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The Price Revolution in the 16th Century: Empirical Results from a Structural Vectorautoregression Model

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Abstract: This paper provides empirical evidence in favor of the hypothesis that the secular price increase in the 16th century is mainly caused by money supply developments as the discovery of new mines in Latin America. First we review price developments for several European countries over the 16th century in the light of this hypothesis. Second the application of a SVAR model to annual time series of price indexes for Old Castile and Leon and New Castile over the 16th century indicates that not only the trend but also the short to medium variability of price movements in 16th century Spain are dominated by permanent money supply shocks.

JEL: E31, E58, E65

Key words: Price revolution, money demand, money supply

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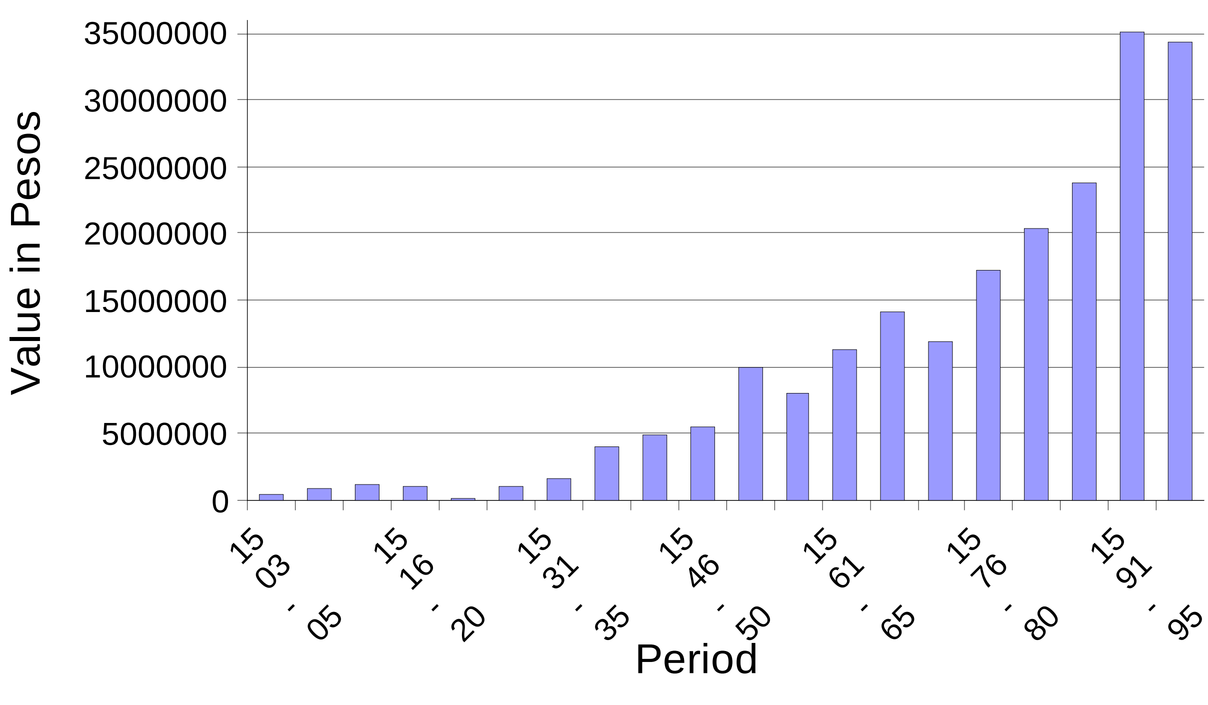
1. Introduction

