ISSN 0111-1760

University of Otago
Economics Discussion Papers
No. 0702

February 2007

Regional Asymmetries in the Impact of Monetary Policy Shocks on Prices: Evidence from US Cities

David Fielding,
Department of Economics,
University of Otago,
PO Box 56, Dunedin,
New Zealand

Kalvinder Shields
Department of Economics,
University of Melbourne,
Victoria 3010,
Australia

First draft: February 2007

Abstract

Deviations from the Law of One Price across US cities are smaller than corresponding international deviations, but nevertheless substantial. We find that a proportion of these deviations can be explained by asymmetric responses to federal monetary policy shocks, and that a large part of the asymmetry can be explained by city-specific economic characteristics.

JEL Classification: E31, E52, R19

Key words: Law of One Price; monetary policy shocks

1. Introduction

Although deviations from the Law of one Price across regions or cities within countries such as the US or Canada are smaller than international deviations, they are still not negligible (Parsley and Wei, 1996; Ceglowski, 2003), and not confined to nontraded goods (Engel and Rogers, 2001). Moreover, convergence towards intra-national purchasing power parity happens very slowly; the half-life of deviations from the Law of One Price in the US is estimated to be about nine years (Cecchetti *et al*., 2002), regardless of the size of the initial deviation. Several of these authors find that transportation costs, proxied by distance, have a large role to play in explaining the average size of the deviations between two specific locations.

One question that has received relatively little attention is the source of the shocks that create deviations from the Law of One Price. Some reasons for the deviations – for example, asymmetric regional supply shocks – may be beyond the control of any policy maker. However, it is also possible that the deviations are partly *caused* by monetary policy makers, because of asymmetries in the response of prices in different locations to the same nationwide monetary policy shock. There is a small literature investigating asymmetries in the response of regional output levels to monetary policy (Toal, 1977; Garrison and Chang, 1979; Carlino and DeFina, 1998), but hardly any attention has been paid to regional price asymmetries.[1] In this paper, we explore the magnitude of price asymmetries by combining existing data from the literature on price deviations with existing data from the literature on monetary policy shocks. The specific questions we address are:

(i) Are asymmetries in the impact of monetary shocks in the US a statistically significant component of variations in cross-city bilateral real exchange rates?

(ii) How much of the variation in real exchange rates is explained by monetary policy shocks?

(iii) How large is the inter-city price divergence caused by a typical monetary policy shock?

(iv) Are there any city-specific characteristics that help to explain such divergence?

[1] Chamie *et al*. (1994) fit an SVAR to price and output data for nine US regions, and use this to estimate the degree of cross-regional correlation in real and nominal structural shocks. However, they do not look explicitly at the impact of shocks on relative prices.

Section 2 presents the time-series analysis that addresses question (i-iii) and Section 3 the cross-section analysis that addresses question (iv). We will see that there is clear evidence for statistically and economically significant asymmetries across US cities that explain a sizeable fraction of the observed variation in real exchange rates. These asymmetries are highly correlated with a number of city-specific economic and demographic characteristics. This suggests that aggregate nationwide prices do not contain all the information we need to calculate the welfare impact of changes in monetary policy.

2. Estimating the Size of Monetary Policy Asymmetries

In this paper we make use of two existing data sets. The price data are taken from Engel and Rogers (2001). Their monthly data set, originating from the Bureau of Labor Statistics, covers 43 components of the consumer price index (CPI) plus aggregate consumer prices for 29 cities over the period 1986(12)-1996(6). The 43 components account for about 99% of the aggregate index; the main item missing is medical insurance. These components are listed in the Appendix; note however that two individual price series are missing: dairy product prices in Tampa and other private transportation prices in Denver. With 29 cities we have 406 bilateral real exchange rates for each CPI component, except in the case of dairy products and other private transportation, where we have 28 cities and therefore 378 bilateral real exchange rates. These relative price series will be used to measure monthly deviations from the (proportional) Law of One Price, as explained below.

The second data set upon which we draw is from Romer and Romer (2004), who construct a monthly time series of US monetary shocks. First of all, the authors infer the Fed's intention for the federal funds rate at the time of FOMC meetings from quantitative and narrative records. A regression of this constructed series on the Fed's internal forecasts then provides a basis for a measure of unanticipated changes in policy that is free of systematic responses to information about the future. The authors show that their measure of policy shocks has large and statistically significant effects on both aggregate US output and aggregate US inflation. The effects are substantially stronger and quicker than those obtained using conventional indicators. The monetary shocks data cover the period 1969(1)-1996(12), which encompasses the period for which we have price data. The estimated monetary shocks for 1986(12)-1996(6) are depicted in Figure 1. The shocks for this period are somewhat smaller on average than for earlier years.

[Figure 1 here]

We combine the two sources of data to estimate the impact of monetary shocks on relative prices in different cities, proceeding as follows. Taking the prices indices from the Engel and Rogers database, we denote the value of CPI component i in city j in month t as p_t^{ij}, and the value of the aggregate CPI as p_t^{j}. We denote the monetary shocks series as $RESID_t$, as in the Romer and Romer data appendix. For every pair of cities (j, k) and for every price series (i) we fit a regression equation of the form

$$\beta^{ijk}(\mathrm{L})\,[\ln(p_t^{ij}) - \ln(p_t^{ik})] = \alpha^{ijk} + \theta^{ijk}(\mathrm{L})RESID_t + u_t^{ijk} \tag{1}$$

and similarly for the aggregate CPI:

$$\beta^{jk}(\mathrm{L})\,[\ln(p_t^{j}) - \ln(p_t^{k})] = \alpha^{jk} + \theta^{jk}(\mathrm{L})RESID_t + u_t^{jk} \tag{2}$$

where the β(L) and θ(L) are lag polynomial operators and u is an error term capturing non-monetary shocks. A lag order of 13 is used in the results reported below, so the sample period is 1988(1)-1996(12). $RESID_t$ is stationary by construction; the evidence from the papers cited in the previous section suggests that the real exchange rate $[\ln(p_t^{ij}) - \ln(p_t^{ik})]$ is stationary, but that convergence to the steady state is very slow. In most cases in our sample, we cannot reject the null that $[\ln(p_t^{ij}) - \ln(p_t^{ik})]$ is I(1), but with only ten year of data this comes as no surprise. When critical values are used in the results reported below it is assumed that $[\ln(p_t^{ij}) - \ln(p_t^{ik})]$ is stationary, but adjustment of the critical values along the lines of Stewart (2006) makes no substantial difference to our conclusions; further details are available on request.

