

where Π_t denotes aggregate inflation in the sample and S_t denotes the skewness statistic as defined above. When skewness is replaced by an alternative measure of asymmetry, the mean-median difference scaled by the standard deviation, the coefficient on asymmetry becomes significant and the goodness-of-fit improves.[13] Next, when relative price asymmetry is proxied by skewness in cross-product relative inflation rates (s_t), the regression results remain qualitatively unchanged[14]

$$\underset{}{\Pi_t} = \underset{(0.339)}{0.566} + \underset{(0.112)}{0.442}\,\Pi_{t-1} + \underset{(0.177)}{0.565}\,s_t, \quad \overline{R}^2 = 0.382 .$$

Overall, besides that the univariate results provide some evidence of persistence in inflation, the contemporaneous relationship between relative price asymmetry and inflation appear to be positive both in sector and microeconomic level data. What is less clear however is the relative importance of idiosyncratic versus aggregate pricing shocks in determining the positive association, especially when the simultaneous determination between the two variables, a central theme in the (S,s) literature is accounted for. This is the issue examined next.

[13] The estimated equation takes the form of

$$\Pi_t = \underset{(0.449)}{0.168} + \underset{(0.136)}{0.377}\,\Pi_{t-1} - \underset{(0.821)}{1.741}\,S_t, \quad \overline{R}^2 = 0.319 .$$

[14] With the exception of Lach and Tsiddon (1992), other studies examining relative price asymmetry in relation to inflation in a univariate statistical model, such as Amano and Macklem (1997), Ball and Mankiw (1995) and Blejer (1983) measure asymmetry in sector-specific inflation rates, as opposed to establishment level relative prices.

6 SVAR SPECIFICATION AND ESTIMATION

To fix notation, first, consider the bivariate, SVAR system of the stationary variables of inflation (Π_{jt}) and relative price skewness (S_{jt}) specified separately for each product j[15]

$$y_{jt} \equiv \begin{bmatrix} \Pi_{jt} \\ S_{jt} \end{bmatrix} = \begin{bmatrix} 0 & G^0_{\Pi S} \\ G^0_{S\Pi} & 0 \end{bmatrix} \begin{bmatrix} \Pi_{jt} \\ S_{jt} \end{bmatrix} + \begin{bmatrix} B_{\Pi\Pi}(L) & B_{\Pi S}(L) \\ B_{S\Pi}(L) & B_{SS}(L) \end{bmatrix} \begin{bmatrix} \Pi_{jt} \\ S_{jt} \end{bmatrix} + \begin{bmatrix} \varepsilon^{\Pi}_{jt} \\ \varepsilon^{S}_{jt} \end{bmatrix} \equiv G^0 y_{jt} + B(L) y_{jt} + \varepsilon_{jt}$$

where $B(L)$ is a pth degree matrix polynomial in the lag operator L with $B(L) = 0$. The diagonal elements of G^0 are normalized to zero. Dynamics in inflation and relative price skewness are assumed to be governed by contemporaneous and past realizations of an unobservable vector of serially uncorrelated and mutually orthogonal structural innovations $\varepsilon_{jt} = [\varepsilon_{jt}^{\Pi}, \varepsilon_{jt}^{S}]$ with variance-covariance matrix $D = \mathrm{E}(\varepsilon_{jt}\varepsilon_{jt}')$. The orthogonality assumption implies that the off-diagonal elements of the variance-covariance matrix are zero. In economic terms, ε_{jt}^{Π} is interpreted as a pure aggregate shock affecting relative prices identically and ε_{jt}^{S} as the outcome of store-specific, idiosyncratic disturbances to pricing policies. The off-diagonal elements of the G^0 matrix represent short-run multipliers, capturing the contemporaneous impact of structural shocks to endogenous variables: $G^0_{S\Pi}$ aggregate shocks to relative price skewness, $G^0_{\Pi S}$ idiosyncratic shocks to inflation.

Under standard regularity conditions, the structural model is transformed to the reduced form autoregressive one as

[15] To ease notation, product-specific parameter indices are omitted. The VAR model at the aggregate level is specified analogously.

$$y_{jt} = H(L) y_{jt} + u_{jt}\,.$$

Here u_{jt} represents reduced form innovations with an unrestricted variance-covariance matrix Σ. Reduced form innovations are linearly related to structural ones by

$$\varepsilon_{jt} = B^0 u_{jt}$$

where $B^0 = I - G^0$.

From the Wold moving average representation $y_{jt} = C(L) u_{jt}$ with $C(L) = (I - H(L))^{-1}$, the infinite order, structural form moving average representation is obtained as

$$y_{jt} = M(L)\varepsilon_{jt}$$

where $M(L) = C(L)(B^0)^{-1}$. As it is related to long-run multipliers, forecast error variance decomposition, impulse responses, this form of the system is of particular interest for model identification and economic inference. First, the long-run multipliers in the system reflecting the cumulative response in endogenous variables to structural shocks correspond to the appropriate elements of the $M(1)$ matrix. Second, the forecast error variance decomposition (*FEVD*) function measuring the quantitative importance of a particular structural shock is also derived from elements of the $M(L)$ matrix. Formally, the statistic gives the percentage of the k-step-ahead (12-month-ahead in the current paper) forecast error variance in variable j attributable to the structural shock i as $FEVD_{ij,k} = d_i^2 \sum_{h=0}^{k-1} m_{ij,h}^2 \Big/ \sum_{l=1}^{I} \left[d_l^2 \sum_{h=0}^{k-1} m_{lj,h}^2 \right]$, where $m_{ij,h}$ is the (i,j)th entry of the infinite moving average matrix $M(h)$, and d_i^2 is the diagonal element of the D matrix comprising of the variance of the structural innovations. Finally, orthogonalized impulse response functions provide an answer to the following question: how does a current unitary structural shock make the econometrician revise the forecast of future realizations of endogenous variables. The answer here is recovered from the appropriate entries of the $M(L)$ matrix again.

To exactly identify the four primitive structural parameters (two of them contained in B^0 and another two in D) from the three reduced form parameter estimates (the elements of Σ), it is necessary to place one extra restriction on the set of structural parameters. The discussion in Section II suggests two alternative restrictions, both of them amounting to particular economic interpretations of the primitive shocks governing the dynamics of the endogenous variables. First, the asymmetry measure in the distribution of relative prices is assumed to be contemporaneously invariant to shocks common to stores resulting in the "Short Run" (*SR*) identifying restriction, $B^0_{S\Pi} = 0$. Alternatively, idiosyncratic shocks have only transitory impact on the aggregate price *level* thus aggregate inflation is governed only by aggregate shocks in the long run, giving rise to the "Long Run" (*LR*) identifying restriction, $M_{\Pi S}(1) = 0$.

6.1 Pre-tests

The VAR model described above assumes the stationarity of the series involved. The stochastic properties of the product-specific inflation and relative price skewness series are examined in a sequence of univariate unit root tests for all the twenty-seven products. The testing procedure is the Augmented Dickey-Fuller (ADF) test used in conjunction with the Schwartz Information Criterion for selecting the number of lags. By default, the maximum number of lags allowed is 12. In general, the results in general suggest the absence of unit root both in the inflation and the skewness series.[16] Three issues in unit-root testing deserve special attention, each of which having a bearing on model specification.

First, with the exception of the skewness variables *s10603* and *s52366*, visual inspection suggests no evidence of a time trend. Therefore, with the exception of these series, the ADF stationarity tests and the VARs do not include a deterministic time trend. Second, preliminary tests did not reject the presence of a unit root in three of the skewness variables, *s10301*, *s14424* and *s66105*. However, eyeballing the series also suggests that they are likely to contain a structural break.[17] To test for the stationarity of these series, along with all the other ones for possible structural breaks, the modified unit root test of Perron (1997) is used. The resulting *t*-statistics and autoregressive roots indicate that the variables of *s10301*, *s14424* and *s66105* can indeed be viewed as stationary with a structural break. The test indicates no structural break in the other series.

Finally, upon further inspection, some of the inflation series seem to exhibit seasonal fluctuations. To confirm this impression, a series of deterministic seasonal regressions are performed with inflation on the left and seasonal dummies on the right hand side. The inflation series with at least two statistically significant monthly dummy coefficients, fifteen out of twenty seven are characterized as ones containing a deterministic seasonal component.[18] To check whether the stochastic element in the inflation series is stationary, a set of standard ADF tests for

[16] The relevant ADF *t*-statistics and the largest autoregressive parameters indicate that the log price level series cannot be rejected to have a unit root. ADF tests also show that the presence of unit-root in the stochastic component of the series can be rejected in all of the series when deterministic seasonal effects are controlled for. The details of the test results are presented in Tables AI through AIII.

[17] Perron (1997) shows that not accounting for a break in the series when it is actually present may result in false acceptance of the unit root in standard ADF tests. To address this issue, he devised a modified ADF procedure choosing endogenously the break point in the series and provided the appropriate critical values for the *t*-statistic. The procedure is based on a regression equation that includes dummies for capturing the break in the series, potentially of three different kinds, a pure intercept, a pure slope or a combination of the two.

the estimated first stage residuals are conducted. Test results show no evidence of non-stationarity in the residuals.

Overall, thirteen of the VAR systems are specified with seasonal dummies, one with a pure time trend, one with a time trend and seasonal dummies, two with dummies for a structural break, and one with dummies for a structural break and deterministic seasonal dummies as well. Nine of them exhibit none of these peculiarities and are estimated with only a constant added to the endogenous variables and their lagged values. Consistent estimates of the reduced form VAR parameters are obtained by Ordinary Least Squares. The number of lags included in each product-specific system is dictated by a series of Likelihood Ratio tests.

7 SVAR RESULTS

The large number of ways the information in the data could be grouped, at the minimum, an overall combination of twenty-seven products, four categories of inference (short-run multipliers, long-run multipliers, forecast error variance decompositions and impulse response functions), two identification schemes, results in a practically non-digestible flow of information.[19] In order to detect the central tendency in the data, the relevant information is merged in three distinct ways.

7.1 Pooled Data

The first approach combines data. Under the assumption that all store level prices are drawn from the same underlying distribution, the parameter estimates are obtained from the bivariate SVAR

[18] All of these series have an R^2 statistic larger than 0.4.
[19] In preliminary calculations, I also experimented with examining food and non-food prices separately. The results for the two groups were similar, so this issue is not elaborated any further.

model of pooled relative price skewness, S_t, and aggregate inflation, Π_t, as endogenous variables, and exogenous seasonal dummies.[20]

The short run and the long run structural multipliers estimated in the pooled system under the two distinct identification assumptions are displayed in the top panel of Table II. Notice first that the short-run coefficients in the first two columns of the table indicate a significant deflationary impact of idiosyncratic shocks. The forecast error variance decompositions are displayed in the top panel of Table III. The results indicate a sizeable relative share of idiosyncratic shocks; under the *LR* restrictions, the relevant figure is as large as 64 percent. The variance decomposition figures are especially remarkable as idiosyncratic shocks changing the shape of the pre-adjustment relative price distribution do not have to lead any aggregate response on impact, they may simply reshuffle relative prices in the middle of the density without pushing any of them beyond the adjustment boundaries.

Figures IIIa and IIIb portray the impulse response functions in the pooled data. Most importantly, independently of the identification assumption, one can detect a statistically significant inflationary effect of idiosyncratic shocks, taking place at the third month following the initial disturbance.[21] Corresponding to the estimated short-run multipliers, there also appears to be a clear initial deflationary effect in the first month in the *SR* identification case and the second month in the *LR* identification case. In addition, the impulse responses show statistically significant positive responses in skewness to idiosyncratic shocks and inflation to aggregate

[20] ADF-test results confirm that both series are stationary with no time trend and structural break.
[21] The 90 percent confidence bands using the Runkle (1987) bootstrap procedure with 500 repetitions are shown on the graphs.

shocks on impact. These latter impulse responses however show relatively little persistence, lasting for only one to two months.[22]

Finally, imposing both identifying restrictions on the bivariate system results in an over-identified VAR model. To test for the relative merit of the two restrictions, a set of exclusion tests are conducted using the pooled specification. The resulting *t*-test statistic indicates that the restriction of the lack of impact of aggregate shocks to relative price skewness cannot be rejected. Similarly, the *F*-test statistic indicates non-rejection for the *LR* over-identifying restriction.

7.2 Product-Specific Model

Next, to maintain the product-level approach to analyzing the aggregate consequences of product-level heterogeneity in pricing behavior, the quartiles, primarily the median of product-specific SVAR estimates are analyzed. To the extent that product-specific shocks generating the relatively high volatility in product level inflation rates are accounted for in this approach, the resulting parameter estimates can be thought of as establishing a lower bound on the importance of idiosyncratic shocks in more aggregated data. The specific findings are as follows.

