Real Exchange Rates in the European Price Revolution

Gregor W. Smith[†]

June 2026

Abstract

The European price revolution of the 16th and 17th centuries is one of the earliest episodes of widely-recognized inflation. I use inter-city real exchange rates to assess how extensively it spread, how rapidly long-run PPP was restored, and whether periodic debasements affected relative prices. Inflation shocks originating in Spain gradually diffused across Europe, but there is little evidence of differential impacts related to distance. Overall the properties of real exchange rates are strikingly similar to those in modern European data.

JEL classification: E31, N13.

Keywords: price revolution; real exchange rate; purchasing power parity

[†]Department of Economics, Queen's University, Canada; gregor.smith@queensu.ca. I thank the Social Sciences and Humanities Research Council of Canada for support of this research and Angela Redish, James McNeil, James Nason, Kivanc Karaman, Mauricio Drelichman, Viktoria Hnatkovska, and seminar participants at the Vancouver School of Economics for comments.

1. Introduction

The term 'price revolution' refers to a period of relatively high inflation in the 16th and 17th centuries in Europe. Historians have long attributed this episode to the flow of precious metals (especially silver) from the Americas. In studying modern economic episodes we are used to the idea that the international spillovers of such a persistent, monetary expansion depend on the exchange-rate regime and on the integration of markets for goods and assets. For the early modern period we may be less sure of how to characterize both the exchange-rate arrangements and the inter-connectedness of economies. Did European countries operate as if on a silver standard or as if they had flexible exchange rates? And were goods markets far less integrated than they are today?

Observing the properties of real exchange rates indirectly answers these questions. I use Allen's (2001) consumer price indexes for European cities to construct real exchange rates. This provides a panel of cities for the period of the price revolution: Amsterdam, Antwerp, Augsburg, Krakow, London, Munich, Paris, Strasbourg, Valencia, and Vienna. I treat Valencia as the base city because it is the only city in the Allen panel that is in Spain—where the price revolution is generally held to originate—with a price index that both spans the historical period and is comparable to those in other European cities.

I first document the stationarity of real exchange rates, which suggests that the price levels had a common trend across cities, possibly extending as far northeast as Krakow. I then report on the half-life of deviations from long-run purchasing power parity (PPP). It is 3.21 years, which is similar to estimates in a wide range of modern studies.

The price indexes are available in silver and in local currency, so one also can study the effects of local depreciations or debasements in the silver content of currencies. I use these measures to constuct a synthetic analogue to modern nominal exchange rates. I find that these played little role in the adjustment of real exchange rates. Debasements were reflected in local-currency prices but had little effect on prices measured in silver. Thus real exchange rates behaved much as in a modern setting with fixed nominal exchange rates.

Finally, I use local projections (LPs) to trace the effect over time of a shock to inflation in Valencia on inflation in the other cities. For a shorter time span I also do the same thing with a shock to inflation in Madrid. Such a shock caused an initial, real appreciation in these Spanish cities but then gradually diffused to other inflation rates. However, there is little evidence that differences in diffusion rates across cities are related to distance from Spain.

As background, Munro (2008) offers a concise introduction to research on the price revolution. The dates here (focusing on 1503–1650) follow the classic book by Earl Hamilton (1934). The chapter by Braudel and Spooner (1967) is a second classic reference that provides both context and detail. I note that the present paper is not about the causes of the price revolution. Palma (2020, 2022) thoroughly documents and discusses the diverse impacts of American precious metals on European countries. I focus instead on the differential inflation experiences of a large panel of cities.

2. Allen Price Indexes

The main data are annual, city-level CPIs from Allen (2001). The CPIs are Laspeyres indexes. The underlying prices come from the records of institutions, as is typical of early price history. The construction of price indexes is comparable across cities, though elements of the baskets differ, just as in modern price indexes. As examples, olive oil and wine enter in Valencia, while butter and beer do so in London, rye bread is consumed in Krakow and wheat bread in Paris, and less fuel is burned in Valencia than in London.