Our first question is: are cross-city asymmetries in the impact of monetary shocks a statistically significant component of variations in the real exchange rate? Figure 2 provides some evidence on this question. This figures shows, for each CPI component (and for the aggregate CPI), the proportion of real exchange rate equations in which the F-statistic for the joint significance of the 13 θ parameters exceeds the 5% critical value. In the absence of any asymmetry, the parameter values in the true data generating process would equal zero. In this case, we should expect the F-statistic to exceed the critical value in 5% of the 406 regressions. However, the proportion exceeds 5% in all expenditure categories, and in most categories by a substantial margin. The proportion for the aggregate CPI is just short of 25%. There is strong evidence that there are statistically significant asymmetries in the way that prices in different cities respond to monetary shocks. There is no obvious pattern in the type of commodity for which the F-statistics are particularly large or small, on average. The two CPI components with largest proportion of rejections are educational expenses and fresh fruit and vegetables; the two with the smallest are household repairs and eggs. The asymmetry in the impact of monetary shocks is a feature of prices across a wide range of goods and services, and not confined to any one single type of expenditure.

[Figure 2 here]

Our second question is: how much of the variation in real exchange rates is explained by monetary policy shocks? Tables 1-2 provide some evidence on this question, suggesting that monetary policy shocks make a sizeable contribution. For every city, we compute the variance in the monthly rate of growth of each of the city's bilateral aggregate CPI real exchange rates over the period 1988(1)-1996(12). For every city there are 28 bilateral real exchange rates (28 other cities), and for each real exchange rate series a corresponding time-series variance. The first column in Table 1 reports, for each city, the average of these variances expressed percentage terms, in other words, $\frac{100^2}{28}\sum_{k \neq j} \text{var}\,[\Delta\ln(p_t^{\,j}) - \Delta\ln(p_t^{\,k})]$ for each j. (On average across all cities, the variance in monthly real exchange rate growth is about one quarter of a percentage point, implying a standard deviation of about one half of a percentage point.) The adjacent column reports in parenthesis the cross-sectional standard deviation corresponding to the city average. The next pair of columns relate to the average variance for a set of simulated real exchange rate series. For each real exchange rate series, the simulation measures the value that the series would have taken each month in the absence of any monetary shocks, using the estimated θ parameters and the non-monetary shocks u in equation (2). The third pair of columns relate to the difference between the average actual and average simulated variances. In all cases except one (Detroit) the average variance from the simulations is smaller than the average variance from the actual series: that is, monetary shocks have led to higher real exchange rate volatility in almost all cities. On average, the monetary shocks contribute about 0.02 percentage points to the total observed variance in a real exchange rate series, which constitutes about one twelfth of the total. The largest estimated effect is in New Orleans, where the monetary shocks add about 0.1 percentage points to the total observed variance.

[Tables 1-2 here]

Table 2 reports corresponding figures for the variance in annual rates of growth of the real exchange rate, that is, $\text{var}\,[\Delta_{12}\ln(p_t^{\,j}) - \Delta_{12}\ln(p_t^{\,k})]$ instead of $\text{var}\,[\Delta\ln(p_t^{\,j}) - \Delta\ln(p_t^{\,k})]$. In this case, the actual variance is about 0.9 percentage points across all cities on average, although for some individual cities the figures are much higher: 2.7 percentage points in Denver and 1.9 percentage points in New Orleans. These numbers are large relative to the US average annual rate of inflation over the period (3.5 percentage points). The overall average contribution of monetary policy shocks to these figures is moderate (about 0.08 percentage points), and in a few cities the variance in annual real exchange rate growth would have been marginally higher in the absence of the monetary shocks. However, there are several individual cities (Honolulu, Milwaukee, New Orleans, Saint Louis) in which the contribution of the shocks to the observed variance in real exchange rate growth exceeds 0.2 percentage points.

Our third question is: how large is the inter-city price divergence caused by a typical monetary policy shock? We answer this question by constructing an impulse response function for each city in which the dependent variable is $[\ln(p_t^{\,j}) - \frac{1}{29}\sum_k \ln(p_t^k)]$, that is, the aggregate price level in city *j* relative to the average aggregate price level across all cities. The impulse responses for each city in turn are depicted in Figure 3; these charts are constructed for a negative *RESID* shock equal to one sample standard deviation (0.15 points), using the estimated θ parameters from each equation (2) regression. In other words, we are plotting the response of relative prices to a generally inflationary policy shock. The figure plots the response of $[\ln(p_t^{\,j}) - \frac{1}{29}\sum_k \ln(p_t^k)]$ in each city up to a 30-month horizon, along with the corresponding two-standard-error bars; Table 3 lists firstly the values at the 12-month horizon and secondly the average of the values for the first 12 months.