The middle panel in Table II shows the median short-run and the long-run cross-multipliers, along with a measure of the extent of the heterogeneity in point estimates, the cross-product standard deviation. First, skewness responds positively to idiosyncratic and inflation to aggregate shocks again. The effects appear to be more persistent here than in the pooled case, they last for three to four months. Next, irrespectively of the identification assumption used, the contemporaneous impact of an aggregate shock to relative price skewness is close to zero. Non-parametric confidence interval sign-tests indicate that this result is statistically significant at the 5

[22] As suggested by the impulse response graphs in Figures AIa through AIIb in the Appendix, these findings are robust to alternative definitions of the relative price and alternative measures of asymmetry in the relative price distribution.

percent level.[23] Finding a small contemporaneous response under the *LR* constraint is reassuring as it suggests that the *SR* identification assumption is a sensible one. The contemporaneous impact of idiosyncratic shocks to inflation is more ambiguous. Although the median of the contemporaneous estimates have positive signs under both identification constraints, the sign test shows that none of them are significantly so. Finally, the absence of a significant long-run response of inflation to idiosyncratic shocks under the *SR* identification scheme is comforting again; it indicates that the *LR* identification assumption is borne out by the data.[24]

The median estimates of the forecast error variance decompositions are displayed in the middle panel of Table III, together with the corresponding cross-product standard deviations of the estimated coefficients. The figures show that idiosyncratic shocks explain about 19 to 26 percent of the variation in inflation forecasts at the product level.[25] Idiosyncratic pricing shocks appear to be fundamental determinants of the forecast error variance in relative price skewness. Under the *LR* identification assumption, their median contribution is 66 percent, while in the *SR* case it is more than 80 percent.

Figures IVa and IVb display the median of the product-specific impulse responses of inflation and relative price skewness to one standard deviation idiosyncratic and aggregate shocks, under both identification schemes. Besides the median, the graphs also display the upper and the lower quartiles of the parameter estimates. First, the top-left panels in the graphs show the 12-month-ahead impulse response of inflation to idiosyncratic shocks. The two impulse response functions portray a remarkably uniform picture with a sizeable peak at about three to four months after the initial shock having occurred. Sign-tests indeed indicate that the positive response at the

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[23] The sign test determines a confidence interval for the median and evaluates the null hypothesis that the median of the point-estimate is not different from zero against a two-sided alternative. See Gibbons and Chakraborti (1992).
[24] Mean reversion in response to idiosyncratic shocks applies to the price level as well.

fourth month is statistically different from zero under both identification assumptions. The impulse responses of skewness to idiosyncratic shocks and inflation to aggregate shocks are significantly positive for a number of periods following the initial shock.

7.3 Panel SVAR

Assuming similar price dynamics across the different product markets, an alternative way of combining information from the individual dynamic systems is to estimate the product-specific VAR models together as a panel. That is, here the reduced form slope parameters are constrained to be the same across products, capturing aggregate effects by generating cross-product homogeneity in price dynamics.

In practice, the VAR systems are estimated as two separate panels together with all the possible exogenous variables by standard Dummy Least Square (DLS) methods. The dependent variable in one of the panels is comprised of all the inflation series, while in the other panel of all the relative price skewness series. As the panel is a long one with a time dimension well exceeding 30 observations, the DLS approach produces asymptotically unbiased estimates even with lagged dependent variables (cf. Judson and Owen (1997)). For identification in the product level systems, the same *SR* and *LR* restrictions are employed as before. For simplicity, the number of lags is set to be the same for all products. The DLS estimation procedure leads to structural parameter estimates, impulse response functions and forecast error variance decompositions that differ across products. To merge information, the cross-product quartiles of parameter estimates, primarily the median one are highlighted again.

First, the bottom panel in Table II displays the median of forecast error variance decompositions. The figures appear to be virtually identical to the ones obtained in the

[25] The total effect of a structural shock to a particular variable does not have to add up to exactly

unconstrained case; idiosyncratic shocks explain about 26-27 percent of the forecast error variance in inflation. Not surprisingly, the panel constraint results in structural parameters with small cross-product variation, ranging from 3 to 5 percent.

The cross-product quartiles of impulse response estimates are depicted in Figures Va and Vb. Both skewness and inflation respond to idiosyncratic and aggregate shocks with the expected sign. Inflation increases on impact to a positive aggregate shock, converging to its mean in a few months afterwards. The impulse response in skewness to an idiosyncratic shock is more persistent, however, not reverting to zero for at least twelve months. The point estimates are fairly tight again, reflecting the restriction that the reduced form VAR parameters are identical across products.

The main focus is again on the impact of idiosyncratic shocks on inflation. The impulse responses are displayed on the top-left graphs. The first point to notice is that the median responses obtained here are remarkably similar to the ones obtained in the unconstrained product-level specification. The upper and lower quartiles of the estimates also show that the panel VAR impulse responses are fairly homogeneous. The contemporaneous impulse responses depend on the identification restriction, however. Under the *LR* scheme, the response of inflation to idiosyncratic shocks starts out to be negative, then turns into significantly positive at horizons of one to four months and again negative in the next six months. Although the emerging picture is less unequivocal under the *SR* constraint, impulse responses in inflation again turn to positive at horizons of two to four months. The sign tests show that these lagged effects are statistically significant under both restrictions. Finally, the size and persistence of impulse responses appear to be somewhat larger in size in the panel model than in either the pooled or the unconstrained product-specific specifications.

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100 percent for the median measure.

8 AN ANTICIPATED CRITICISM

Bryan and Cecchetti (1999) argue that the empirical results in Ball and Mankiw (1995) documenting a positive and statistically significant contemporaneous correlation between inflation and relative price asymmetry are merely statistical artifacts and suffer from small-sample bias. Their argument is summarized through the following thought experiment. Consider a sample of price changes drawn from a zero-mean symmetric distribution, having a sample mean of zero. By construction, the mean and the skewness of the distribution are uncorrelated. It is straightforward to show that if an extra outlier draw is made from the far positive (or negative) tail of the underlying distribution then it may induce a simultaneous increase (or fall) in measured inflation and skewness. The example illustrates the possibility of a spuriously measured positive unconditional correlation between inflation and the skewness in the distribution of price changes, when the distribution has fat tails. Motivated by these considerations, Bryan and Cecchetti employ Monte Carlo simulations to demonstrate that the suspected bias is an actual concern in the Ball and Mankiw data set. After correcting for the small-sample bias, they actually find negative correlation between skewness and inflation. To explain their findings, Bryan and Cecchetti suggest that if price setters were fully reluctant to cut their nominal prices, a fall in aggregate inflation would induce the distribution of nominal price changes bunching around zero implying increased skewness. They reach the conclusion that "the recent focus on the correlation between the mean and skewness of the cross-sectional distribution of inflation is unwarranted".

Though the criticism of Bryan and Cecchetti appears to invalidate the empirical results of Ball and Mankiw, its basic main thrust is not applicable in the present context. First, Bryan and Cecchetti study the properties of two moments of the *same* underlying distribution of price changes. In contrast, the current paper examines the relationship between variables drawn from two distinct distributions, so there is no reason to hypothesize a bias of the sort described above. Second, any contemporaneous, unconditional correlation between inflation and relative price skewness does not preclude the presence of more complex dynamic relationship between the two variables. Third, indeed, to the extent that they underscore the lagged response of inflation to idiosyncratic shocks and the potential negative contemporaneous correlation between inflation and relative price skewness, the findings of the current paper may be viewed as complementary to findings of Bryan and Cecchetti.[26]

Lastly, the particular construct Bryan and Cecchetti (and Ball and Mankiw) use in measuring the relative price renders their empirical results directly immaterial to the assessment of (S,s) pricing models motivating the study of asymmetry in relative price distributions. There are two points to highlight in this regard, both of them being pertinent to the correspondence between theory and measurement. First, although the idea of downward rigid price adjustment is intuitively appealing, so far only models of the (S,s) family have been able to model rather than merely assume downward price rigidity.[27] Therefore, it is difficult to determine how arguments regarding the distribution of *price changes* would bear on the dynamics in the object envisioned by (S,s) theory, the distribution of *relative prices*. In addition, the kinked demand curve theory of asymmetric price adjustment actually predicts, if anything upward, rather than downward price rigidity. Second, as argued before, identifying relative prices with sector-specific inflation rates misses an important element of microeconomic reality, the within-sector heterogeneity in price setting.

9 CONCLUSIONS

This study has explored some of the consequences of lumpiness and heterogeneity in microeconomic price setting for aggregate price changes. What do we learn from microeconomic

[26] Nonetheless, it remains to be seen how the small-sample bias argument applies to the distribution of relative prices in the current work. This exercise is the subject of future research.
[27] See Ball and Mankiw (1994) and Tsiddon (1993).

price data placed in a dynamic, simultaneous model of inflation and relative price asymmetry? Three points.

First, the baseline univariate estimates confirm and extend the earlier findings of Amano and Macklem (1997), Ball and Mankiw (1995), Lach and Tsiddon (1992) and Suvanto and Hukkinen (2002) in that the contemporaneous correlation between relative price asymmetry and inflation is positive, both in sector and store level data. Second, robustly to alternative identification schemes dictated by two-sided (S,s) pricing theory, definitions of the relative price and measures of asymmetry in the relative price distribution, idiosyncratic pricing shocks explain a substantial portion of the forecast error variance in price changes. This set of results in general provides support for theories emphasizing the aggregate importance of idiosyncratic elements in lumpy and heterogeneous microeconomic behavior.

Finally, while they increase relative price asymmetry on impact, identified idiosyncratic shocks lead to a sizeable response in inflation only with a lag of two to five months. The corresponding direct impact tends to be either insignificant or ambiguous in sign. These results place the univariate evidence into new perspective. The positive contemporaneous correlation between inflation and relative price asymmetry cannot be attributed solely to asymmetry in idiosyncratic shocks, but rather to some non-trivial interaction between the underlying disturbances. The finding also implies that standard two-sided (S,s) pricing models may miss an important element of reality, the delayed response in inflation to pure idiosyncratic shocks. One possible rationale for the hump-shaped response might be that it takes time to collect and incorporate them into pricing plans as idiosyncratic news are less visible for price setters, especially when compared to news affecting all stores equally. This type of sluggishness to acquire and process new information might be thought of as a friction in information perception and processing in the spirit of Woodford (2001) or Mankiw and Reis (2002).

An alternative story could be that non-adjustment (S,s) thresholds are not fixed, instead they respond to pricing shocks. Analogously to Carroll and Dunn (1997), for instance, besides reshuffling the distribution by making relative prices bunch closer to the lower adjustment threshold, a local inflationary shock may also result in the concurrent widening of the non-adjustment (S,s) band. Intuitively, this effect could be due to increased uncertainty about the target price resulting in a more cautious attitude towards initiating a price change. Overall, the reported results give emphasis to conducting further research on the dynamic consequences of lumpy and heterogeneous pricing behavior, with a special emphasis on the sluggishness in price setters' response to idiosyncratic disturbances.

References

Amano, Robert A. and R. Tiff Macklem (1997): "Menu Costs, Relative Prices, and Inflation: Evidence for Canada," *Bank of Canada Working Paper 97-14*

Atkeson, Andrew and Lee Ohanian (2001): "Are Phillips Curves Useful for Forecasting Inflation?," *Federal Reserve Bank of Minneapolis Quarterly Review*, Spring, pp. 2-11

Balke, Nathan S. and Mark A. Wynne (2000): "An Equilibrium Analysis of Relative Price Changes and Aggregate Inflation," *Journal of Monetary Economics*, pp. 269-292

Ball, Laurence and N. Gregory Mankiw (1994): "Asymmetric Price Adjustment and Economic Fluctuations," *Economic Journal*, pp. 247-261

Ball, Laurence and N. Gregory Mankiw (1995): "Relative Price Changes as Aggregate Supply Shocks," *Quarterly Journal of Economics*, pp. 161-193

Bils, Mark and Peter J. Klenow (2002): "Some Evidence on the Importance of Sticky Prices," *NBER Working Paper #9069*

Blejer, Mario I. (1983): "On the Anatomy of Inflation: The Variability of Relative Commodity Prices in Argentina," *Journal of Money, Credit and Banking*, pp. 469-482

Bryan, Michael F. and Stephen G. Cecchetti (1999): "Inflation and the Distribution of Price Changes," *Review of Economics and Statistics*, pp. 188-196

Caballero, Ricardo J., Eduardo M. R. A. Engel and John C. Haltiwanger (1995): "Plant Level Adjustment and Aggregate Dynamics," *Brookings Papers on Economic Activity*, pp. 1-39

Carroll, Chris D. and Wendy E. Dunn (1997): "Unemployment Expectations, Jumping (S,s) Triggers, and Households Balance Sheets," *NBER Macroeconomics Annual*, pp. 165-229

Danziger, Leif (1987): "Inflation, Fixed Cost of Price Adjustment, and Measurement of Relative-Price Variability: Theory and Evidence," *American Economic Review*, pp. 704-713

Fengler, Matthias R. and Joachim K. Winter (2001): "Price Variability and Price Dispersion in a Stable Monetary Environment: Evidence from German Retail Markets," *manuscript*

Gibbons, Jean Dickinson and Chakraborti, Subhabrata (1992): *Nonparametric Statistical Inference* (New York: Marcel Dekker Inc.)