I include cities for which there are data over 1503–1650: Amsterdam, Antwerp, Augsburg, Krakow, London, Munich, Paris, Strasbourg, Valencia, and Vienna. Table 1 gives the cities, their 3-letter labels, and their full time spans. I omit Allen's data for cities where the CPI begins midway through the price revolution: Lwow (1520), Gdansk (1535), Naples (1548), Warsaw (1558), and Leipzig (1566) and for Northern Italy where the series ends in 1620. Madrid plays a lesser role in this study because its price data begin in 1551. Madrid (in Castile) and Valencia (in Aragon) did not form a currency union during this period. But I treat Valencia as the benchmark city because of the long time span of its price index.

The Allen data do not include rent but several recent studies have carefully measured early modern residential rent and included it in price indexes. Drelichman and González

Agudo (2014) measure rents in Toledo for 1490–1600. They combine this measure with a price index for Madrid that excludes rent to produce an overall index. They also integrate rent measures into Allen's indexes for Amsterdam and Antwerp, for comparison. González Agudo (2019) constructs a new price index for Toledo for 1521–1650 that features a rent measure, some manufactured goods, and several basket changes over time.

I do not study real exchange rates using these broader indexes, for two reasons. First, Allen (2001) and Drelichman and González Agudo (2014) estimate that rent makes up 5–10% of the urban consumption basket. Both studies conclude that the effect of including rent on the price indexes is 'detectable but not dramatic.' Second, for measuring things like distance effects in reversion to PPP (as in this study) a large cross-section dimension is important. To my knowledge there is no panel of price indexes that includes rent on a consistent basis for many cities.

It is possible that the speed of reversion to PPP would be slower in indexes using shelter costs, since shelter is a nontraded good. However, other components of the baskets, such as perishable foods, were nontraded across these cities before artificial refrigeration. Allen calculates the price of bread by combining grain prices with wages, which also introduces a nontraded component.

This study mainly focuses on the period of the price revolution, between 1503 and 1650. The time span is chosen to match the span in the classic study by Hamilton (1934), except that I begin in 1503 rather than 1501 to produce a balanced panel that includes Amsterdam and Augsburg.

The prices are recorded both in silver and in local currency. Studying the prices in silver makes them directly comparable across cities, while studying those in local currency allows one to see the effects of debasement or depreciation. Figure 1 shows log CPIs in local currency (in black) and in silver (in red). The time span in the figure is 1450–1765 (where data are available) to put the price revolution in context by including data before and after it. The general upward trend over this period of course constitutes the price revolution. Notice that for Valencia the inflation rate is lower both before and after 1503–1650, a feature not shared by all cities.

For Krakow prices are quoted only in silver. For Strasbourg the silver price does not fluctuate. For all other cities log prices in local currency (black line, right axis) and log prices in silver (red line, left axis) are shown on the same vertical scale specific to that city. Thus the overall inflation in local currency can be compared to that in silver. In each case the black line rises more than the red line. That difference reflects periodic depreciations or debasements of the local currency. I study the effects of these depreciations below in section 3.5.

There also are some not-so-gradual changes. For Augsburg, Strasbourg, and Vienna these spikes appear in prices in both local currency and in silver. For Augsburg the outliers appear as values in 1621–1623 and for Vienna 1621–1624. These are the early years of the Thirty Years War which brought cataclysmic events to both cities. For London, Munich, and Paris there are spikes in the value of the local currency relative to silver. For example, in London the price level appears unusually low in 1546–1551. In the Allen files this episode is due to a jump up in the price of silver, measured in pence/oz. His original source—Feaveryear (1931)—confirms these values, which occur at the time of the Tudor debasement. The debasement began in 1544, late in the reign of Henry VIII, and ended with the Elizabethan restoration of the coinage in 1560–1561. Gould (1970), Outhwaite (1982), and Challis (1989) provide the detailed history.

It is possible, though, that this spike in the price of silver in London is partly due to measurement error. The price of silver is in newly minted pence, but multiple coinages (and currencies) circulated and were not all debased. In other words, there were coins with the same face value that varied in silver content. To allow for the possibility that spikes like these are due to measurement error, I comment below on estimates of mean-reversion for these cities that are robust to outliers.