During the 16th century we observe a steady price increase in all European countries: at the end of the century prices were three to four times higher than at the beginning of the century. This amounts to an annual inflation rate in the range of 1% to 1.5% which looks small given the inflationary experience of the 20th century. In deed this inflation rate is close to the target inflation rate of price stability oriented central banks of today. However, given the fact that the monetary regime of the 16th century was a metallic standard mainly based on silver this secular inflation is remarkable as such a monetary regime has an intrinsic price level stabilization property: rising commodity prices lead to a fall in the purchasing power of the monetary metal and correspondingly less incentive to mine it and more incentive to use it for non monetary purposes. This endogenous adjustment of the money supply leads to long run stability of the price level even if permanent shifts in money demand occur over time. Therefore long run inflation can only be explained either by debasement of coins or by shifts in the supply of the monetary metal: a permanent ceteris paribus increase in the productivity of mining the monetary metal leads to a fall of its relative price which is accomplished by rising prices for the other commodities. This process only comes to an end if the purchasing power of the monetary metal is equal to its production costs.

Therefore the obvious candidate for the great price revolution from the economic viewpoint is the discovery of new silver and gold deposits and the productivity increase in the silver mining industry in the 16th century. This process started in Central Europe around the beginning of the century. According to an estimation by North (1994, p. 74) central European silver output doubled between 1470 and 1520, and increased further in the 1520s with the new mine of Joachimsthal. In the 1530s, however, it declined sharply. Between 1492 and 1550 a substantial amount of gold, looted in the New World was brought to Europe by the Spanish and Portuguese. From the 1540s a rising supply of silver was shipped to Europe from the newly discovered mines in Mexico and Peru, where the silver mountain of Potosi was discovered in 1545. The output of the Potosi mine rose strongly in the 1560s after mercury deposits had been discovered in the Andes. Mercury was necessary to process the silver. According to Hamilton (1934) the total imports of treasure (gold and silver) from the New World during the 16th century amounted to 206.6 million pesos, of which only about 15 % were imported before 1555 (Figure 1.1). A peso had a silver content of 18.95 g. Thus the total amount corresponded to about 3915 metric tons of silver. But as Kindleberger mentions: "It is generally recognized that Hamilton's estimates understate imports for the early years of the seventeenth century, because he counted only those imports recorded by the official Casa de Contratacion in Seville. Dutch and English East India ships directly in Cadiz, downstream from Seville. In addition, considerable amounts of silver went from Peru to Acapulco in Mexico between 1573 and 1815, and from thence to the Philippines ..." (1998, p.3). The same should also hold for the second half of the 16th century, though to a lesser degree.

Figure 1.1
Total Imports of Treasure from the New World to Spain in Pesos, 16th Century



Note: Five year periods.

Source: Hamilton (1934)

The “monetary” view of the price revolution is questioned by historians who attribute the secular increase of commodity prices to population and income growth as well as urbanization and wars (see, e.g. Outhwaite 1969 and North 1990). Thus it is argued that the rise of prices began well before the arrival of gold and silver from the New World. But as has already been mentioned the production of silver grew substantially in Central Europe at least since the beginning of the century, and quite substantial amounts of gold were imported during its first half. Further, the critics argue that their hypothesis that the rise of the price level was a consequence of population and income growth is supported because the prices of victuals rose more than that of other goods and wages (see e.g. Figure 2.1). But with this argument only the change of relative prices can be explained. The critics' view is based on a misinterpretation of Irving Fisher’s equation of exchange, as has been clearly outlined by Flinn(1984). In fact most of the causes mentioned lead to an increase in money demand and therefore to short to medium deflationary pressure with a tendency for long run price stability. For instance, if the population is growing with the same real per capita income, the demand for money will increase. For with an unchanged money supply a pressure on the general price level will result, though it is possible that the prices for some foodstuffs, notably grain may rise even in absolute terms. Nevertheless, together with strong fluctuations in harvests these money demand shifts, though they cannot explain the rising price level, may be important for the considerable variability of price changes we observe during the 16^{th} century.

This paper provides an empirical analysis of the relative importance of money demand and supply shocks using price data for Old Castile and Leon and New Castile over the 16^{th} century. To this end we apply the Structural Vectorautoregression model (SVAR) which allows us to identify a permanent (money supply caused) and transitory (money demand caused) component of the price level.