Judson, Ruth A. and Ann L. Owen (1997): "Estimating Dynamic Panel Data Models: A Practical Guide for Macroeconomists," *Finance and Economics Discussion Series 1997-3*

Lach, Saul and Daniel Tsiddon (1992): "The Behavior of Prices and Inflation: An Empirical Analysis of Disaggregated Data," *Journal of Political Economy*, pp. 349-389

Mankiw, N. Gregory and Ricardo Reis (2002): "Sticky Information Versus Sticky Prices: A Proposal to Replace the New Keynesian Phillips Curve," *Quarterly Journal of Economics*, pp. 1295-1328

Mills, Frederick (1927): *The Behavior of Prices* (New York: NBER, Inc.)
Parsley, David C. (1996): "Inflation and Relative Price Variability in the Short and Long Run: New Evidence from the United States," *Journal of Money Credit and Banking*, pp. 323-341
Perron, Pierre (1997): "Further Evidence on Breaking Trend Functions in Macroeconomic Variables," *Journal of Econometrics*, pp. 355-385
Rátfai, Attila (2003): "Linking Individual and Aggregate Price Changes," *working paper*
Rátfai, Attila (2004): "The Frequency and Size of Price Adjustment: Microeconomic Evidence," *Managerial and Decision Economics*, forthcoming
Reinsdorf, Marshall (1994): "New Evidence on the Relation between Inflation and Price Dispersion," *American Economic Review*, pp. 720-731
Runkle, David E. (1987): "Vector Autoregression and Reality," *Journal of Business and Economic Statistics*, pp. 437-442
Silver, Mick and Christos Ioannidis (2001): "Intercountry Differences in the Relationship between Relative Price Variability and Average Prices," *Journal of Political Economy*, pp. 355-374
Suvanto, Antti and Juhana Hukkinen (2002): "Stable Price Level and Changing Prices," *working paper*
Tommasi, Mariano (1993): "Inflation and Relative Prices: Evidence from Argentina," in Eytan Sheshinski and Yoram Weiss, eds., *Optimal Pricing, Inflation, and the Cost of Price Adjustment* (Cambridge, MA: MIT Press)
Tsiddon, Daniel (1993): "The (Mis)Behavior of the Aggregate Price Level," *Review of Economic Studies*, pp. 889-902
Vining, Daniel R., Jr. and Thomas C. Elwertowski (1976): "The Relationship between Relative Prices and the General Price Level," *American Economic Review*, pp. 699-708
Weiss, Yoram (1993): "Inflation and Price Adjustment: A Survey of Findings from Micro-Data," in Eytan Sheshinski and Yoram Weiss, eds., *Optimal Pricing, Inflation, and the Cost of Price Adjustment* (Cambridge, MA: MIT Press)
Wolman, Alexander L. (2000): "The Frequency and Costs of Individual Price Adjustment," *Federal Reserve Bank of Richmond, Economic Quarterly*, pp. 1-22
Woodford, Michael (2001): "Imperfect Common Knowledge and the Effects of Monetary Policy", *working paper*

APPENDIX

To evaluate the robustness of the SVAR results, alternative definitions of the relative price and measures of the asymmetry in the relative price distribution are examined. For simplicity, I focus on only results with pooled relative prices and aggregate inflation.

Alternative Measure of Asymmetry

A potential problem with the standard skewness statistic designed to represent the degree of bunching of relative prices in the tails of the distribution is that it could be sensitive to outliers in the distribution. To evaluate if the main results of the SVAR analysis are robust to the precise definition of asymmetry in the pooled relative price distribution, an alternative non-parametric measure is examined, the difference between the mean and the median scaled by the standard deviation, *mm*.[28] The statistic is expected to be larger, the more intensive the bunching of relative prices in the lower tail of the distributions is.

Using the alternative asymmetry measure, the VAR model of inflation and relative price asymmetry is estimated subject to the same identification restrictions as the ones used earlier. The top panel of Table AIV shows the decompositions of 12-month-ahead forecast error variances. The figures corroborate the baseline results in that idiosyncratic shocks are quantitatively important determinants of aggregate inflation dynamics. The impulse response functions are depicted in Figures AIa and AIb. A direct comparison of these graphs to Figures IIIa and IIIb reveals that the impulse response functions obtained for the skewness statistic and the alternative asymmetry measure are strikingly similar to each other.

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[28] This alternative asymmetry variable is positively correlated with the standard skewness statistic in the data, with a partial correlation coefficient of 0.58. I experimented with yet another alternative measure, $W = (Q1 + Q3 - 2M)/(Q3-Q1)$, where $Q1$ and $Q3$ are the lower and the upper quartiles and M is the median of the distribution. As W and the scaled mean-median difference measures are virtually identical, I confine my attention to the latter one.

Alternative Timing in the Measurement of Relative Prices

Another potential objection to the baseline results is that they are contingent on a particular timing convention in the definition of the target price. To address this issue, the relative price is alternatively defined as $z_{ijt} = p_{ij,t-1} - p_{j,t-1}$, where $p_{j,t-1}$ is the log average product-specific price lagged by one period.

As in the previous subsection, the impulse response functions and forecast error variance decompositions are examined solely for the pooled relative price measure and aggregate inflation. Figures AIIa and AIIb display the impulse responses functions. Under the *LR* scheme the impulse responses are indistinguishable from the ones obtained in the baseline case. Aggregate inflation responds to idiosyncratic shocks with a five months lag following the shock in a statistically significant way. Impulse responses under the *SR* assumption slightly differ from the *LR* case, the lagged response of inflation to idiosyncratic shocks materializes two months after the initial disturbance, and there is a significant direct impact as well. The forecast error variances displayed in the bottom panel of Table AIV again confirm that idiosyncratic shocks are in general important determinants of inflation dynamics.

TABLE I
PRODUCTS IN THE SAMPLE

Product Code	Product Name	Absolute Weight	Relative Weight	Mean	Standard Deviation
10001	Pork, Chops	0.49	9.39	1.22	5.09
10002	Spare Ribs, with Bone	0.19	3.64	1.55	5.21
10003	Pork, Leg without Bone and Hoof	0.77	14.75	1.18	5.36
10102	Beef, Round	0.04	0.77	1.83	2.38
10103	Beef, Shoulder with Bone	0.04	0.77	1.97	2.46
10301	Pork Liver	0.12	2.30	1.75	2.86
10401	Chicken Ready to Cook	0.41	7.85	1.43	1.54
10601	Sausage, Bologna Type	0.25	4.79	1.54	2.96
10603	Sausage, Italian Type	0.17	3.25	1.71	2.88
10605	Sausage, Boiling	0.17	3.26	1.80	2.91
10801	Carp, Living	0.06	1.15	1.93	1.87
11302	Curd, 250g	0.16	3.07	2.33	3.67
12101	Lard, Pork	0.13	2.49	2.10	8.06
12201	Fat Bacon	0.07	1.34	2.81	3.85
12203	Smoked Boiled Bacon	0.07	1.34	2.62	3.90
12301	Sunflower Oil	0.37	7.09	1.83	3.82
13002	Flour, Prime Quality	0.28	5.36	1.98	3.35
13301	Roll, 52-56g, 10 pieces	0.21	4.02	2.17	3.50
13501	Sugar, White, Granulated	0.53	10.15	1.59	2.55
13801	Dry Biscuits, without Butter, Packed	0.05	0.96	1.63	1.86
14424	Tomato Paste	0.03	0.57	1.41	1.91
15208	Vinegar, 10 hydrate	0.05	0.96	1.19	2.99
17001	Coffee, Omnia Type, 100g	0.21	4.02	1.87	4.02
19001	Cigarette, Kossuth Type, 25 pieces	0.17	3.26	1.72	2.90
52366	Broom, Horsehair-Synthetic Mix	0.01	0.19	1.38	1.65
66105	Car Driving School, Full Course	0.16	3.07	2.58	7.85
66301	Movie Ticket, Evening, 1-6 Rows	0.01	0.19	2.19	1.64
		5.22	100.00	1.82	3.44
	All Products in CPI	100.00	100.00	1.76	1.17

Notes: 1 Figures are compiled from various consumer price statistic booklets of the Central Statistical Office, Hungary.
2 Products are narrowly defined according to size, branding, type and flavor.
3 Weights are expenditure-based. *Absolute Weight* is taken from the 1995 CPI. *Relative Weight* reflects weight of the item in this particular sample.
4 *Mean* is the average of monthly inflation rates. *Standard Deviation* is the standard deviation of monthly inflation rates.

Table II

SHORT-RUN AND LONG-RUN MULTIPLIERS

POOLED DATA ESTIMATES

	Short Run		Long Run	
Identification	$G^0{}_{\Pi S}$	$G^0{}_{S\Pi}$	$M(1)_{\Pi S}$	$M(1)_{S\Pi}$
SR: $B^0{}_{S\Pi} = 0$	-2.38	0	2.15	0.08
	(1.52)	(0)	(1.23)	(0.19)
LR: $M(1)_{\Pi S} = 0$	-1.25	1.01	0	0.27
	(0.53)	(1.27)	(0)	(0.07)

Note: *SR* and *LR* refer to the identification scheme. Standard errors are in parentheses.

PRODUCT-SPECIFIC ESTIMATES

	Short Run		Long Run	
Identification	$G^0{}_{\Pi S}$	$G^0{}_{S\Pi}$	$M(1)_{\Pi S}$	$M(1)_{S\Pi}$
SR: $B^0{}_{S\Pi} = 0$	0.74	0	-0.16	-0.16
	[1.45]	[0]	[2.81]	[1.08]
LR: $M(1)_{\Pi S} = 0$	0.18	-0.01	0	-0.14
	[3.03]	[0.07]	[0]	[1.49]

Note: *SR* and *LR* refer to the identification restriction imposed. The figures are the cross-product median of the estimated parameters. The corresponding standard deviations are in square brackets.

PANEL ESTIMATES

	Short Run		Long Run	
Identification	$G^0{}_{\Pi S}$	$G^0{}_{S\Pi}$	$M(1)_{\Pi S}$	$M(1)_{S\Pi}$
SR: $B^0{}_{S\Pi} = 0$	0.14	0	0.33	-0.25
	[0.56]	[0]	[0.78]	[0.14]
LR: $M(1)_{\Pi S} = 0$	-0.18	0.03	0	-0.17
	[0.06]	[0.08]	[0]	[0.23]

Note: *SR* and *LR* refer to the identification restriction imposed. The figures are the cross-product median of the estimated parameters. The corresponding standard deviations are in square brackets.

Table III

FORECAST ERROR VARIANCE DECOMPOSITION

POOLED DATA ESTIMATES

		Variance Share in Percentage Terms 12 month horizon	
Identification	Source of Shocks	Inflation	Relative Price Asymmetry
SR: $B^0_{S\Pi} = 0$	Aggregate (Π)	0.66	0.10
	Idiosyncratic (S)	0.34	0.90
LR: $M(1)_{\Pi S} = 0$	Aggregate (Π)	0.36	0.58
	Idiosyncratic (S)	0.64	0.42

Note: *SR* and *LR* refer to the identification restriction imposed. Asymmetry is measured by skewness.

PRODUCT-SPECIFIC ESTIMATES

		Variance Share in Percentage Terms 12 month horizon	
Identification	Source of Shocks	Inflation	Relative Price Asymmetry
SR: $B^0_{S\Pi} = 0$	Aggregate (Π)	0.81	0.19
		[0.13]	[0.23]
	Idiosyncratic (S)	0.19	0.82
		[0.13]	[0.23]
LR: $M(1)_{\Pi S} = 0$	Aggregate (Π)	0.74	0.34
		[0.22]	[0.22]
	Idiosyncratic (S)	0.26	0.66
		[0.22]	[0.22]

Note: *SR* and *LR* refer to the identification restriction imposed. The figures are the cross-product median of the estimated parameters. The corresponding standard deviations are in square brackets. Asymmetry is measured by skewness.

PANEL ESTIMATES

		Variance Share in Percentage Terms 12 month horizon	
Identification	Source of Shocks	Inflation	Relative Price Asymmetry
SR: $B^0_{S\Pi} = 0$	Aggregate (Π)	0.73	0.19
		[0.04]	[0.05]
	Idiosyncratic (S)	0.27	0.82
		[0.03]	[0.05]
LR: $M(1)_{\Pi S} = 0$	Aggregate (Π)	0.74	0.24
		[0.04]	[0.02]
	Idiosyncratic (S)	0.26	0.76
		[0.05]	[0.03]

Note: *SR* and *LR* refer to the identification restriction imposed. The figures are the cross-product median of the estimated parameters. The corresponding standard deviations are in square brackets. Asymmetry is measured by skewness.