[Figure 3 and Table 3 here]

By construction, the average value of $[\ln(p_t^{\,j}) - \frac{1}{29}\sum_k \ln(p_t^k)]$ is zero. However, as can be seen from the figure, the impulse response at the 12-month horizon is significantly different from zero in the majority of cities. Table 3 shows the largest increase in relative prices at the 12-month horizon (0.47 percentage points) is for Honolulu, and that the largest decrease (0.27 percentage points) is for San Diego. Since we are looking at an inflationary monetary policy shock, the table indicates that in absolute terms the largest 12-month price response is in Honolulu, and the smallest in San Diego. A typical inflationary monetary policy shock will raise prices in Honolulu relative to the US as a whole by 0.47 percentage points, and depress relative prices in San Diego by 0.27 percentage points. Prices in Honolulu relative to those in San Diego will rise by 0.75 percentage points. For a larger monetary policy shock – say, one equal to two sample standard deviations (0.3 points) – the change in relative prices will be twice as large. Again, these figures are substantial relative to the average US inflation rate. For such a shock, even prices in the city tenth from the top of the list (San Francisco) relative prices in the city tenth from the bottom (Pittsburgh) rise by over 0.2 percentage points. It remains to see why there are such asymmetries in the impact of monetary policy; this is the subject of the next section.

3. Explaining Monetary Policy Asymmetries

In this section we explore the causes of the cross-sectional variation in the data indicated in Table 3; our aim is to explain why an unanticipated fall in the interest rate raises prices more in some parts of the US than its does in others. We consider the following causes of asymmetries in the response of city-specific prices. Descriptive statistics for the variables below appear in Table 4.

(i) The share of interest-sensitive industries

As noted by Carlino and DeFina (1998), some industries may be more interest-sensitive than others. Therefore, total output and aggregate demand (and hence prices) will be more sensitive to monetary policy shocks in areas that are relatively intensive in the interest-sensitive industries. We are using data only from metropolitan areas, where agriculture and mining make up a negligible fraction of output. Nevertheless, we can control for the share of manufacturing in total employment, and also for the share of the wholesale and retail trade sector, which may be more sensitive to short-run fluctuations in consumer demand than other sectors. The Bureau of Economic Analysis publishes annual data on sectoral employment in each US metropolitan statistical area (MSA), and we use these employment shares, averaged over 1987-1996, to measure of the relative sizes of the manufacturing and trade sectors in each city.

(ii) Firm size

Small firms are more reliant on banks than are large ones, and monetary policy may affect the cost of bank loans more directly than it does the opportunity cost of other sources of finance (Bernanke and Blinder, 1988; Bernanke, 1993; Gertler and Gilchrist, 1993). Moreover, Oliner and Rudebusch (1995) produce evidence that the rise in the cost of borrowing for small firms following a monetary contraction is particularly large, possibly because credit market imperfections arising from informational asymmetries are particularly severe for small firms. In this case, cities in which small firms make up a larger share of the total should manifest a particularly high level of interest sensitivity. We control for variation in average firm size across different cities by using data from the US Census Bureau industrial census. In the results reported below, the measure used is the fraction of firms with fewer than 20 workers in each MSA; changing the cut-off point to ten, 50 or 100 workers makes no significant difference to our results, nor does using the average number of workers per firm.

(iii) Bank size

Kashyap and Stein (1995) argue that when monetary policy is tightened and reserves are restricted, large banks can find alternative sources of funding more easily and cheaply than small ones. In this case, aggregate demand in cities in which small banks predominate may be particularly sensitive to monetary policy shocks. In order to control for bank size, we use data published by the Federal Deposit Insurance Corporation. The FDIC Summary of Deposits lists the asset value of every bank operating in the US, along with the location of the bank's branches. Our bank size indicator for each city is the mean of the log of the asset value of all banks with branches in the MSA. (The means are not weighted by the number of branches; using weighted means produces similar results but with a somewhat larger standard error.) The data are for 1996, the last year of our time-series sample; data prior to 1994 are not available.

(iv) House prices

According to Caplin *et al*. (1997), property values affect the sensitivity of households to interest rate changes. Lower property values make it relatively difficult for households to refinance their mortgages to take advantage of a fall in interest rates. They show that refinancing is much more limited in states with lower property values. Therefore, aggregate demand in cities in which property values are relatively high should exhibit more sensitivity to monetary policy shocks. Our measure of property values, taken from the 1990 census, is the median value of specified owner-occupied housing units for each MSA.

(v) Demography

Inter-temporal substitution of consumption may be more difficult for certain age groups. It is difficult to defer expenditure on some children's goods and services, such as education, and, for a given level of bequests, risky to defer consumption in old age, when the probability of death in the near future is quite high. We therefore include two demographic measures, both taken from the

1990 census: the fraction of the population of the MSA under the age of 21 and the fraction over the age of 60. Changing these cut-off points to 16 and 70 respectively does not make any substantial difference to our results, and additional age dummies are statistically insignificant.[2]

[Tables 4-5 here]

Table 5 reports the result of a cross-section regression equation for the relative size of the effect of monetary policy shocks. The dependent variable is the average price effect over the first year, as listed in the second set of figures in Table 3. A larger value of the dependent variable indicates a relatively large rise in prices in response to a monetary expansion. Therefore, a positive coefficient on an explanatory variable indicates that a larger value of this variable is associated with greater sensitivity to monetary policy; a negative coefficient indicates that the variable is associated with less sensitivity. Note that the regression equation passes a Jarque-Bera test for residual normality only when we include a dummy variable for Buffalo. Buffalo has a relatively small number of banks (17), but these include some of the largest in the US (for example, Chase Manhattan, Chemical Bank and Citibank), so the bank size measure for Buffalo is highly atypical of a city of its size. All of the explanatory variables are statistically significant, except firm size and the proportion of the population over 60. The removal of these variables from the regression equation has no substantial impact on the size of the other coefficients. All together, the explanatory variables account for about three quarters of the cross-sectional variation in the data.