The content of the paper is organized as follows: section 2 contains a brief discussion of the price revolution in the 16th century. The SVAR approach is described and applied in section 3. Section 4 concludes.

2. The Price Revolution in the 16'th Century

The tendency of rising prices during the 16th century is a general feature observed in different countries of Europe. It has been noted since more than a century, and has been amply documented for France, the Netherlands and what is now Belgium (brabant and Flanders) besides for the different regions of Spain. Already in the Handwoerterbuch der Staatswissenschaften of 1901 (vol. VI, p. 220) we find the following figures for the development of the price of wheat in several countries (Table 2.1).

Table 2.1

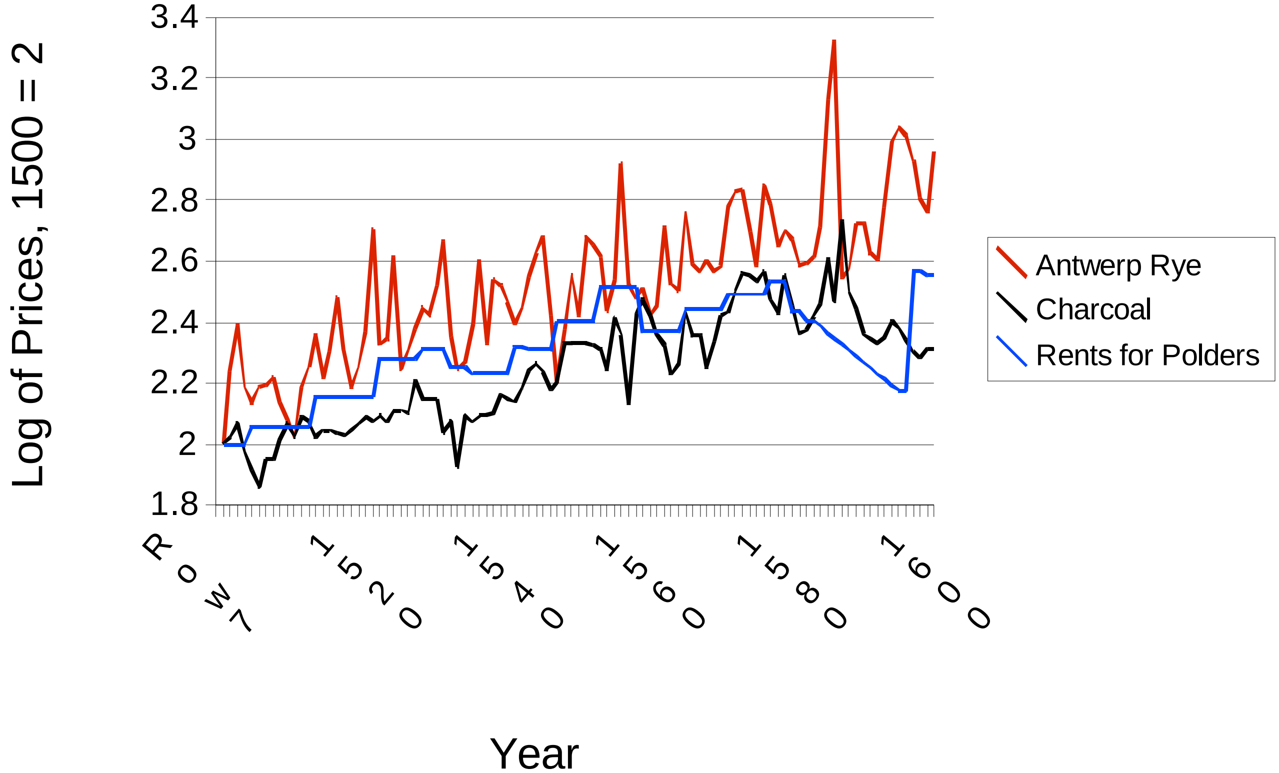
The Development of Wheat Prices in Several European Countries