The vertical scales in figure 1 differ across cities, so that one can observe city-specific outliers like these. But the focus here is on *relative* price revolutions. To begin to see how these measures may be connected, I next construct log real exchange rates. Call Valencia city val and label any other city by j. Label the log price as $p_{j,t}$, measured in silver for comparability across cities. The log real exchange rate or relative price is:

$$q_{j,t} = p_{val,t} - p_{j,t}, \tag{1}$$

so that an increase is a real appreciation in Valencia relative to city j.

Figure 2 graphs $q_{j,t}$ for nine cities. Now the graphs apply to 1503–1765, a span which provides the largest balanced panel. Data for Munich end in 1765 and those for Paris end in 1786. The long time span is designed to enhance power in testing for unit roots in section 3.1 below. The real exchange rates for most (perhaps all) cities display no trend. That suggests there was a European-wide price revolution. The rest of the paper constructs statistics to provide details on how this unfolded.

3. Five Features of Real Exchange Rates

This section assesses five features of real exchange rates: (1) whether they are stationary (sometimes referred to as long-run PPP); (2) how quickly mean-reversion occurs; (3) whether mean-reversion is faster the nearer a city is to Valencia; (4) whether adjustment occurs in prices in Valencia or elsewhere; and (5) whether currency depreciations relative to silver play a role in the adjustment of real exchange rates. One of the goals is to compare these properties with those found in modern real exchange rates.

3.1. Real Exchange Rates are (Probably) Stationary

Tests for the stationarity of $q_{j,t}$ are sometimes referred to as tests of long-run purchasing power parity (PPP). But note that stationarity does not imply long-run equality of the price indexes in levels, only that there is reversion towards some mean. That is appropriate here too, for the consumption baskets of course differ across cities in the Allen data.

To test for unit roots in $q_{j,t}$ I calculate the *DF-GLSu* and Q_T statistics derived by Elliott (1999). These offer higher power than the traditional ADF test. They draw the initial value from the unconditional distribution rather than using a value of zero as Elliott-Rothenberg-Stock tests do. The lag length is selected by the BIC. I use the span 1503–1765, which is the longest span common to all cities. This span extends beyond the traditional span of the price revolution also to enhance test power. To make the same point in a different way, the long span allows for the possibility that the price level in Krakow, say, only slowly responds to shocks to the price level in Valencia, including those late in the period of the price revolution.

Table 2 contains the results. Evidence against the unit root null appears in the left tail of each distribution. For Antwerp and Krakow one statistic has a p-value below 10% and the other has a p-value below 5%. For both cities there is some real depreciation over this time span: Their inflation rates (in silver) tended to be below those in Valencia. For the other 7 cities both statistics have p-values below 1%. That suggests that there are not unit roots in these real exchange rates. Based on these results, I treat them as stationary and use standard tools of inference in studying $q_{j,t}$ below.

Historically, these findings suggest that the price revolution applied to all of these European cities. Had there been a city immune to this pattern, it would have experienced a trend real depreciation, but such a pattern is not evident from Table 2 or Figure 2.

In fact, Figure 2 provides an additional finding. Suppose one hypothesized that the price revolution first affected Valencia, and that its price level followed an S-shaped (logistic) pattern over time. Suppose further that one imagined other cities followed this same pattern but with a time delay. In that case their real exchange rates would jump up early in the 16th century and then gradually decline as their inflation rates first caught up to those in Valencia and then later exceeded them. But Figure 2 does not show such a general pattern in the real exchange rates.

3.2. Mean Reversion is Relatively Rapid

To characterize the mean reversion in each real exchange rate, I next zoom in on the span 1503–1650, as in Hamilton's (1934) study. I use this statistical model:

$$\Delta q_{j,t} = \alpha_j + \beta_j q_{j,t-1} + \epsilon_{j,t}. \tag{2}$$

To check on the validity of this first-order model, I also calculate the Ljung-Box test statistic for residual autocorrelation up to lag 4, as modifed by West and Cho (1995) to be robust to heteroskedasticity. The p-values show there is little evidence of remaining dynamics in $\hat{\epsilon}_{j,t}$.