Table AI

UNIT ROOT TESTS FOR INFLATION AND RELATIVE PRICE SKEWNESS

Inflation			Skewness		
Product Code	ADF t-statistic	Largest AR Root	Product Code	ADF t-statistic	Largest AR Root
dp10001	-4.94	0.45	s10001	-2.61	0.74
dp10002	-4.95	0.44	s10002	-3.77	0.56
dp10003	-4.88	0.46	s10003	-2.71	0.73
dp10102	-3.92	0.53	s10102	-4.32	0.54
dp10103	-4.02	0.44	s10103	-8.83	0.70
dp10301	-4.06	0.50	s10301[b]	-8.91	-0.87
dp10401	-5.83	0.21	s10401	-4.26	0.47
dp10601	-4.50	0.43	s10601	-3.86	0.55
dp10603	-4.50	0.43	s10603[a]	-3.22	0.69
dp10605	-4.24	0.48	s10605	-2.67	0.63
dp10801	-4.19	0.46	s10801	-5.16	0.23
dp11302	-7.21	0.07	s11302	-4.71	0.39
dp12101	-3.88	0.64	s12101	-3.67	0.59
dp12201	-3.98	0.52	s12201	-3.70	0.57
dp12203	-4.78	0.39	s12203	-3.03	0.69
dp12301	-5.77	0.21	s12301	-3.10	0.66
dp13002	-4.14	0.48	s13002	-3.96	0.50
dp13301	-6.37	0.11	s13301	-3.88	0.55
dp13501	-4.46	0.35	s13501	-5.91	0.19
dp13801	-5.95	0.18	s13801	-3.59	0.60
dp14424	-4.48	0.42	s14424[c]	-5.02	0.38
dp15208	-5.40	0.27	s15208	-2.81	0.75
dp17001	-3.46	0.63	s17001	-4.17	0.68
dp19001	-7.03	0.02	s19001	-2.68	0.78
dp52366	-8.20	0.12	s52366[a]	-3.71	0.56
dp66105	-6.66	0.07	s66105[d]	-5.58	0.43
dp66301	-7.06	0.02	s66301	-2.70	0.75

[a] ADF regression includes deterministic time trend.
[b] ADF regression includes dummies for a structural "intercept and slope" break at 94:12. The 5% t-sig critical value is –5.59 for T = 70. See Perron [1997].
[c] ADF regression includes dummies for a structural "intercept break" at 93:01. The 5% t-sig critical value is –4.83 for T = 100. See Perron [1997].
[d] ADF regression includes dummies for a structural "slope break" at 93:01. The 5% t-sig critical value is –5.23 for T = 60. See Perron [1997].

Notes: 1. *dp*<code> refers to the monthly percentage change in the average price level of the product denoted by <code>. Similarly, *s*<code> refers to the relative price skewness measure of the product denoted by <code>.
2. The number of lags in the regressions is based on the Schwartz Information Criterion allowing for a maximum number of lags of 12.
3. Unless otherwise indicated, regressions do not include a time trend.

Table AII
UNIT ROOT TESTS FOR LOG PRICE LEVELS

Product Code	ADF *t*-statistic	Largest AR Root	ADF *t*-statistic	Largest AR Root
	Deterministic time trend included		No deterministic time trend	
log_p10001	-3.85	0.81	-1.85	0.96
log_p10002	-2.92	0.82	-1.47	0.97
log_p10003	-3.82	0.82	-1.81	0.96
log_p10102	-1.42	0.94	-1.24	0.99
log_p10103	-1.34	0.94	-1.16	0.99
log_p10301	-2.93	0.86	-1.44	0.98
log_p10401	-1.68	0.90	0.03	1.00
log_p10601	-1.62	0.91	-1.97	0.97
log_p10603	-1.92	0.89	-1.78	0.98
log_p10605	-1.62	0.92	-1.86	0.98
log_p10801	-2.76	0.84	-0.00	1.00
log_p11302	-6.22	0.52	-0.57	0.98
log_p12101	-3.82	0.83	-2.45	0.96
log_p12201	-2.79	0.85	-1.64	0.98
log_p12203	-3.39	0.82	-1.06	0.99
log_p12301	-2.48	0.84	0.31	0.99
log_p13002	-1.74	0.94	-0.79	0.99
log_p13301	-3.40	0.79	-2.25	0.96
log_p13501	-1.96	0.94	-1.65	0.98
log_p13801	-2.19	0.86	-1.26	0.99
log_p14424	-0.63	0.98	-2.41	0.98
log_p15208	-2.11	0.88	-1.61	0.95
log_p17001	-2.19	0.93	-0.71	0.99
log_p19001	-2.73	0.78	0.39	1.01
log_p52366	-2.15	0.88	1.72	1.02
log_p66105	-1.81	0.89	-1.51	0.96
log_p66301	-0.97	0.94	-1.18	0.99

Notes: 1. *log_p*<code> refers to the log of the average price level of the product denoted by <code>.
2. The number of lags in the regressions is based on the Schwartz Information Criterion with a maximum number of lags of 12.

Table AIII
UNIT ROOT TESTS FOR RESIDUALS FROM SEASONAL DUMMIES REGRESSIONS

Residuals from Seasonal Dummy Regressions		
Product Code	ADF t-statistic	Largest AR Root
res_dp10001	-3.95	0.54
res_dp10002	-4.30	0.48
res_dp10003	-3.98	0.53
res_dp10102	-3.79	0.46
res_dp10103	-4.19	0.49
res_dp10301	-4.25	0.48
res_dp10401	-5.30	0.28
res_dp10601	-4.70	0.40
res_dp10603	-4.95	0.35
res_dp10605	-4.50	0.43
res_dp10801	-4.58	0.40
res_dp11302	-9.51	-0.13
res_dp12101	-3.67	0.59
res_dp12201	-3.57	0.52
res_dp12203	-4.51	0.43
res_dp12301	-2.90	-0.26
res_dp13002	-3.77	0.54
res_dp13301	-7.15	0.00
res_dp13501	-4.79	0.29
res_dp13801	-6.20	0.16
res_dp14424	-4.46	0.44
res_dp15208	-5.22	0.30
res_dp17001	-3.26	0.65
res_dp19001	-7.71	-0.07
res_dp52366	-8.44	-0.14
res_dp66105	-6.50	0.09
res_dp66301	-7.31	0.00

Notes: 1. *res_dp*<code> refers to the residual obtained from a seasonal dummy regression of the change in the log average price level of the product denoted by <code>.
2. ADF regressions include a constant and no time trend.
3. The number of lags in the regressions is based on the Schwartz Information Criterion with a maximum number of lags of 12.

Table AIV

FORECAST ERROR VARIANCE DECOMPOSITION - POOLED DATA

ALTERNATIVE MEASURE OF ASYMMETRY: SCLALED MEAN-MEDIAN DIFFERENCE

		Variance Share in Percentage Terms 12 month horizon	
Identification	Source of Shocks	Inflation	Relative Price Asymmetry
SR: $B^0_{S\Pi} = 0$	Aggregate (Π)	0.72	0.18
	Idiosyncratic (S)	0.28	0.82
LR: $M(1)_{\Pi S} = 0$	Aggregate (Π)	0.62	0.20
	Idiosyncratic (S)	0.38	0.80

Note: *SR* and *LR* refer to the particular identification scheme. Asymmetry is measured by the mean-median difference scaled by the standard deviation.

ALTERNATIVE MEASURE OF TIMING: LAGGED TARGET PRICE

		Variance Share in Percentage Terms 12 month horizon	
Identification	Source of Shocks	Inflation	Relative Price Asymmetry
SR: $B^0_{S\Pi} =0$	Aggregate (Π)	0.68	0.21
	Idiosyncratic (S)	0.32	0.79
LR: $M(1)_{\Pi S} = 0$	Aggregate (Π)	0.66	0.40
	Idiosyncratic (S)	0.34	0.60

Note: *SR* and *LR* refer to the particular identification scheme. The relative price is measured as the log difference between actual and average price, both lagged by one period. Asymmetry is measured by skewness.