As anticipated, an increase in the share of wholesale and retail firms in total employment is associated with greater sensitivity to monetary policy shocks. A one percentage point increase in the share leads a rise in relative prices over the next year that is greater by about 0.05 percentage points per month on average. The coefficient on the manufacturing employment share is negative, however. A one percentage point increase in this share leads to an average price effect that is about 0.01 percentage points lower. That manufacturing intensity should be associated with less monetary policy sensitivity in our data contrasts with the result of Carlino and DeFina (1998), who find that the response of output to monetary policy is relatively large in manufacturing-intensive areas. However, the Carlino and DeFina regressions are for output rather than prices, and are based

[2] We did also explore the possibility that interest elasticities vary with net wealth over the life cycle, but this avenue did not lead to any significant interpretable results.

on state data rather than city data. (City data look very different from aggregate state data: compare the positions of Los Angeles, San Francisco and San Diego in Table 3.)

As anticipated, there is a negative coefficient on the bank size variable, indicating that cities with larger bank exhibit less sensitivity to monetary policy shocks. The regression coefficient is about -0.3. In order to interpret this coefficient, consider the estimated effect of reducing the bank size measure from its mean value (12.8) to its smallest observed value (11.7, Anchorage). This is predicted to increase the magnitude of a rise in relative prices following a monetary shock by 0.33 percentage points per month ([12.8 – 11.7] × 0.3). Consider also the estimated effect of increasing the bank size measure from its mean value to its largest observed value (13.7, San Francisco). This is predicted to reduce the magnitude of a rise in relative prices by 0.27 percentage points per month ([13.7 – 12.8] × 0.3).

The coefficient on house prices is positive, as anticipated, with a value of about 0.14. Reducing this variable from its sample mean (1.10) to the smallest observed value (0.67, Saint Louis) reduces the magnitude of a rise in relative prices following a monetary shock by 0.06 percentage points per month ([1.10 – 0.67] × 0.14). Increasing the house prices variable from its sample mean to the largest observed value (2.98, Minneapolis) increases the magnitude of a rise in relative prices by 0.26 percentage points per month ([2.98 – 1.10] × 0.14).

Finally, there is a negative coefficient on the fraction of the city population below the age of 21. This means that there is less interest sensitivity in cities with a relatively high proportion of children. A one percentage point increase in the fraction of the population below the age of 21 leads to an average price effect that is about 0.03 percentage points lower.

4. Conclusion

Evidence indicates that the marked deviations from the Law of One Price across US cities can be explained partly by asymmetric responses to monetary policy shocks. These asymmetries are both economically and statistically significant, and can be partly explained by cross-city variations in a number of economic and demographic characteristics.

Such asymmetry has potentially important consequences for monetary policy. Just as the theory of optimal monetary policy is founded on the idea of maximising the welfare of a representative agent, real-world monetary policy targets make reference to the rate of inflation of a national consumer price index, that is, the rate of growth of the cost of living for an “average” consumer. But if consumers in different regions face different prices then concerns about the distribution of inflation rates resulting from a given monetary policy become relevant. De Grauwe (2000) and Gros and Hefeker (2002) show how in theory a monetary policy rule that ignores information at the regional level may lead to welfare losses when there are asymmetries in the transmission mechanism. An optimal monetary policy reaction function needs to put more weight on regions where prices are relatively unresponsive (Benigno, 2004). Recent empirical studies of monetary transmission asymmetries in the EMU (Cecchetti, 1999; Elbourne and de Haan, 2004) reflect concern about this problem. Our evidence suggests that the concern is not necessarily limited to international monetary unions, and that national central banks may also need to review their reliance on national aggregate measures when conducting monetary policy.

Moreover, one should not imagine that asymmetries in the response of prices to a monetary shock will remain constant over the long run. The regional economic and demographic characteristics that explain the asymmetries are likely to change only slowly, but they will eventually change in response to regional economic development and structural adjustment. A deeper understanding of these processes could lead to significant improvements in the efficiency of monetary policy.

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Appendix: CPI Components and Codes

SA0	all items	SE29	furniture and bedding
SA1111	cereals and bakery products	SE32	other house furnishings
SA111211	meats	SA3111	men's and boys' apparel
SE06	poultry	SA3112	women's and girls' apparel
SE07	fish and seafood	SA3114	other apparel commodities
SE08	eggs	SE40	footwear
SA1113	dairy products	SE41	infants' and toddlers' apparel
SA11141	fresh fruits and vegetables	SE44	apparel services
SA11142	processed fruits and vegetables	SE45	new vehicles
SA1115	other food at home	SE46	used cars
SE19	food away from home	SE4701	motor fuel
SE20	alcoholic beverages	SA415	other private transportation
SE22	homeowners' costs	SE49	automobile maintenance and repairs
SE2101	residential rent	SE53	public transportation
SA2112	other rental costs	SA51	medical care commodities
SA221	fuels	SE56	professional medical services
SE27	other utilities and public services	SE57	hospital and related services
SA213	household maintenance and repairs	SA61	entertainment commodities
SE34	housekeeping services	SE62	entertainment services
SE33	housekeeping supplies	SA712	personal care
SA2313	household appliances	SA713	personal and educational expenses
SE28	textile house furnishings	SE63	tobacco and smoking products

Table 1: Actual and Hypothetical Variances in [$\Delta\ln(p_t^j) - \Delta\ln(p_t^k)$]