	Country		
Period	Saxony	Paris	England
1455-80	13.9		
1511-20			20.59
1521-30		31.28	21.96
1531-40	29.8	32.96	21
1541-50	25.8	34.33	17.02
1551-60	31.59	36.03	28.85
1561-70	41.32	64.05	26.22
1571-80	53.04	71.08	34
1581-90	60.43	82.33	
1583-92			48.32
1590-1600		145.13	
1593-1602			70.68
1601-1610		63.48	
1603-1612			69.88

(Measured in Silver)

In looking at the figures two points are noteworthy. First, dramatic price increases begin only in the second half of the century. This should rather support the hypothesis that the inflow of silver from the New World has been decisive for this development. Second, that though five year averages have been used, some strong fluctuations are still visible.

Figure 2.1 Development of Log of Prices in Silver in Brabant, 16th Century



Notes: For wheat and charcoal nnual figures provided in Brabant groat have been converted according to the silver and gold content of the groat. This was necessary since the groat was devalued several times. For instance, whereas its silver content had been 0.33 g in 1499, it had shrunk to 0.19 g in 1600.
The figures for the price of polders were already provided as an index in the source. But the data for 1584-94 are missing, since the dikes were broken. They have been interpolated in the graph.
Source: van der Wee (1963, pp. 479 ff., 255 f., 177 f., 133 ff. and 128 f.).

A similar development of prices can be observed in two towns of the Netherlands (Table 2.2). In this case indexes are presented. Here again we note the acceleration of inflation in the second half of the century. Strong fluctuations are also present, though besides five year averages several goods are now contained in the indexes.

Table 2.2

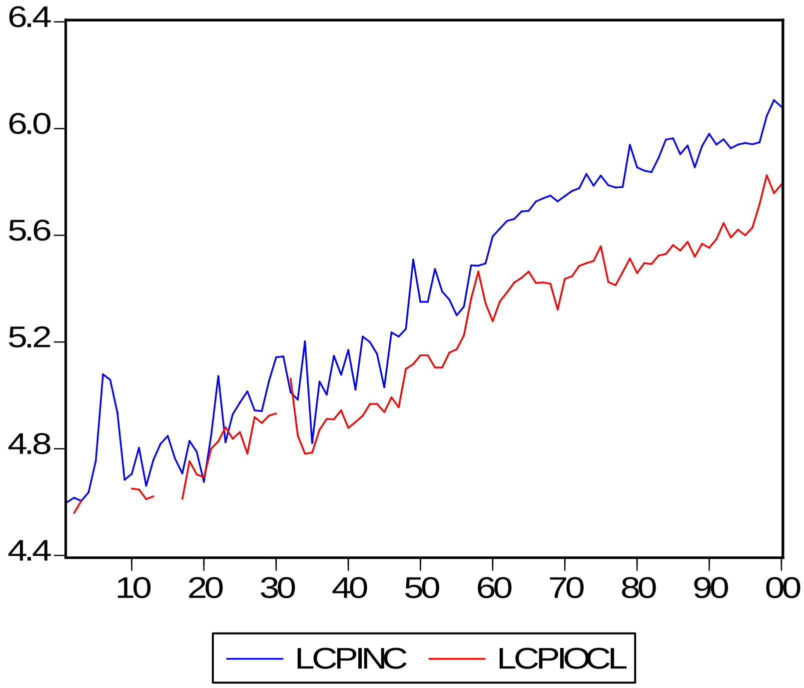
Development of Price Index in Two Dutch Cities

	City	
Period	Leyden	Utrecht
1501-05	100	100
1506-10	106.23	113.42
1511-15	111.92	118.28
1516-20	117.79	118.22
1521-25	139.59	138.22
1525-30	123.67	155.69
1531-35	141.28	122.4
1536-40	140.3	118.34
1541-45	146	127.38
1546-50	144.04	128.43
1551-55	187.72	138.58
1556-60	198.75	163.88
1561-65	211.12	172.92
1566-70	233.45	187.08
1571-75	297.69	251.63
1576-80	344.4	276.62

Source: Posthumus (1964).