Figure Ia

EVOLUTION OF CROSS-SECTIONAL DENSITY – AGGREGATE SHOCK

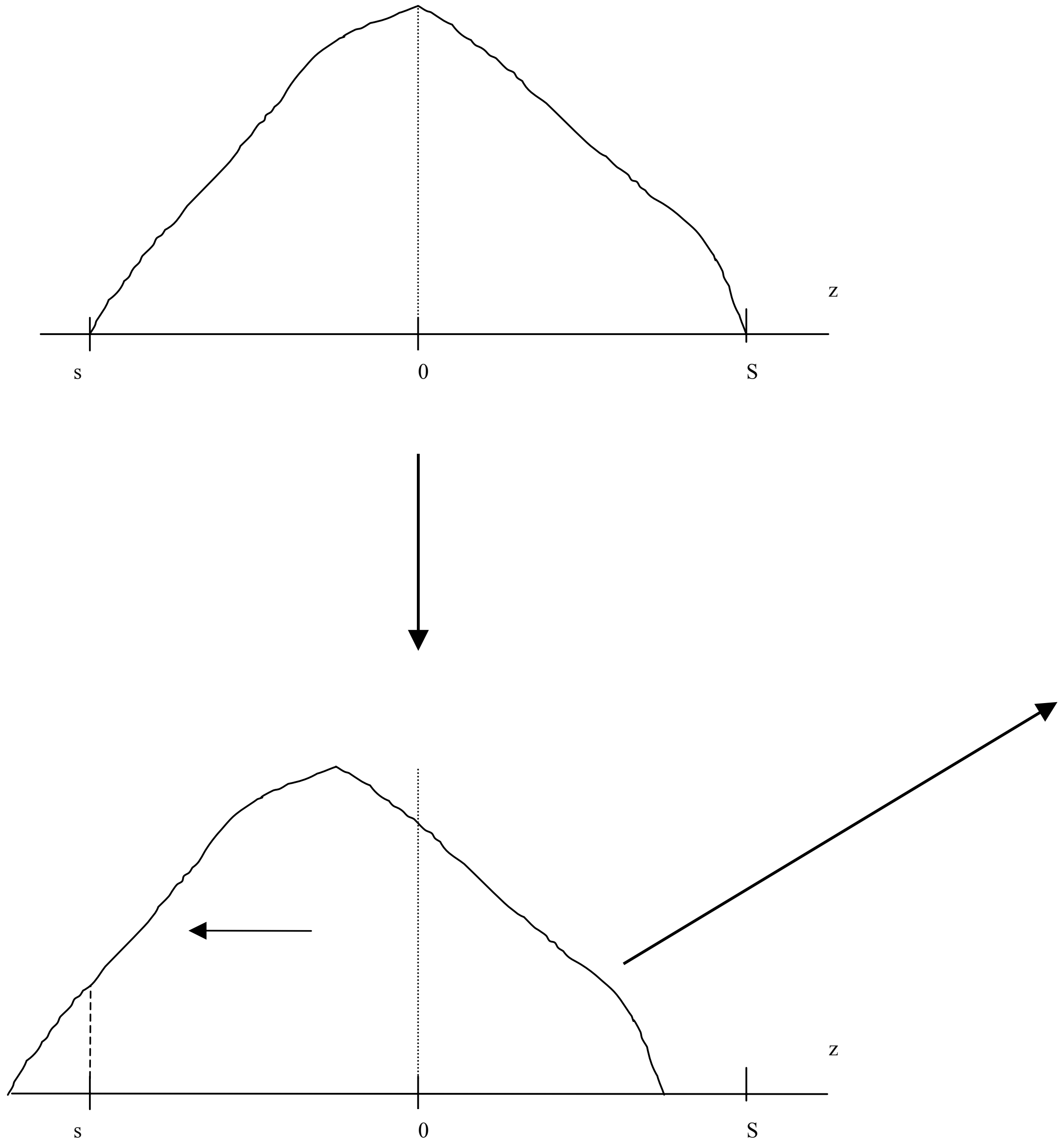


Figure Ib

EVOLUTION OF CROSS-SECTIONAL DENSITY – AGGREGATE SHOCK, CONT'D

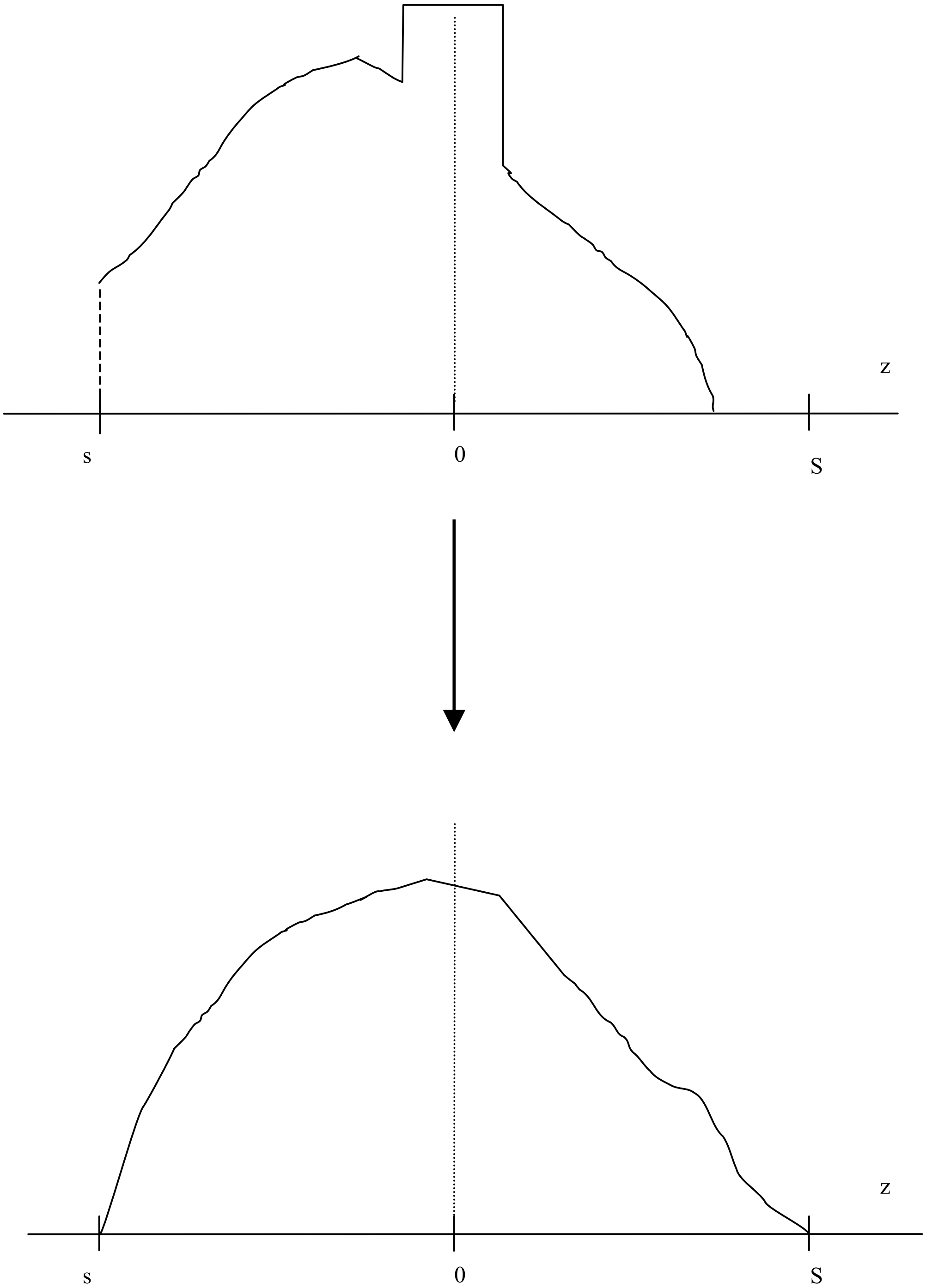


Figure Ic

EVOLUTION OF CROSS-SECTIONAL DENSITY – IDIOSYNCRATIC SHOCK

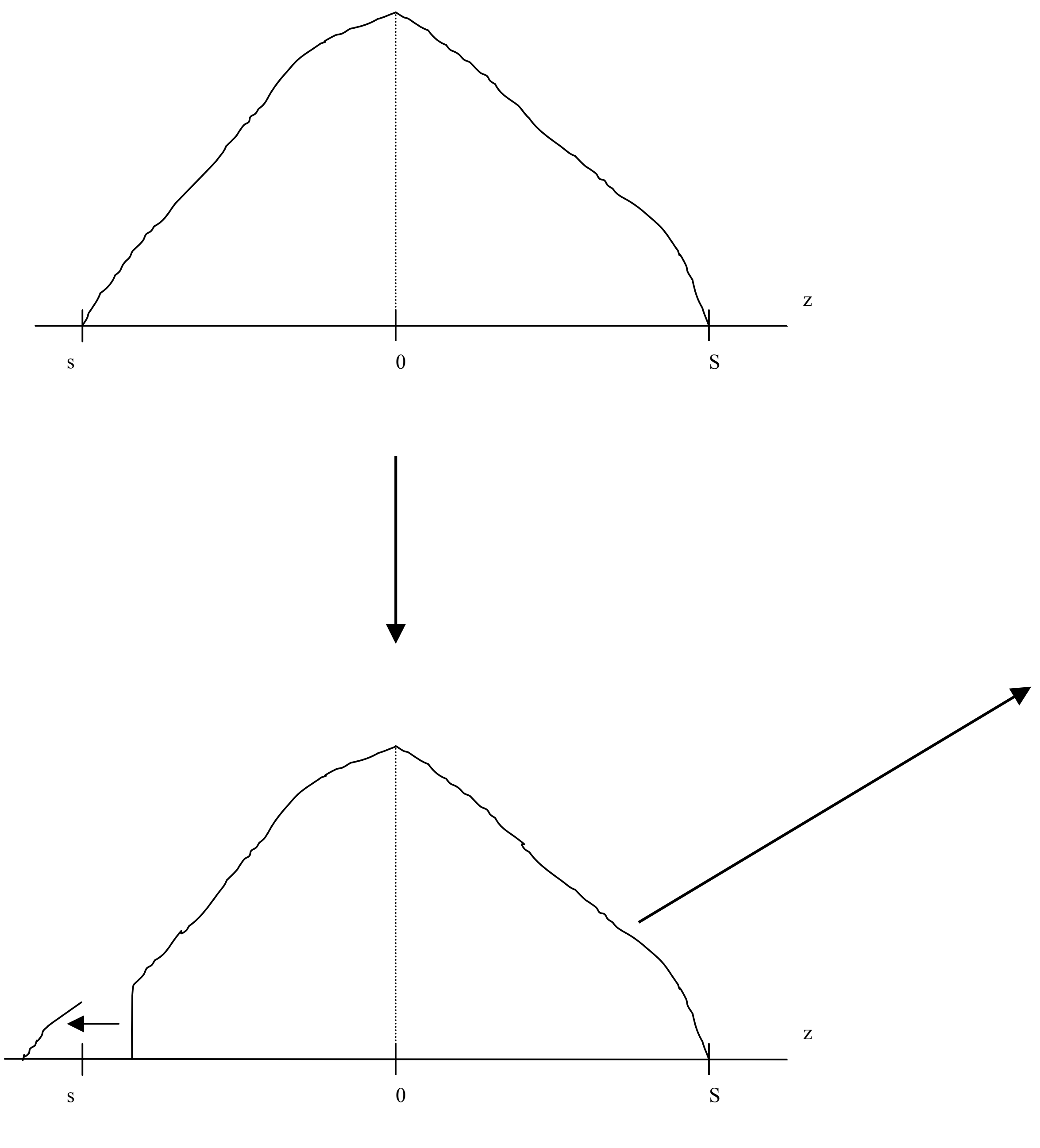


Figure Id

EVOLUTION OF CROSS-SECTIONAL DENSITY – IDIOSYNCRATIC SHOCK, CONT'D

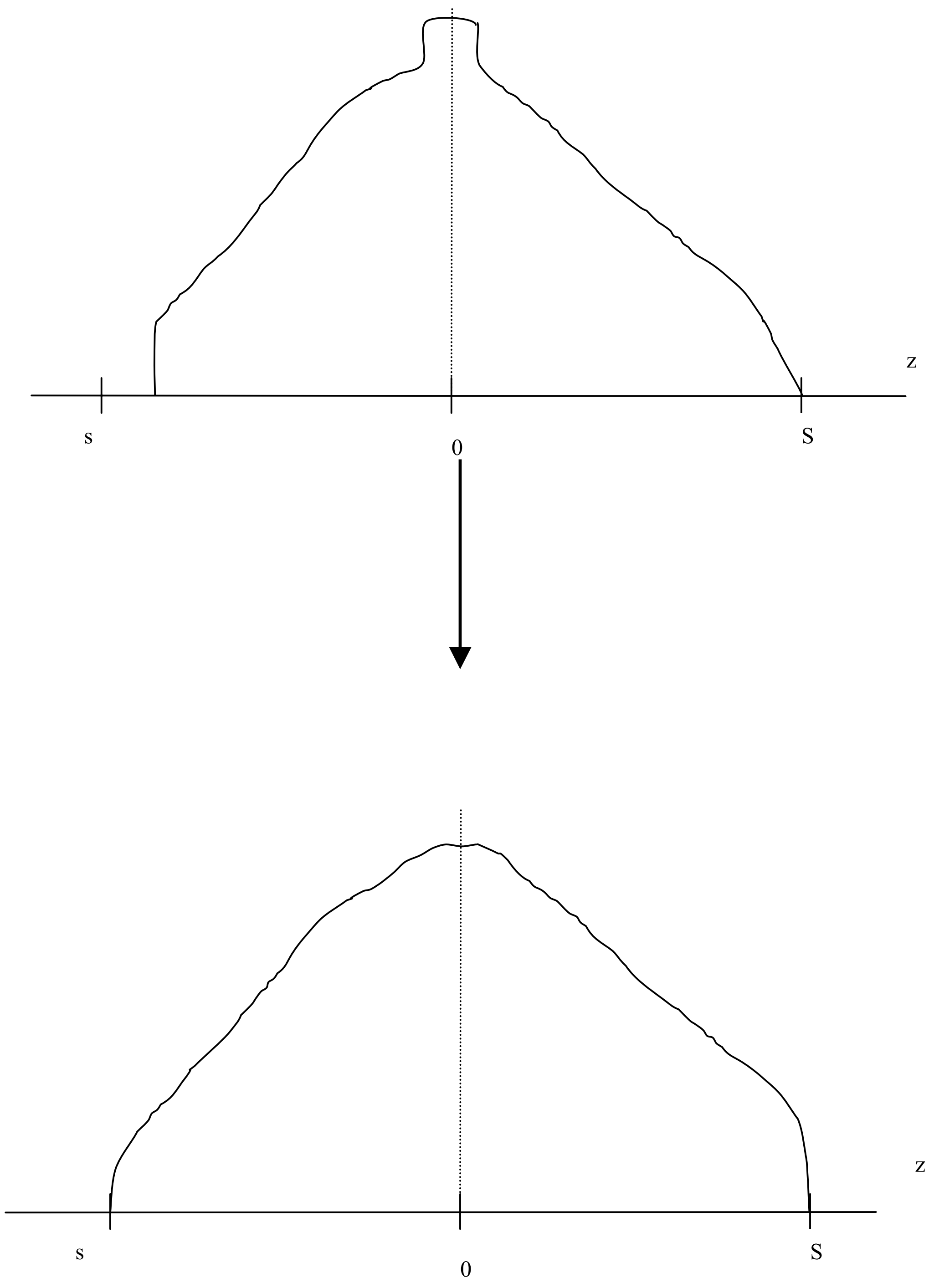


Figure II

ANNUAL CPI INFLATION IN HUNGARY