	Average actual variance		Average hypothetical variance		Difference	
Anchorage	**0.186**	(0.012)	**0.170**	(0.012)	**0.017**	(0.004)
Atlanta	**0.469**	(0.022)	**0.427**	(0.022)	**0.042**	(0.005)
Baltimore	**0.176**	(0.021)	**0.161**	(0.020)	**0.016**	(0.006)
Boston	**0.243**	(0.017)	**0.193**	(0.021)	**0.050**	(0.008)
Buffalo	**0.283**	(0.014)	**0.261**	(0.013)	**0.022**	(0.004)
Chicago	**0.136**	(0.011)	**0.120**	(0.010)	**0.016**	(0.005)
Cincinnati	**0.281**	(0.018)	**0.205**	(0.013)	**0.076**	(0.009)
Cleveland	**0.195**	(0.013)	**0.154**	(0.012)	**0.042**	(0.006)
Dallas	**0.232**	(0.021)	**0.176**	(0.019)	**0.056**	(0.008)
Denver	**0.331**	(0.018)	**0.318**	(0.021)	**0.014**	(0.006)
Detroit	**0.247**	(0.022)	**0.277**	(0.030)	**-0.030**	(0.011)
Honolulu	**0.347**	(0.013)	**0.267**	(0.010)	**0.081**	(0.008)
Houston	**0.353**	(0.021)	**0.319**	(0.020)	**0.035**	(0.006)
Kansas City	**0.178**	(0.011)	**0.156**	(0.011)	**0.022**	(0.005)
Los Angeles	**0.084**	(0.009)	**0.074**	(0.008)	**0.010**	(0.003)
Miami	**0.186**	(0.013)	**0.153**	(0.012)	**0.033**	(0.005)
Milwaukee	**0.267**	(0.012)	**0.214**	(0.012)	**0.053**	(0.006)
Minneapolis	**0.366**	(0.021)	**0.290**	(0.018)	**0.076**	(0.008)
New Orleans	**0.669**	(0.023)	**0.567**	(0.023)	**0.102**	(0.010)
New York	**0.078**	(0.006)	**0.062**	(0.005)	**0.016**	(0.004)
Philadelphia	**0.126**	(0.010)	**0.107**	(0.008)	**0.018**	(0.004)
Pittsburgh	**0.158**	(0.015)	**0.134**	(0.016)	**0.024**	(0.007)
Portland	**0.215**	(0.017)	**0.205**	(0.019)	**0.010**	(0.007)
Saint Louis	**0.269**	(0.021)	**0.217**	(0.020)	**0.052**	(0.008)
San Diego	**0.310**	(0.018)	**0.260**	(0.016)	**0.010**	(0.007)
San Francisco	**0.145**	(0.009)	**0.125**	(0.007)	**0.021**	(0.005)
Seattle	**0.181**	(0.013)	**0.149**	(0.010)	**0.032**	(0.005)
Tampa	**0.545**	(0.025)	**0.509**	(0.028)	**0.036**	(0.007)
Washington	**0.152**	(0.013)	**0.143**	(0.016)	**0.009**	(0.005)
Overall average	***0.255***		***0.205***		***0.018***	

Table 2: Actual and Hypothetical Variances in $[\Delta_{12}\ln(p_t^j) - \Delta_{12}\ln(p_t^k)]$

	Average actual variance		Average hypothetical variance		Difference	
Anchorage	**1.970**	(0.152)	**1.931**	(0.154)	**0.039**	(0.029)
Atlanta	**1.095**	(0.066)	**1.099**	(0.064)	**-0.004**	(0.028)
Baltimore	**0.373**	(0.054)	**0.354**	(0.062)	**0.019**	(0.021)
Boston	**1.242**	(0.140)	**1.244**	(0.154)	**-0.002**	(0.044)
Buffalo	**0.714**	(0.090)	**0.667**	(0.098)	**0.046**	(0.024)
Chicago	**0.308**	(0.039)	**0.248**	(0.045)	**0.060**	(0.024)
Cincinnati	**0.545**	(0.058)	**0.444**	(0.058)	**0.101**	(0.025)
Cleveland	**0.490**	(0.049)	**0.369**	(0.052)	**0.121**	(0.029)
Dallas	**0.521**	(0.088)	**0.543**	(0.087)	**-0.022**	(0.018)
Denver	**2.706**	(0.219)	**2.703**	(0.222)	**0.003**	(0.029)
Detroit	**0.523**	(0.058)	**0.351**	(0.046)	**0.172**	(0.039)
Honolulu	**1.975**	(0.163)	**1.764**	(0.164)	**0.211**	(0.044)
Houston	**0.960**	(0.065)	**0.898**	(0.079)	**0.062**	(0.036)
Kansas City	**0.834**	(0.085)	**0.789**	(0.099)	**0.046**	(0.036)
Los Angeles	**0.807**	(0.137)	**0.665**	(0.131)	**0.143**	(0.031)
Miami	**0.734**	(0.082)	**0.792**	(0.092)	**-0.058**	(0.037)
Milwaukee	**0.844**	(0.072)	**0.601**	(0.081)	**0.243**	(0.034)
Minneapolis	**1.274**	(0.093)	**1.192**	(0.086)	**0.082**	(0.023)
New Orleans	**1.852**	(0.123)	**1.636**	(0.117)	**0.216**	(0.044)
New York	**0.370**	(0.088)	**0.316**	(0.089)	**0.054**	(0.023)
Philadelphia	**0.328**	(0.066)	**0.279**	(0.067)	**0.049**	(0.017)
Pittsburgh	**0.327**	(0.058)	**0.191**	(0.054)	**0.137**	(0.025)
Portland	**0.424**	(0.069)	**0.424**	(0.074)	**-0.001**	(0.019)
Saint Louis	**0.718**	(0.068)	**0.434**	(0.055)	**0.283**	(0.044)
San Diego	**1.237**	(0.149)	**1.228**	(0.143)	**0.009**	(0.030)
San Francisco	**0.546**	(0.089)	**0.425**	(0.075)	**0.122**	(0.030)
Seattle	**0.876**	(0.098)	**0.862**	(0.108)	**0.014**	(0.026)
Tampa	**1.257**	(0.090)	**1.175**	(0.105)	**0.082**	(0.032)
Washington	**0.551**	(0.094)	**0.559**	(0.099)	**-0.008**	(0.015)
Overall average	***0.910***		***0.834***		***0.077***	