Finally we turn to the development of prices in Spain. In Figure 2.2 the logarithmic price indexes of Old Castile (LPIOCL) and Leon and New Castile (LPINC) are displayed for the years 1500-1600. Two characteristics of price developments in these two parts of Spain are evident from this graph: There is a positive trend inflation of about 1.1% pa and 1.4% pa in Old Castile and Leon and New Castile respectively. This trend behavior, which appears to be across the two regions, is not smooth but characterized by large shocks, that even lead to sometimes deflationary episodes despite the positive trend inflation rate. The common trend is to be expected from economic theory as both regions had the same metallic standard. The high volatility of price developments is, however, remarkable and leaves room for the importance of money demand shocks. In particular, the deflationary episodes may be explained by good harvests or (and) by a scarcity of money as often alleged in historical sources. However, the positive price trend can, according to economic theory, only be explained by a trend growth of money supply caused by reduced production costs of the monetary metal.

Figure 2.2 Log Commodity Price Index for Old Castile (LPIOCL) and Leon and New Castile (LPINC)



3. Econometric Results from an SVAR Model

This section presents the results of the application of SVAR models to the log price indexes of Old Castile (*y*) and Leon and New Castile (*x*). As mentioned in section two this model allows to identify a permanent shock, which is interpreted as a permanent productive shock in the production of the monetary metal, as well as a transitory shock representing mainly money demand disturbances. The model is based on the assumption that the series investigated follow a stochastic trend. That means that the series have to be differenced once in order to get a stationary stable series. This framework implies that the series do not revert to a linear deterministic trend as random shocks have a permanent effect on the level of the series. In other words there is a random walk with drift component of the series. The application of the usual statistical tests (unit root tests, stationarity tests) leads to the conclusion that the hypothesis of a stochastic trend in the data series cannot be rejected at usual significance levels. Moreover, a cointegration test indicates that the two logarithmic price series are driven by one common trend as we expect from economic theory. The detailed results of these tests are found in an appendix.

Our analysis is based on the following non-stationary VAR model with lag length 1[1]:

$$\begin{aligned} y_t &= b_{10} + a_{11} y_{t-1} + a_{12} x_{t-1} + \varepsilon_{1t} \\ x_t &= b_{20} + a_{21} y_{t-1} + a_{22} x_{t-1} + \varepsilon_{2t} \end{aligned} \qquad (1)$$

This model is the reduced form of an unspecified dynamic model of price dynamics. Therefore, the error terms are in general linear combinations of two structural shocks *u* which are mutually uncorrelated with expected value zero and variance 1:

[1] The lag length of one is indicated to be appropriate according to the Akaike, Schwarz and Hannan-Quinn criterion as well as a sequential likelihood ratio test.

$$\varepsilon_{1t} = b_{11}u_{1t} + b_{12}u_{2t}$$
$$\varepsilon_{2t} = b_{21}u_{1t} + b_{22}u_{2t} \qquad (2)$$

Iterating (1) backwards and substituting (2) provides the vector moving average representation of the VAR model:

$$y_t = c_{10} + \sum_{i=0}^{\infty} c_{11}(i)u_{1t-i} + \sum_{i=0}^{\infty} c_{12}(i)u_{2t-i}$$
$$x_t = c_{10} + \sum_{i=0}^{\infty} c_{21}(i)u_{1t-i} + \sum_{i=0}^{\infty} c_{22}(i)u_{2t-i} \qquad (3)$$
$$coefficients\ in\ matrix\ notation : C(i) = A^i B$$