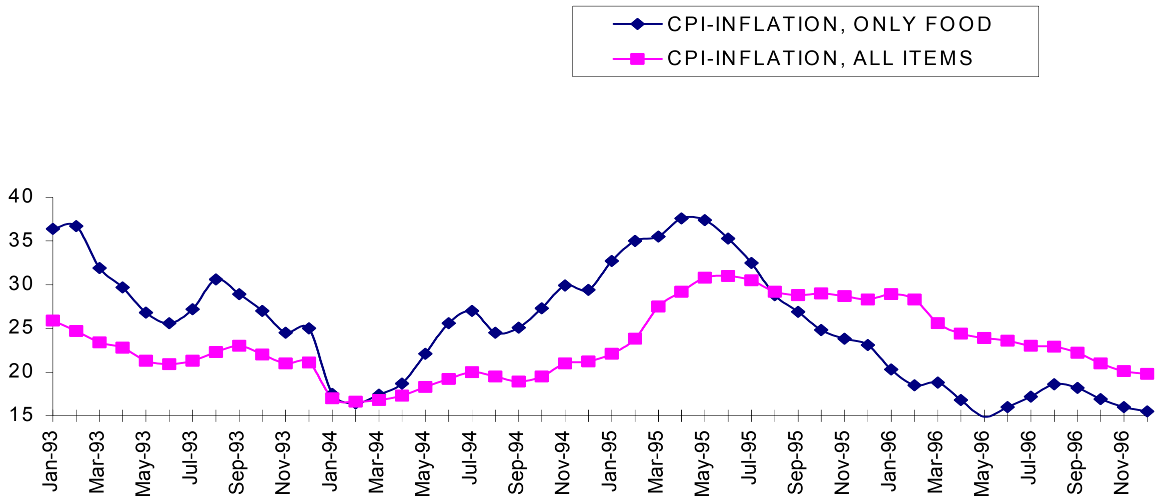
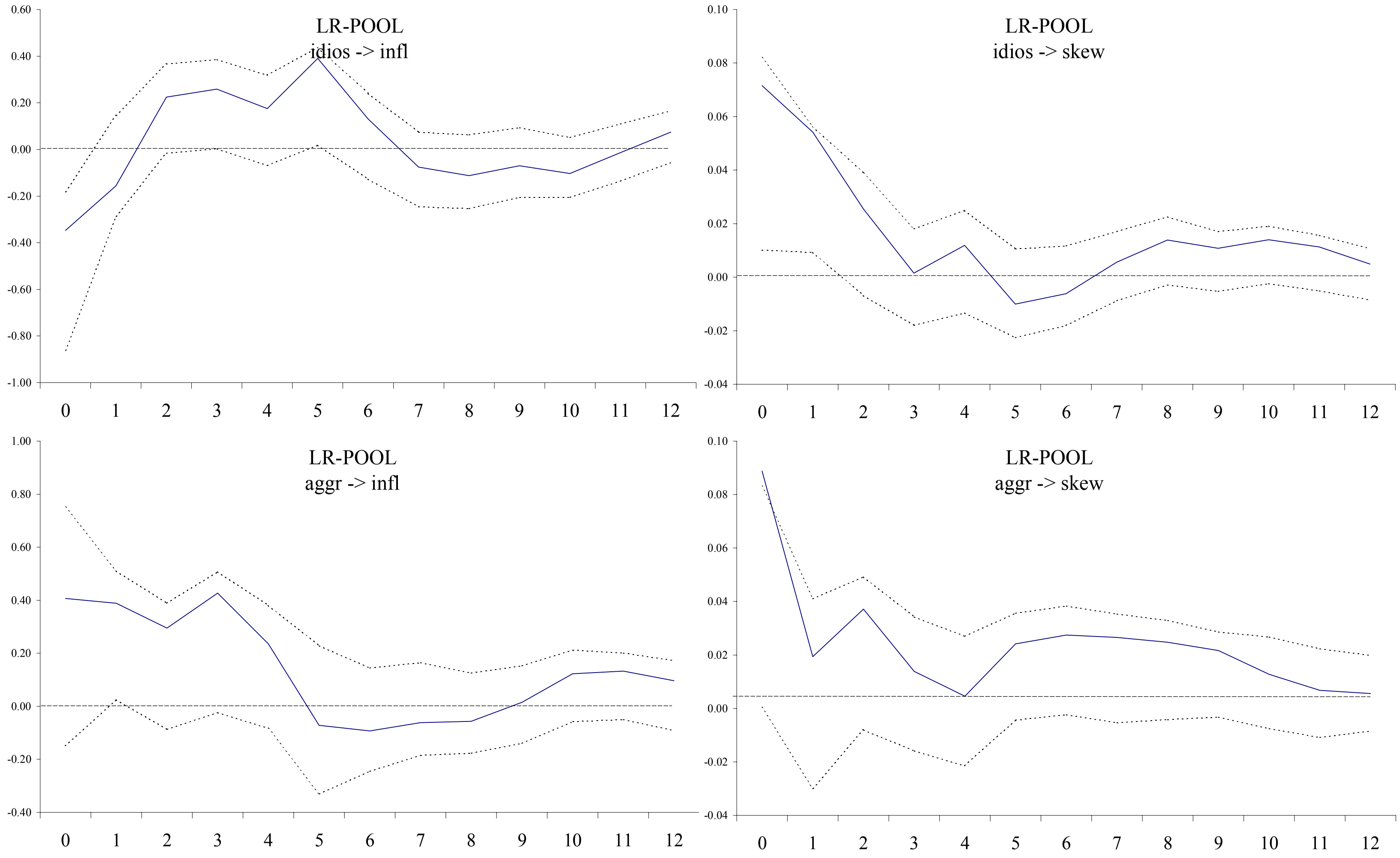
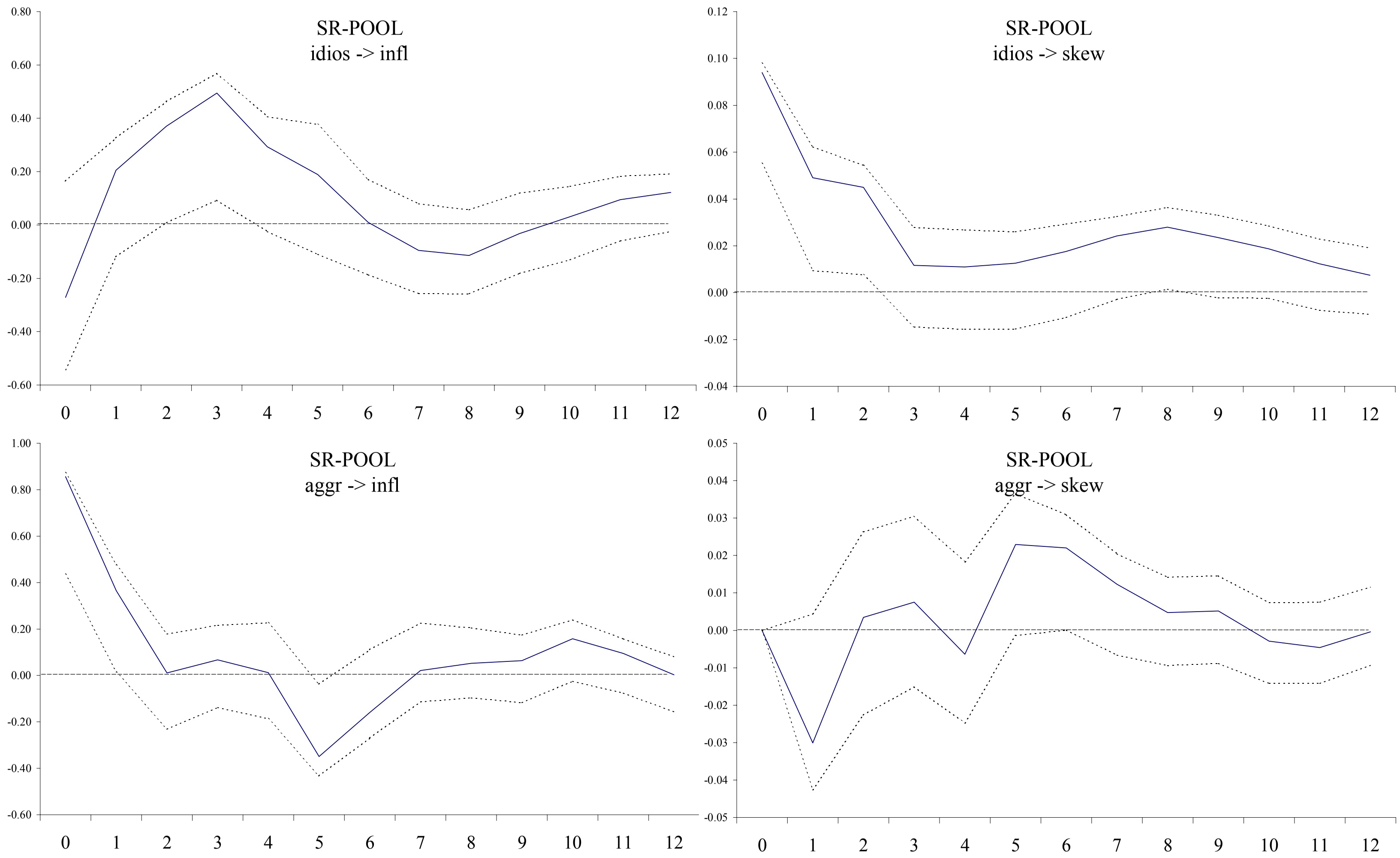


Figure IIIa
Impulse Response Functions



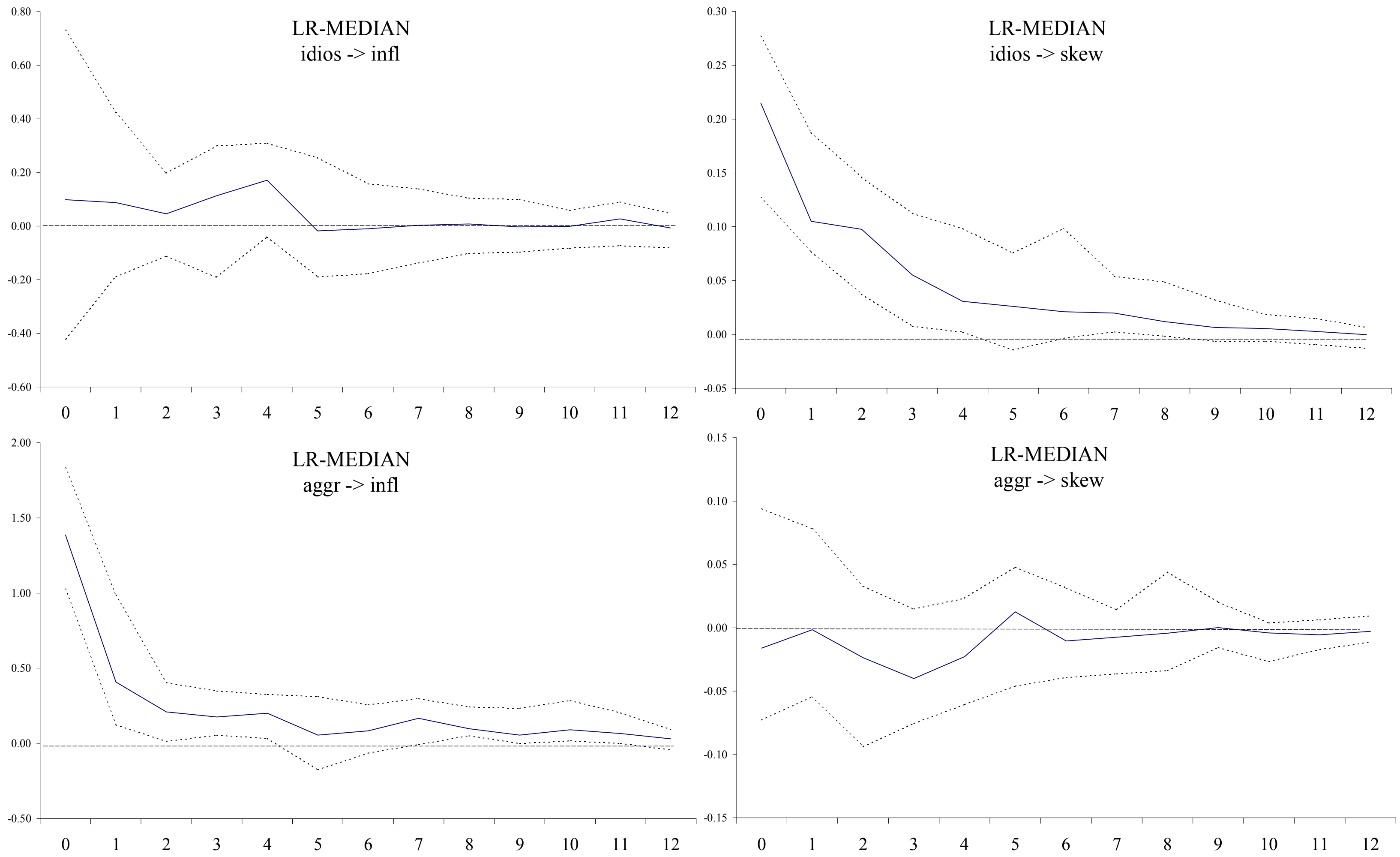
Note: Dashed lines are 90 percent Runkle (1987) confidence bands.
Horizontal axis: months following shock

Figure IIIb
Impulse Response Functions



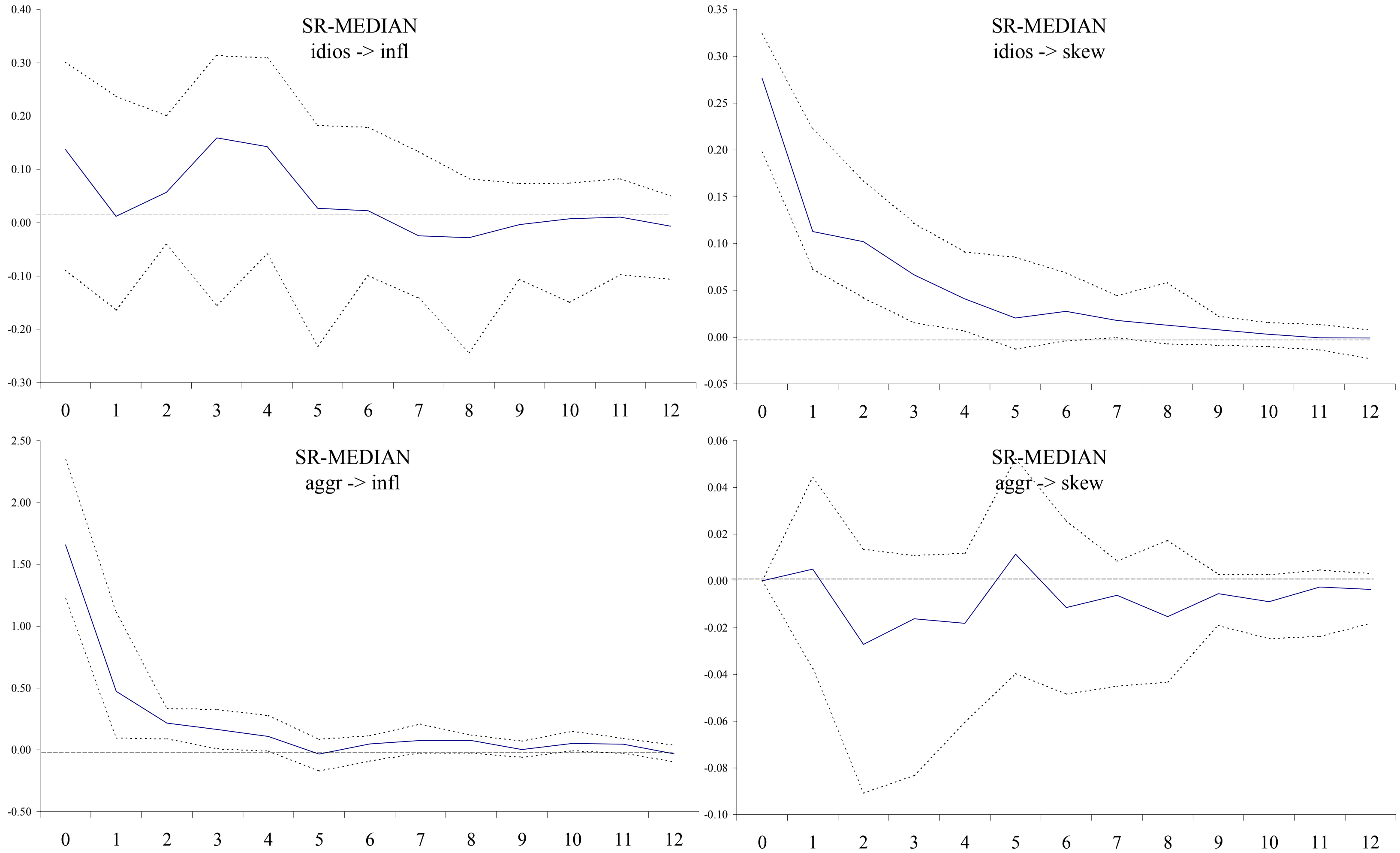
Note: Dashed lines are 90 percent Runkle (1987) confidence bands.
Horizontal axis: months following shock