Table 3: The Effect of a Typical Monetary Policy Shock on Relative Prices

	price effect at one year	average price effect in first year		price effect at one year	average price effect in first year
Honolulu	0.474	0.520	Chicago	-0.001	-0.091
Los Angeles	0.214	0.204	Kansas City	-0.016	-0.011
Dallas	0.166	0.122	Cleveland	-0.020	-0.151
Boston	0.142	0.226	Baltimore	-0.021	-0.073
Miami	0.118	0.181	Pittsburgh	-0.027	-0.054
Tampa	0.102	0.096	Atlanta	-0.073	-0.047
Buffalo	0.088	0.195	Anchorage	-0.084	-0.017
Denver	0.080	0.042	Minneapolis	-0.138	-0.282
Houston	0.077	-0.178	Milwaukee	-0.165	-0.270
San Francisco	0.075	-0.026	Detroit	-0.176	-0.275
Portland	0.064	0.064	Seattle	-0.180	-0.102
Washington	0.059	-0.003	Saint Louis	-0.196	-0.309
Philadelphia	0.057	-0.072	New Orleans	-0.217	-0.384
New York	0.035	0.128	San Diego	-0.272	-0.313
Cincinnati	0.033	-0.119			

Table 4: Descriptive Cross Section Statistics

	mean	standard deviation
% employment in trade	22.0	1.7
% employment in manufacturing	11.7	4.7
% firms with fewer than 20 staff	85.3	1.6
banks' mean log $ asset value	12.8	0.4
median house value ($100K)	1.1	0.7
% population 21 or under	30.4	2.3
% population 60 or over	16.0	4.1

Table 5: Determinants of the Relative Size of the Effect of Monetary Policy Shocks

	coeff.	t ratio	coeff.	t ratio
intercept × 100	0.0354	1.38	0.0346	6.39
% employment in trade	0.0526	3.12	0.0533	3.54
% employment in manufacturing	-0.0090	-1.77	-0.0084	-2.37
% firms with fewer than 20 staff	-0.0016	-0.08		
banks' mean log $ asset value	-0.3016	-6.92	-0.2972	-7.54
median house value ($100K)	0.1470	3.48	0.1446	3.61
% population 21 or under	-0.0285	-1.63	-0.0312	-3.97
% population 60 or over	0.0025	0.24		
Buffalo dummy	0.4807	11.95	0.4827	13.33
R^2	0.775		0.774	
σ	1.345		1.284	
Schwartz-Bayesian Criterion	1.266		1.037	
Akaike Criterion	0.842		0.707	
Jarque-Bera test (p value)	0.379		0.323	
Ramsey RESET test (p value)	0.386		0.378	

Figure 1: Monetary Shocks (*RESID*) over 1986(12)-1996(6)

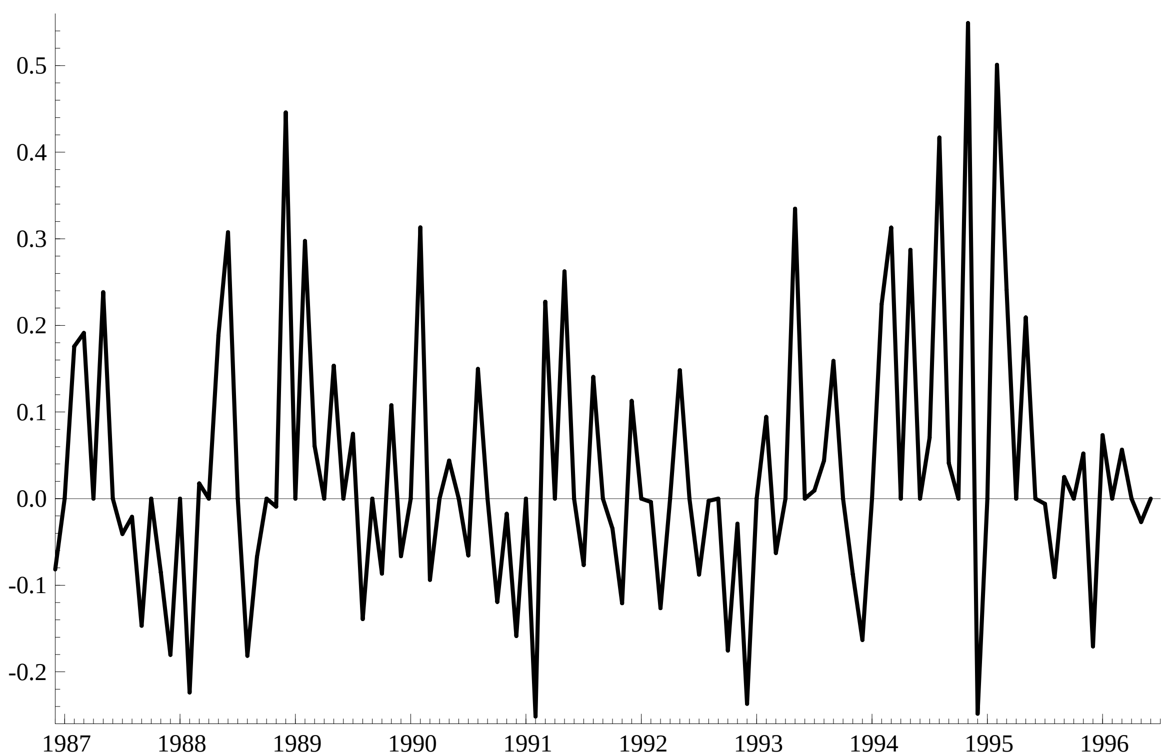

Figure 2: The Fraction of Cases in which Monetary Shocks are a Significant Determinant of Commodity-Specific Cross-City Real Exchange Rates, by CPI Item

See Appendix 1 for an explanation of the CPI component codes in this figure.