Table 3 contains OLS estimates $\hat{\beta}_j$ and their HAC standard errors. All the estimates are negative and most are estimated quite precisely. In absolute value the smallest values are for Antwerp (-0.100) and Krakow (-0.117), where the unit root tests found the most persistence in $q_{j,t}$ or the least evidence of a shared price revolution. The highest values are for Paris (-0.358) and Amsterdam (-0.334). Overall, cheap (expensive) cities experience real appreciation (depreciation), restoring long-run PPP.

Recall from section 2 that there are outliers in the series particularly for Augsburg, London, and Vienna. For those cities, I also estimate equation (2) by LAD (least absolute deviations) which is less sensitive to these outliers. The LAD (OLS) values for $\hat{\beta}_j$ are for Augsburg -0.345 (-0.312), for London -0.194 (-0.294), and for Vienna -0.085 (-0.191). The LAD estimates are negative and comparable in scale to those from OLS, and the point estimate of mean reversion is the lowest for Vienna.

To study the heterogeneity in $\hat{\beta}_j$ across cities, I then estimate the restricted system with $\beta_j = \beta$:

$$\Delta q_{j,t} = \alpha_j + \beta q_{j,t-1} + \epsilon_{j,t}. \tag{3}$$

As with other studies of PPP with panel data, there is cross-city correlation in the residuals in part because of the common base city, Valencia. The cross-section dimension $J = 9$ is small relative to the time-series dimension $T = 148$, so estimation by Zellner's (1962) original SUR yields standard errors robust to this spatial dependence, as noted by Driscoll and Kraay (1998).

Estimation yields $\hat{\beta} = -0.194$ with standard error 0.016. A likelihood ratio test of the cross-city restrictions has a p-value less than 0.01, so that this pooled model cannot fully represent the adjustment overall. The next sub-section studies one possible pattern in the heterogeneity of adjustment speeds.

Meanwhile, from the restricted panel (3) the half-life of departures from PPP, denoted h, is given by:

$$\hat{h} = \frac{\ln(0.5)}{\ln(1 + \hat{\beta})},$$

This formula gives a value of 3.21 years with a standard deviation of 0.30 years. In equation (3) I have corrected standard errors rather than estimating by feasible GLS, so the choice of base city does effect the estimates $\hat{\beta}$ and $\hat{h}$. However, the estimated half life is very similar if I use as a base Amsterdam (3.06), Antwerp (3.51), Paris (3.20), or other cities.

To put this finding in context, I note that Rogoff's (1996) summary stands up after 30 years: Across many studies the half-life of PPP deviations is in the range of 3 to 5 years. This consensus is based both on the ongoing floating-rate era and on some historical studies with a long time spans. For example, Diebold, Husted, and Rush (1991) study 6 countries during the classical gold standard. They find that the half-life of departures from PPP is approximately 3 years. Lothian and Taylor (1996) study the sterling/dollar real exchange rate for 1791–1990 and the sterling/franc real exchange rate for 1803–1990 and find half-lives of 6 years and 3 years respectively.

To my knowledge, only one previous study estimates half-lives for the early modern period. Çürük, Karaman, and Yıldırım-Karaman (2022) aggregate data from Allen (2001) and other sources over 12 commodities and multiple cities to produce national indexes for Austria, the Dutch Republic, England, France, Poland, Portugal, and Spain for 1500–1800. Pooling across countries, they find a half-life of about 2 years. (Their method uses LPs with a variety of controls and so is comparable to the method in section 4 below.)

Two studies focus on differences in estimated half-lives across regimes. Taylor (2002) studies the persistence in real exchange rates for a panel of twenty countries over 1892–1996. Pooling across regimes and countries, he finds a half-life of 3.4 years. The estimated half-life is lower (1.6 years) for the Bretton Woods period and for the subsequent float (2.1 years). Bergin, Glick, and Wu (2014) study twenty industrialized countries (now excluding developing countries) over 1949–2010, and thus with more observations in the floating-rate era. They find a half-life under the Bretton Woods regime of 1.9 years and under the subsequent float of 3.6 years. Overall, then, the early modern adjustment half-lives found in Table 3 are comparable to those in recent data, though not quite as low as in the Bretton Woods era.