Equation (3) shows us the reaction of the two observed variables to current and past shocks and the *c*-functions are the impulse response function. In order to calculate this impulse response functions we need the *a*-coefficients as well as the *b*-coefficients. The former can be directly estimated by application of OLS to the two equations in (1). However, for the estimation of the *b*-coefficients we need an additional restriction as OLS estimation of (1) provides us only with three estimates, namely the variances and the covariance of the error-terms which depend in turn on the four *b*-coefficients. Usually this problem is solved by restricting the *B*-matrix to be lower triangular, i.e. a recursive ordering of the contemporaneous impact of the shocks. This is not very meaningful in our application as we have to expect that the two price indexes are hit by money demand and money supply shocks the same time. However, there are meaningful long run restrictions for our shocks, namely the only money supply shock ($u2$) has a long run impact on the two price indexes and that the long run effect of the money demand shock ($u1$) is zero. This approach corresponds to the structural VAR model with long run restrictions introduced by Blanchard and Quah (1989). In fact we need only one of the two restrictions to identify the impulse response functions. We choose the restriction that the long run effect of the money demand shock on log prices in New Castile is zero. His approach has the advantage that we can check empirically the second restriction that long run effect of the money demand shock on log prices in Old Castile and Leon is zero.

The VAR-estimates for the sample covering the years from 1501-1600 are reported in Table 3.1. These coefficient estimates show a feedback pattern between the two series. There is a strong and highly significant effect of lagged prices in Old Castile and Leon on prices in New Castile, which in turn have a significant but less strong effect with one year lag on prices in Old Castile and Leon.

The SVAR impulse response functions and variance decomposition estimates are shown in Figures 3.1 and 3.2. The effects of the shock $u1$ which is interpreted as money demand shock according to our long run identification scheme quickly goes to zero within three years. Note that we restricted the model only with respect to the long run effect of the money demand shock to LPINC. The fast convergence to zero and the near zero long run effect of the money supply shock on LPIOCL is not imposed but is delivered the data. By contrast the permanent money supply shock $u2$ affects both prices series quickly and is approximately the same for both series in the long run. The variance decomposition which shows us the contribution of the money demand and supply shock to the forecasting variance for the twos series confirms the results of the impulse response analysis: with a forecasting horizon of 3 and more years we see a clear dominance of the money supply shock in the sense that at least three thirds of variability of the two price indexes are caused by money supply shocks. Thus not only the trend but also the short to medium variability of price movements in 16th century Spain are dominated by the permanent money supply shocks. Finally we should note that our results are robust with respect to the sample size: estimating the model with data from 1540 to 1600 (the period during which most of the inflow of Latin American silver occurred) produced essentially the same impulse response and variance decomposition results[2].

Table 3.1: VAR Estimates for LCPIOCL, LCPINC; 1500-1600

Vector Autoregression Estimates
Sample (adjusted): 1503 1600
Included observations: 85 after adjustments
Standard errors in ( ) & t-statistics in [ ]

	LCPIOCL	LCPINC
LCPIOCL(-1)	0.763604	0.746447
	(0.08904)	(0.12775)
	[8.57560]	[5.84312]
LCPINC(-1)	0.175583	0.394452
	(0.06939)	(0.09955)
	[2.53038]	[3.96228]
C	0.288288	-0.574488
	(0.12628)	(0.18117)
	[2.28299]	[-3.17107]
R-squared	0.968904	0.960335
Adj. R-squared	0.968146	0.959368
Sum sq. resides	0.291865	0.600736
S.E. equation	0.059660	0.085592
F-statistic	1277.511	992.6662
Log likelihood	120.5401	89.86091
Akaike AIC	-2.765650	-2.043786
Schwarz SC	-2.679439	-1.957575
Mean dependent	5.218989	5.457356
S.D. dependent	0.334273	0.424620
Determinant resid covariance (dof adj.)		2.25E-05
Determinant resid covariance		2.10E-05
Log likelihood		216.6121
Akaike information criterion		-4.955578
Schwarz criterion		-4.783156

[2] The likelihood ratio test of coefficient stability over the 1500-1539 and 1540-1600 periods rejects the null hypothesis of the same coefficients in both periods (chi-squared 22.64 with 6 degrees of freedom). This indicates that there is a change in price dynamics during the century in Spain which does not strongly effect our IR and VD results. This can be probably explained by the fact that before 1540 the Central European mines were the main source of new silver in Europe.