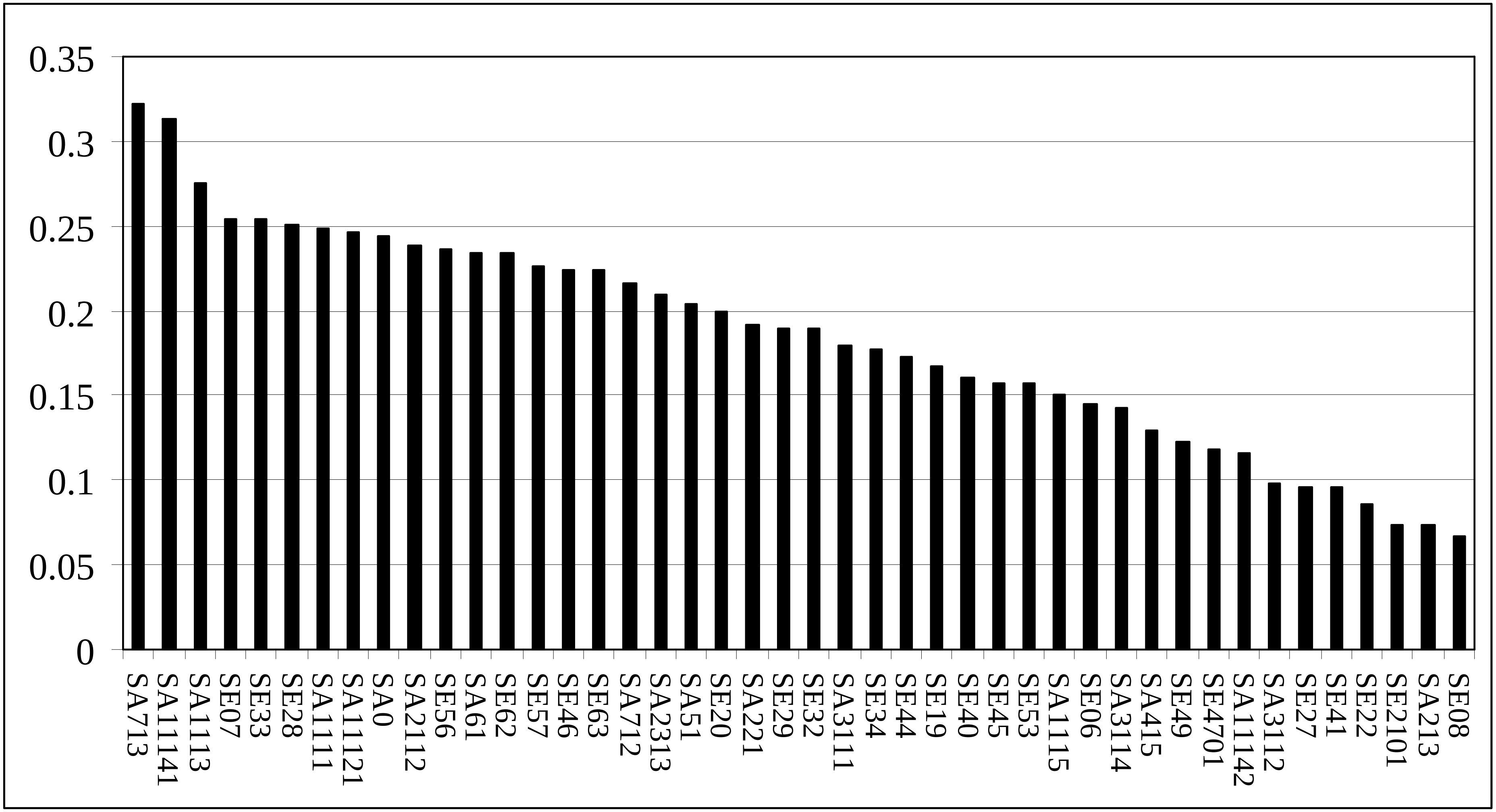

Figure 3: Impulse Responses for Relative Prices with a Typical Monetary Policy Shock