In the AR(1) model or its transformed version (2), the OLS estimator of persistence is biased down ($\hat{\beta}$ is biased away from zero), so OLS under-estimates persistence and the half-life. However, most other studies (including those surveyed by Rogoff) also use OLS, so this fact may not affect the comparison of half-lives. The bias is greatest for roots near unity and for small sample sizes, e.g. less than 100 observations. Rossi (2005) also shows that confidence intervals for the half-life can be very wide for roots near unity. But these conditions are unlikely to affect the findings here. The sample size is $T = 148$ and the pooled, point estimate of persistence is $1 + \hat{\beta} = 0.806$.

3.3. Mean Reversion Does Not Vary with Distance from Spain

The point estimates in Table 3 show that mean reversion is faster for Paris than for London or Vienna. This pattern suggests that the adjustment to long-run PPP may be related to distance from Valencia. It seems interesting to check for distance effects because the test of H_0: $\beta_j = \beta$ in Table 3 found evidence of heterogeneity in the responses.

To my knowledge, there are no good sources on early modern travel times or routes between these cities. The remarkable Viabundus project of Holterman et al. (2022) estimates some travel times and routes. But it includes only northern European cities and so far excludes almost all of those used in this study. I therefore measure distance by the great circle distance between the principal church or cathedral in each city, which plays the role of the city centre. Each of these church locations existed in 1503 at the start of the sample. (St Paul's Cathedral in London was built between 1675 and 1710 but it stands on the site of Old St Paul's.) Table 4 lists the church or cathedral in each city, its latitude and longtitude, and the great circle distance from Valencia Cathedral. Distance is rounded to the nearest kilometre. For simple visualization, Figure 3 plots the longitude and latitude of each city, along with its label from Table 1.

Call d_j the distance to city j from Valencia. The statistical model then is:

$$\Delta q_{j,t} = \alpha_j + (\beta_0 + \beta_d d_j) q_{j,t-1} + \epsilon_{j,t}, \tag{4}$$

so that the speed of adjustment can vary with distance. Two issues arise with this setup. First, the great circle distance of course may measure the economic distance or travel cost with error. Other measures of the economic distance between cities may come to mind, for example using modern surface travel routes that reflect geographical barriers. Second, we have no reason to believe the effect is linear in distance. I thus repeat the estimation with distance d_j replaced by its rank r_j, shown in the last column of Table 4. That replacement may resolve the first issue and does resolve the second one, since rank regression is linear for any monotone function.

The aim here is to see whether the adjustment speed declines as distance increases and to report the scale of that effect. However, no such pattern appears. In each model the coefficients $\hat{\beta}_d$ on the interaction terms are positive, so that the speed of mean-reversion declines as distance increases. But the coefficients are very small and are imprecisely estimated.

3.4. Most Adjustment is by Non-Spanish Prices

I next report the role played by the local inflation rate in adjustment to long-run PPP. Denote by P the linear projection operator. Then from the definition of $q_{j,t}$ (1):

$$P[\Delta q_{j,t}|q_{j,t-1}] = P[\Delta p_{val,t}|q_{j,t-1}] - P[\Delta p_{j,t}|q_{j,t-1}].$$

Thus the mean reversion in the real exchange rate naturally arises from the difference between the responses of the two rates of change of prices. The last term is estimated with:

$$\Delta p_{j,t} = \gamma_j + \omega_j q_{j,t-1} + \epsilon_{j,t}. \tag{5}$$

The final column of Table 3 contains estimates $\hat{\omega}_j$ and their standard errors for each city and for the system restricted to have a common value ω. The coefficients $\hat{\omega}_j$ are all positive so that cities in which $p_{j,t-1} < p_{val,t-1}$ tend to then experience price increases. The pooled version of the price-change response equation (5) gives $\hat{\omega} = 0.137$ with standard error 0.029. However the cross-equation restrictions again are rejected at conventional significance levels, as the p-value of the likelihood ratio test is less than 0.01.