Figure 3.1: Impulse Response for LCPIOCL, LCPINC; 1500-1600

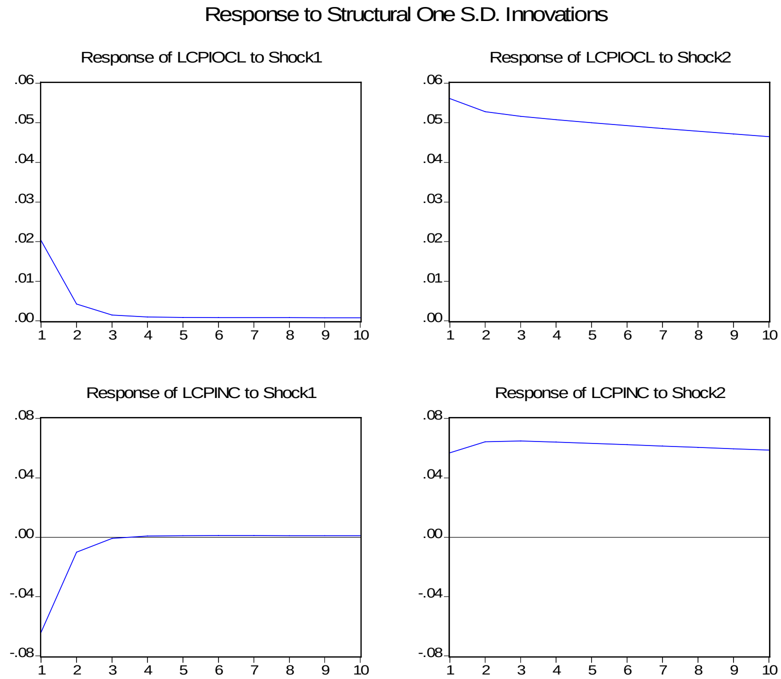
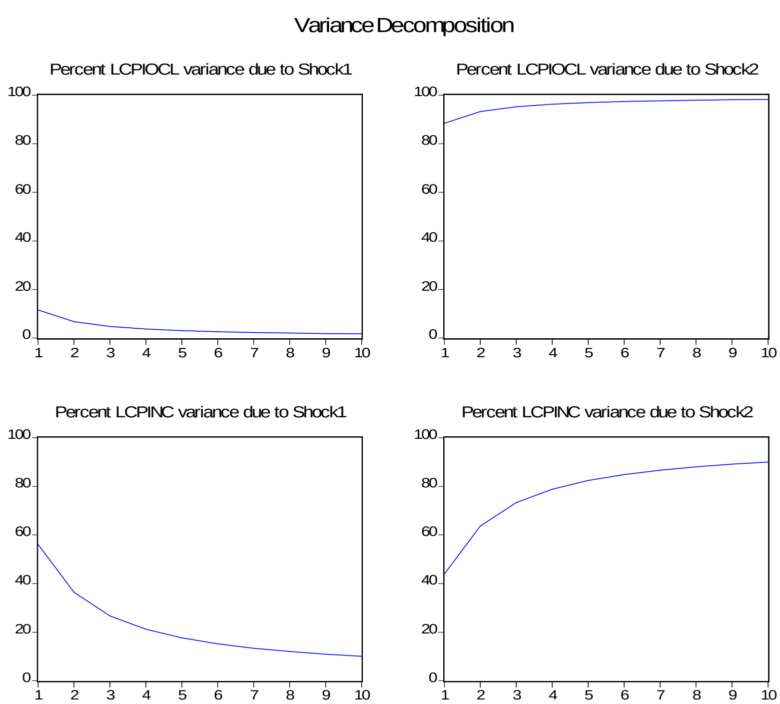