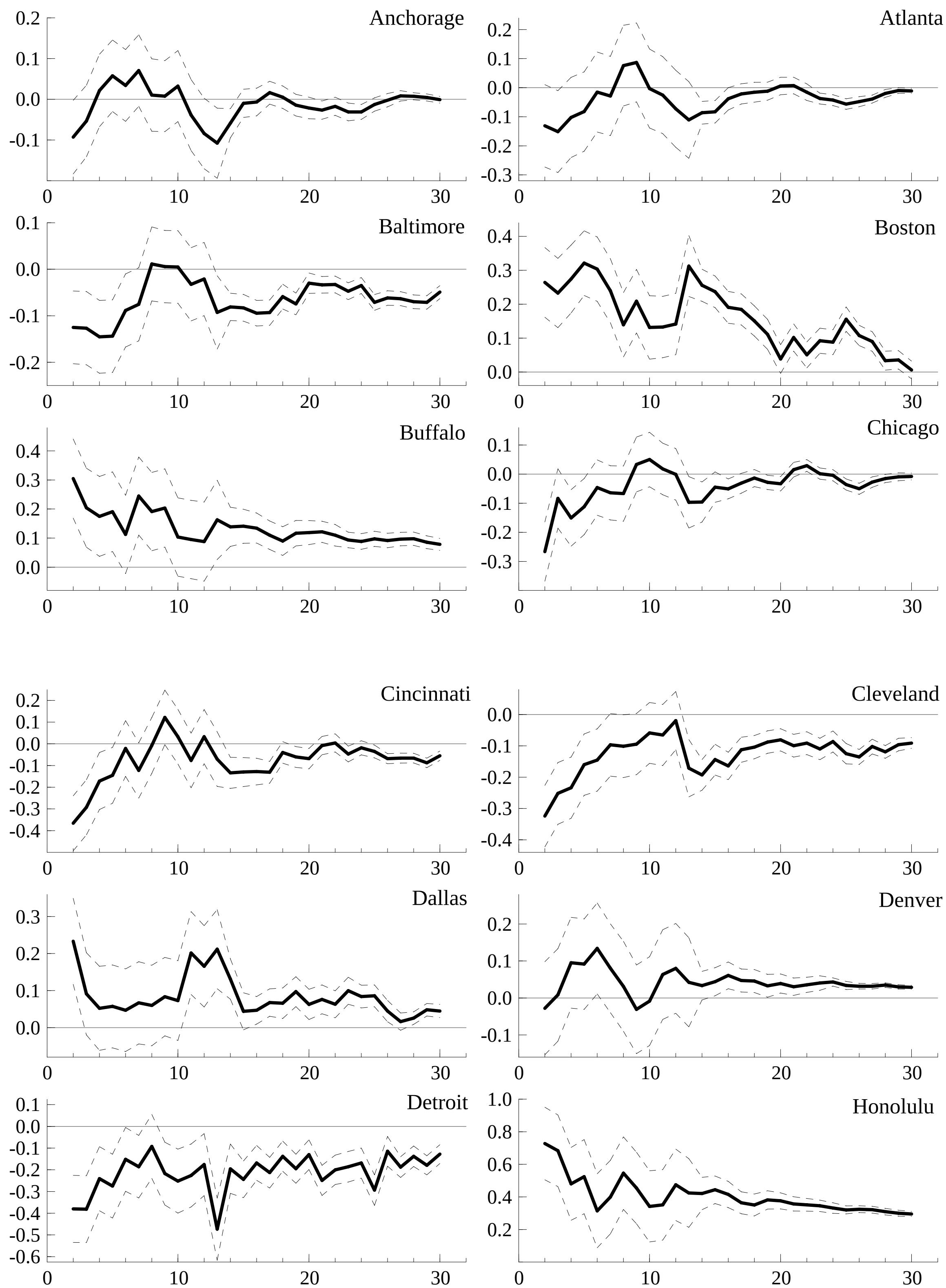

Figure 3 (continued)

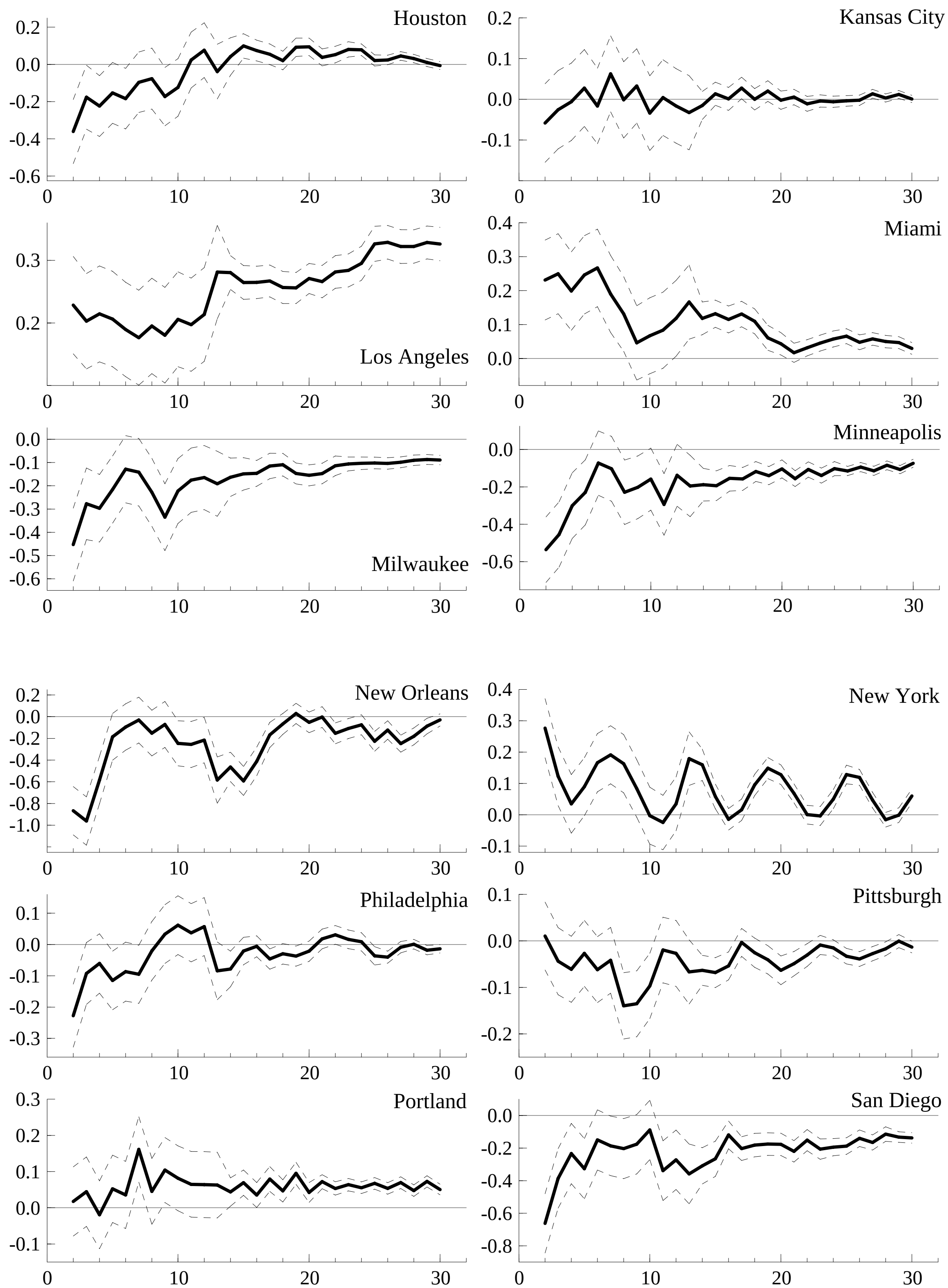

Figure 3 (continued)