When I mildly unrestrict the projection (5) by entering $p_{val,t-1}$ and $p_{j,t-1}$ separately, I find that both are useful predictors and that they enter with coefficients that are approximately equal and opposite (and hence the restriction to combining them as $q_{j,t-1}$ appears to hold). Thus the coefficient $\hat{\omega}_j$ is not simply capturing mean reversion in the inflation rate in each city. One might also wonder whether the $q_{j,t-1}$ term matters only when it is positive, in that city-j inflation catches up to Valencia inflation as Spain leads the price revolution. But a formal test with a dummy variable distinguishing positive from negative values of $q_{j,t-1}$ shows that is not the case. Inflation does tend to be high after years in which $p_{j,t}$ is less than $p_{val,t}$ but it also tends to be low in the opposite case.

The percentage share of the adjustment contributed by local price changes is $100 \times -(\hat{\omega}_j/\hat{\beta}_j)$. That ratio is 38% for Krakow but is greater than 50% for every other city, with a maximum value of 86% for Vienna. In the pooled version (which implicitly weights cities by goodness-of-fit) the ratio is 71%. Thus the majority of price adjustment that restores long-run PPP is done by non-Spanish prices.

3.5. Real Exchange Rates Evolve as if in a Silver Standard

I next study what role adjustments in nominal exchange rates played in the adjustment towards PPP. This too is a standard question asked in modern data. Constructing nominal exchange rates begins with Allen's (2001) data on values of local currencies in terms of silver, which he collected from a wide range of sources. These series align with the cities and time periods studied so far. The currencies are guilders for Amsterdam and Antwerp, pfennings for Augsburg, pence for London, denars for Munich, livres tournois for Paris, and kreuzer for Vienna along with pence for Valencia (as listed in Table 1). I omit Krakow and Strasbourg in this section because their prices are quoted only in silver and not in local currency.

Allen records the value of local currency in terms of silver, for example in grams per guilder for Amsterdam. These values tend to fall over time, reflecting debasements for example. Denote the log value of the currency in silver by $a_{j,t}$. Thus the analogue to a log nominal exchange rate is:

$$s_{j,t} = a_{val,t} - a_{j,t}. \tag{6}$$

An increase in $s_{j,t}$ is a nominal depreciation in city j, just as an increase in $q_{j,t}$ is a real depreciation. If Amsterdam has faster debasements than does Valencia then the guilder will depreciate.

This is a synthetic nominal exchange rate for two reasons. First, foreign currencies circulated in most cities and so could be exchanged for local ones. For example, through much of this period the Castilian real was widely used in Valencia. Second, actual, larger-scale transactions in foreign exchange were conducted through high-denomination gold coins or bills of exchange, which could have a maturity of a month or more. (Brzezinski et al. (2024) provide some prices of bills of exchange for a number of the cities in this panel, but they begin in 1575.)

Figure 1 showed that prices in local currency rose more than did prices in silver, which means that the values of the local currencies fell in terms of silver over time. Table 5 gives cumulative, percent decreases in these currency values for each city over 1503–1650. For example, in Valencia the value of the local currency in silver fell by 4.3% in this period. All other cities had greater decreases (larger falls in $a_{j,t}$). Thus they all experienced nominal depreciations relative to Valencia, as measured by the synthetic exchange rates.

Feaveryear (1931) noted that physical wear on coins gradually eroded their silver content so that the value of local currency in silver tended to drift down (i.e. there were gradual debasements). But this trend would lead to a trend in the nominal exchange rate only if there were *differences* in the rate of physical wear across cities.

In contrast, periodic debasements directed by monetary authorities led to irregular jumps down in $a_{j,t}$ which are thus reflected in $s_{j,t}$. Debasements occurred as a sovereign reduced the amount of silver in a coin while minting new silver supplies or reminting older silver. As a result the value of those coins measured in silver fell. Fiat money was not yet in use in this period. During the 1600s coins combining copper and silver (called vellon) were issued in Castile however. In that case reducing the share of silver in those coins also counts as a depreciation of the local currency.