Figure IVa
Impulse Response Functions



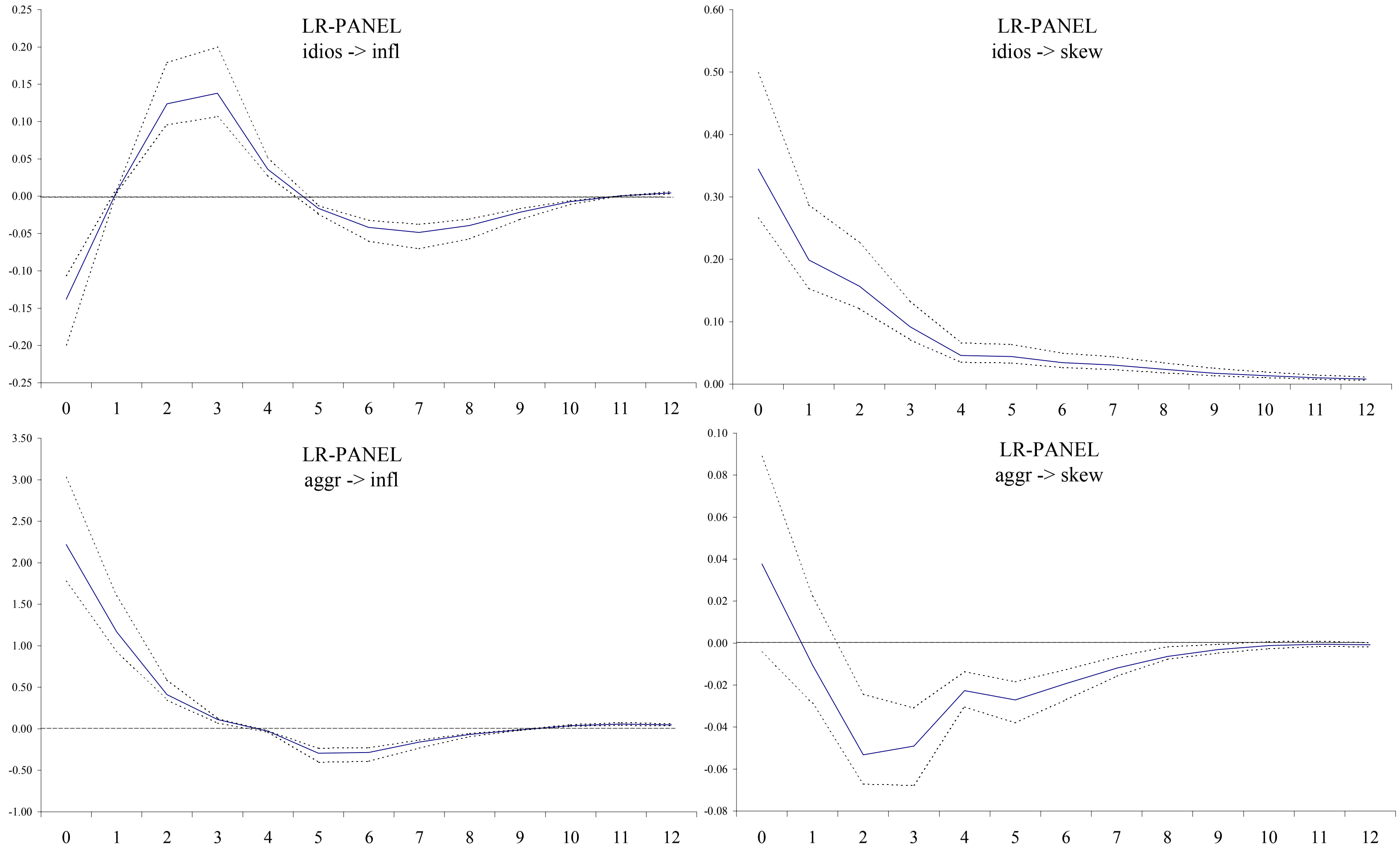
Note: Dashed lines are the upper and lower quartiles, the solid line is the median of impulse responses across products.
Horizontal axis: months following shock

Figure IVb
Impulse Response Functions



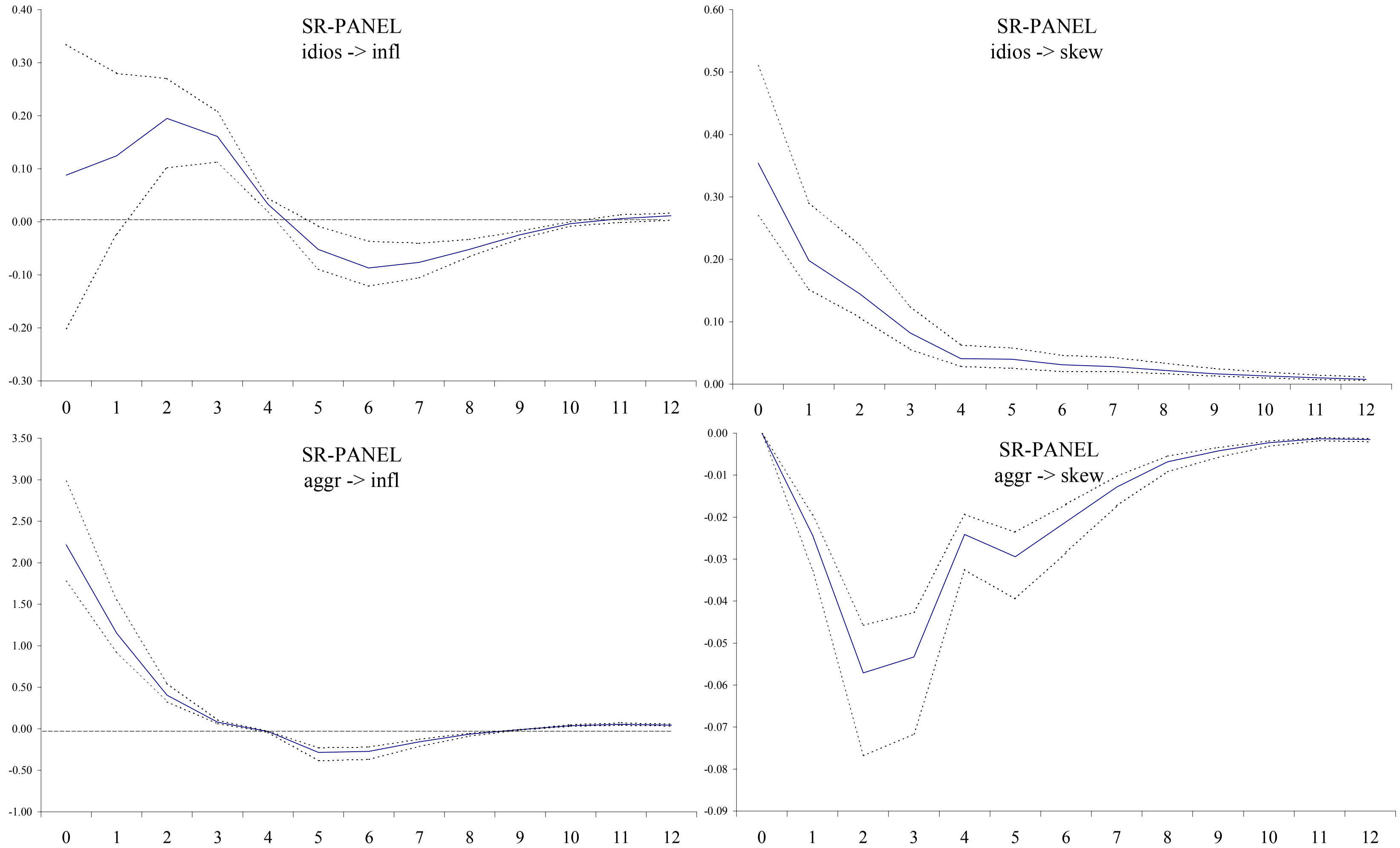
Note: Dashed lines are the upper and lower quartiles, the solid line is the median of impulse responses across products.
Horizontal axis: months following shock

Figure Va
Impulse Response Functions
Panel VAR



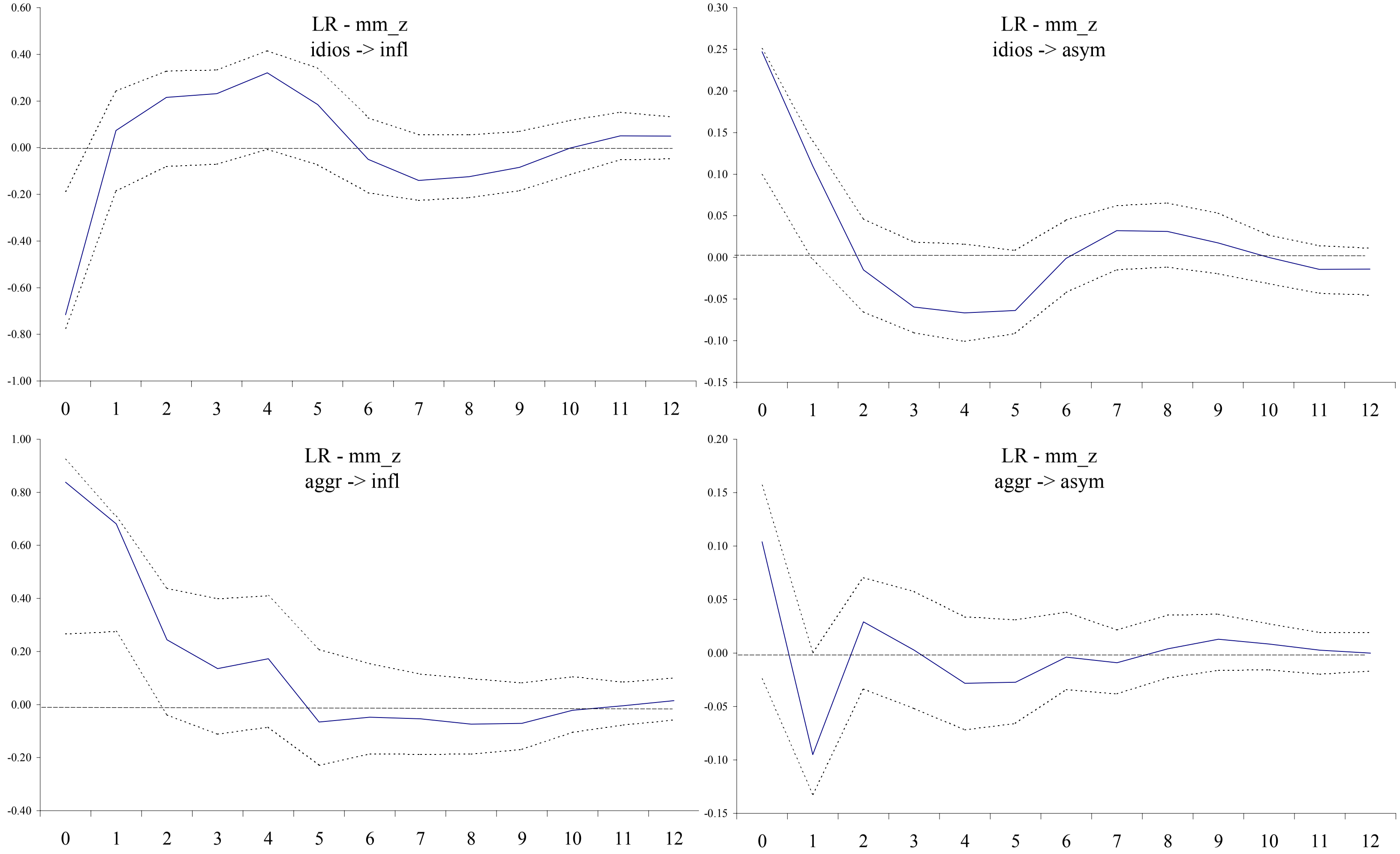
Note: Dashed lines are the upper and lower quartiles, the solid line is the median of impulse responses across products.
Horizontal axis: months following shock