Figure 3.2: Variance Decomposition for LCPIOCL, LCPINC; 1500-1600



4. Conclusion

Economic theory suggests that the secular price level increase during the 16th century in Europe (the price revolution) is caused by the discovery of new silver and gold deposits and the productivity increase in the silver mining industry in the 16th century. This “monetary” view of the price revolution is questioned by historians who attribute the secular increase of commodity prices to population and income growth as well as urbanization and wars. This paper provides empirical evidence in favor of the economic point of view. First we review price developments for several European countries over the 16th century in the light of this controversy. Second we provide an empirical analysis of the relative importance of money demand and supply shocks using annual time series of price indexes for Old Castile and Leon and New Castile over the 16th century. To this end we apply the Structural Vectorautoregression model (SVAR) which allows us to identify a permanent (money supply caused) and transitory (at least partly money demand caused) component of the price level. This analysis indicates that not only the trend but also the short to medium variability of price movements in 16th century Spain are dominated by the permanent money supply shocks.

Appendix: Unit root and cointegration properties of Spanish price indexes in the 16th century

Before modeling the dynamic interrelationship between the different price indexes we have briefly to analyze their unit root and cointegration properties. First we test the unit root hypothesis (i.e. the series follow random walk trends and have infinite variance) against the alternative of trendstationarity (i.e. the series are reverting to a deterministic linear trend and have finite variance). We used the logarithms of the time series. The first column of Table A.1 reports the augmented Dickey Fuller t-Test for the unit root hypothesis, which cannot be rejected at the 10 percent level in both cases. As it is well known that unit root test suffer from size and power problems in small samples we consider secondly the alternative test of the null of stationarity suggested by Kwiatkowski, Phillips, Schmidt and Shin. This approach leads to a rejection of the hypothesis of trend stationarity at least at the 10 percent level in both cases.

Table A.1: Results of Unit Roots and (Trend-)Stationarity Tests, 1501-1600

Series	ADF	KPSS
LCPIOCL	-1.4167	0.1193*
LCPINC	-3.2459	0.1267*

*, ** and *** indicates statistical significance at the 10%, 5% and 1% level, respectively.

Test regression includes a constant and a deterministic trend. The ADF lag length is determined by the Modified Akaike Information Criterion (Ng and Perron, 2001). The band width for the Bartlett window in the KPSS test is 6.

Moreover, we would expect all these series share a common trend in the sense that there is a long run convergence of price developments in the regions. That means that even if all the series follow a unit root trend there are linear combinations (e. g. LCPIOCL-LCPINC) which are stationary and the series are called cointegrated. This hypothesis can be tested by the multivariate cointegration test developed by Johansen: The results obtained are reported in Table A.2. This test clearly rejects the hypotheses of no (stationary linear combinations). These findings support the hypothesis of a common random walk trend of all three price series and are in line with economic theory.

Table A.2: Johansen Test: LCPIBR, LCPIOCL, LCPINC; 1501-1539

Date: 12/28/06   Time: 15:18
Sample (adjusted): 1503 1600
Included observations: 85 after adjustments
Trend assumption: Linear deterministic trend
Series: LCPINC LCPIOCL
Lags interval (in first differences): No lags

Unrestricted Cointegration Rank Test (Trace)

Hypothesized No. of CE(s)	Eigenvalue	Trace Statistic	0.05 Critical Value	Prob.**
None *	0.436019	49.41429	15.49471	0.0000
At most 1	0.008572	0.731773	3.841466	0.3923

Trace test indicates 1 cointegrating eqn(s) at the 0.05 level
* denotes rejection of the hypothesis at the 0.05 level
**MacKinnon-Haug-Michelis (1999) p-values

Normalized cointegrating coefficients (standard error in parentheses)

LCPINC	LCPIOCL
1.000000	-1.257819
	(0.03505)

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