Figure Vb
Impulse Response Functions
Panel VAR



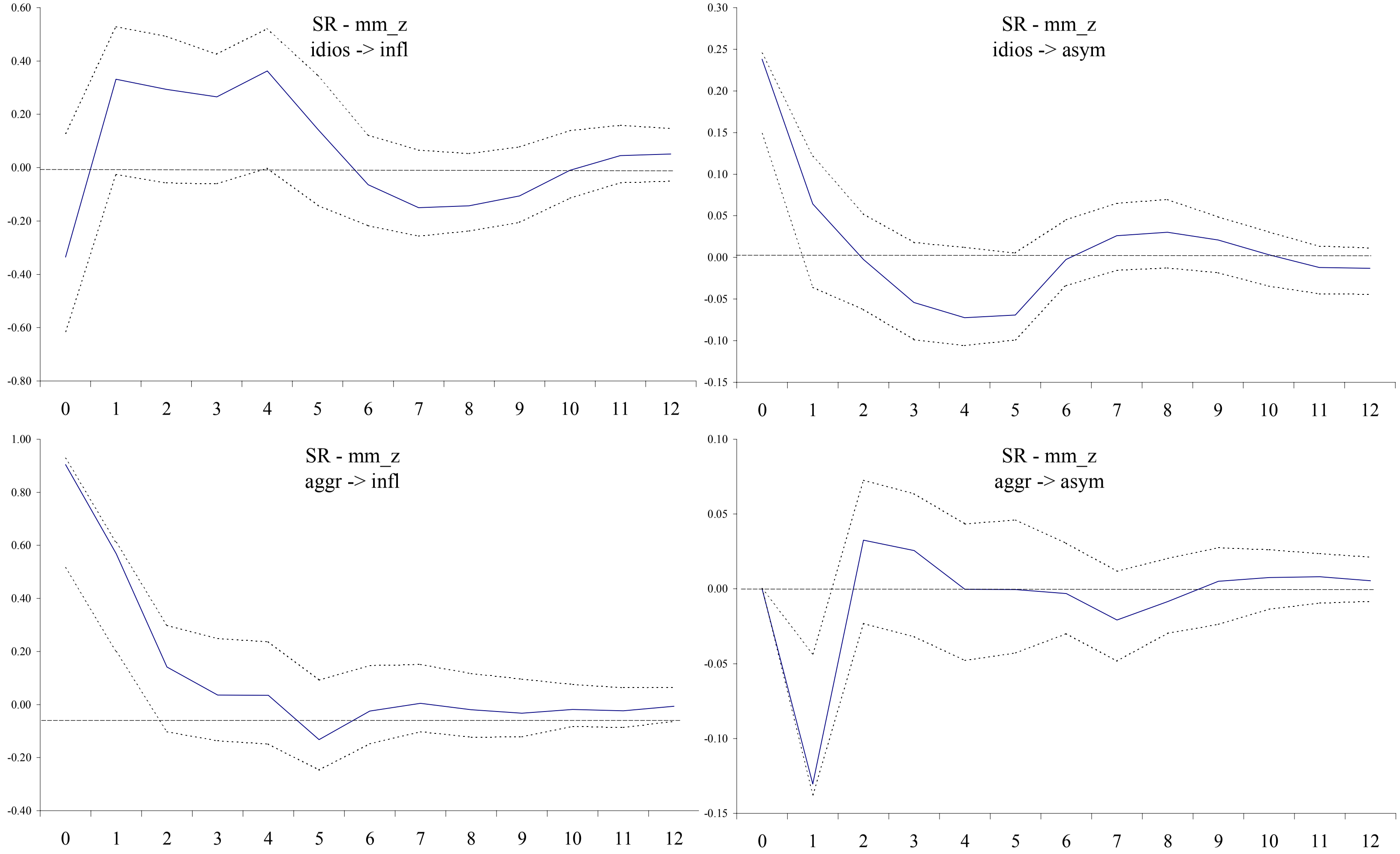
Note: Dashed lines are the upper and lower quartiles, the solid line is the median of impulse responses across products.
Horizontal axis: months following shock

Figure AIa
Impulse Response Functions
(Pooled Data)



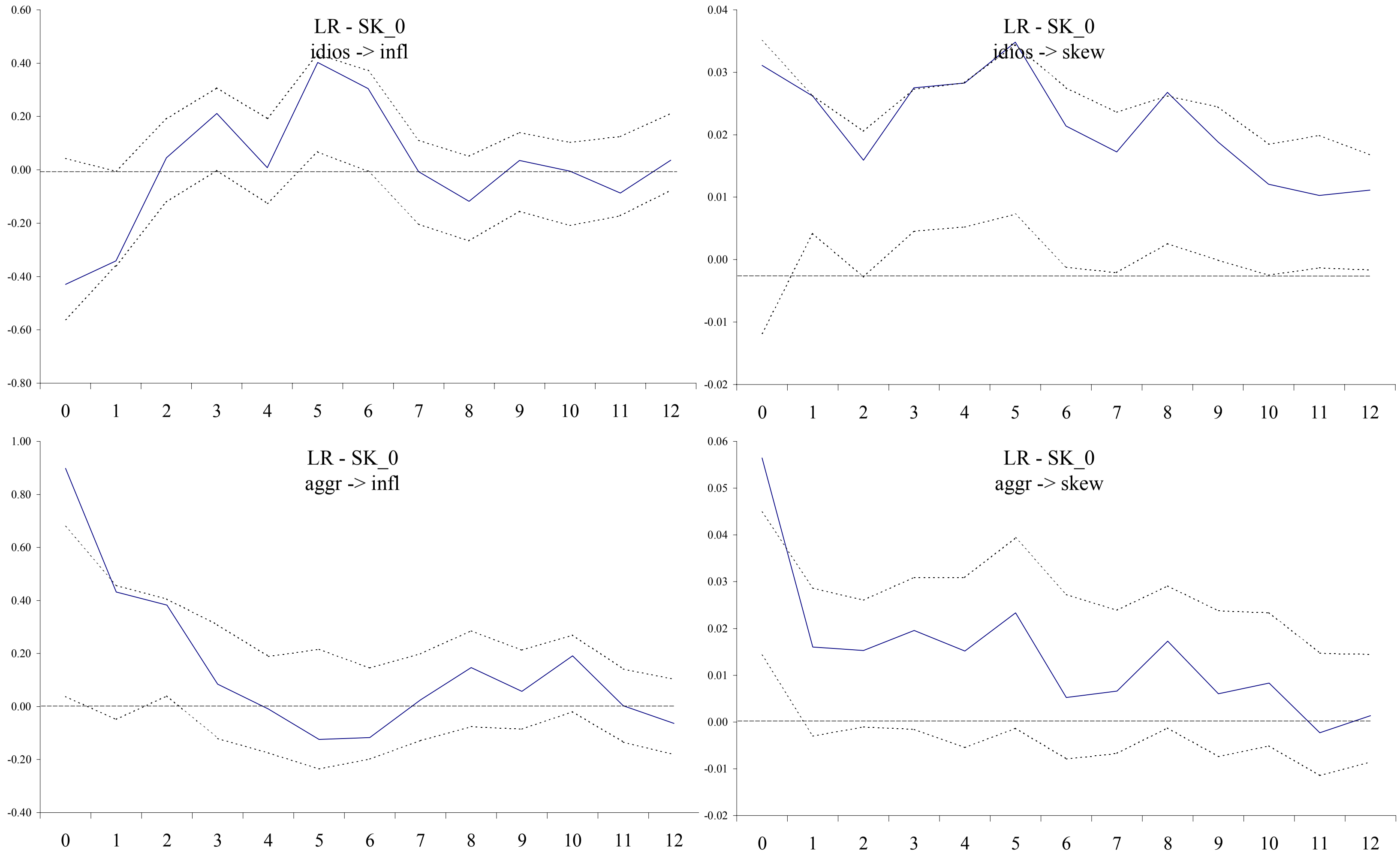
Note: Dashed lines are 90 percent Runkle (1987) confidence bands.
Horizontal axis: months following shock

Figure AIb
Impulse Response Functions
(Pooled Data)



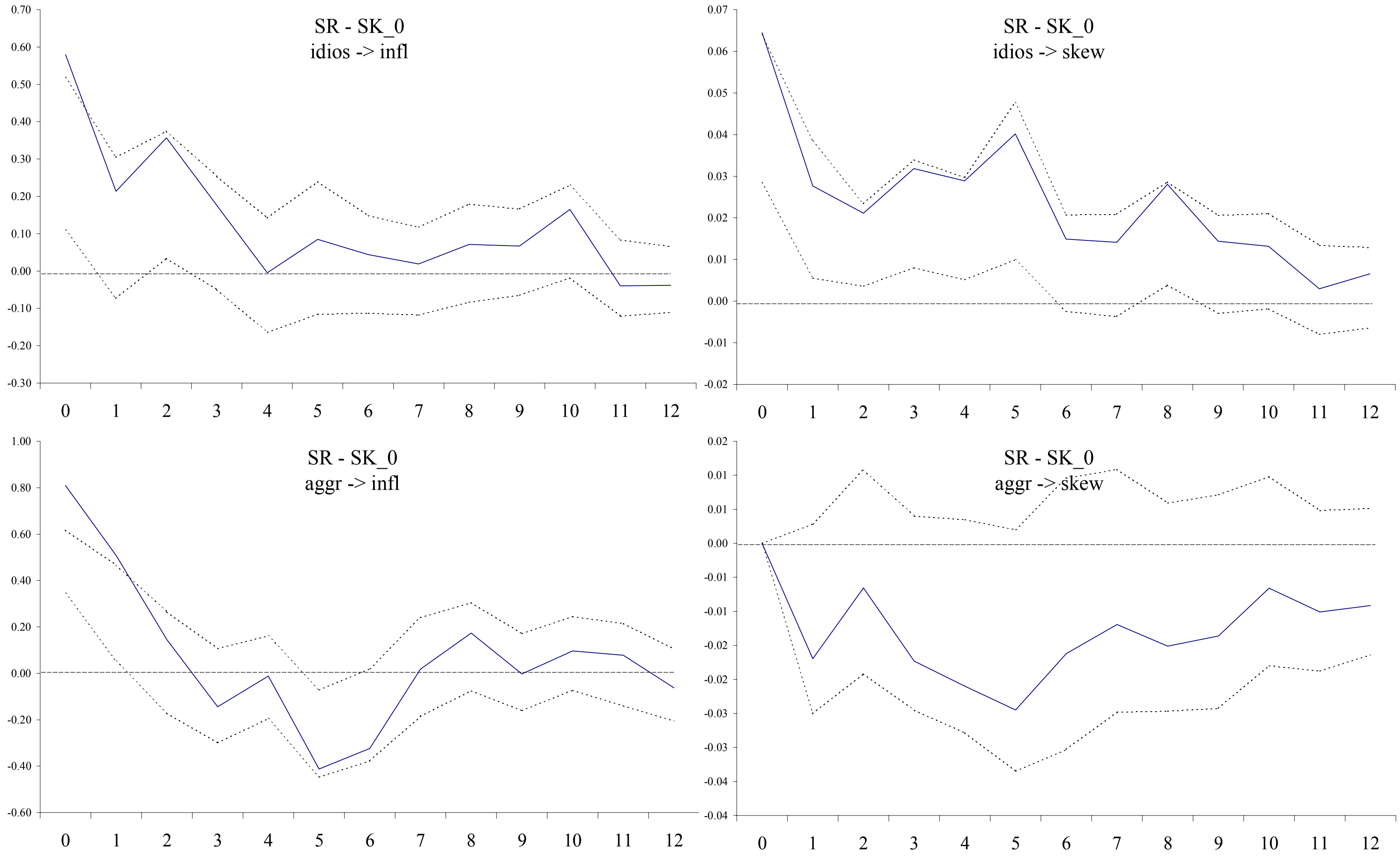
Note: Dashed lines are 90 percent Runkle (1987) confidence bands.
Horizontal axis: months following shock

Figure AIIa
Impulse Response Functions
(Pooled Data)



Note: Dashed lines are 90 percent Runkle (1987) confidence bands.
Horizontal axis: months following shock

Figure AIIb
Impulse Response Functions
(Pooled Data)



Note: Dashed lines are 90 percent Runkle (1987) confidence bands.
Horizontal axis: months following shock

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