Not all of the paths are monotone though. As noted in section 2, London, Munich, and Paris exhibit sharp, V-shaped troughs in the value of the local currency: faster debasements followed by restorations. For London, the trough in 1551 marks the end of the Great Debasement (1544–1551) begun under Henry VIII and continued then reversed under Edward VI. Feaveryear (1931), Gould (1970), and Outhwaite (1982) provide data and histories of this episode. For Munich, the trough is in 1622, early in the Thirty Years War. Kindleberger (1991) argued that sovereigns anticipated the war and accelerated minting and clipping to accumulate seigniorage. He suggested that the debasements then spread to other cities (especially in the Habsburg Empire) via Gresham's Law, because many currencies circulated simultaneously, especially in smaller states. However, Figure 1 does not show similar changes in currency values for Antwerp or Vienna, though prices in both silver and local currency spiked upwards in Augsburg. For Paris the trough is in 1720, which marked the end of John Law's System and its associated paper money. Velde (2007) provides a concise summary of this dramatic episode.

Karaman, Pamuk, and Yıldırım-Karaman (2020) provide a complete history of European debasements in the early modern period. They find that debasements were predicted by political events, such as wars, and not by economic factors involving the prices of commodities or precious metals. In other words, debasements were designed to raise revenue not to deliver real depreciations. Palma (2022, p 1609) finds that debasements were not predicted by inflows of precious metals to Europe.

Modern studies of international macroeconomics often find a correlation between the nominal and real exchange rates. To allow for possible non-stationarity in the nominal rates I calculate the correlations between changes in the log real and nominal exchange rates, $\Delta q_{j,t}$ and $\Delta s_{j,t}$. London is the only city for which this correlation is large: The value is 0.71 over 1503–1650. For the other cities these values range from -0.5 to 0.24. Overall, then, they are lower than those for modern, floating exchange-rate regimes.

I next explore formally whether these movements in nominal exchange rates played a role in the adjustment of real exchange rates. The statistical model is:

$$s_{j,t+h} - s_{j,t} = \xi_{j,h} + \theta_{j,h} q_{j,t} + \epsilon_{j,t+h}, \tag{7}$$

where the new subscript h measures the horizon in years. From section 3.2 we know that cities with low prices tend to experience real appreciations ($\hat{\beta}_j < 0$). We now see what part of this adjustment occurs as a nominal appreciation. The change will be in that direction for city j if $\theta_{j,h} < 0$. This is the Big Mac hypothesis long *avant la lettre*.

This estimation is inspired by two previous studies, both of which compare the response of the nominal exchange rate to the lagged real exchange rate with the overall response of the real exchange rate itself. Cheung, Lai, and Bergman (2004) study France, Germany, Italy, Japan, and the UK relative to the US for 1973 to 1998. They find that 60–90% of adjustment to PPP occurs through the nominal exchange rate. Eichenbaum, Johannsen, and Rebelo (2021) estimate equation (7) for a range of countries in data up to 2008. They focus on the fact that for inflation-targeting countries $\hat{\theta}_j$ is negative, rises in absolute value as the horizon rises, and accounts for most of the adjustment in the real exchange rate. That means that the real exchange rate forecasts changes in the nominal exchange rate. In the context here, that means that relative debasements or depreciations are predictable.

In the early modern data in this study, for individual cities, the coefficient $\hat{\theta}_{j,h}$ is generally negative, though estimated with varying precision. To report on the role played by adjustment in the nominal exchange rate I therefore estimate the system:

$$s_{j,t+h} - s_{j,t} = \xi_{j,h} + \theta_h q_{j,t} + \epsilon_{j,t+h} \tag{8}$$

so that $\hat{\theta}_h$ is the pooled estimate of the response at horizon h. To assess the scale of $\hat{\theta}_h$ I also estimate the pooled version of the original mean-reversion equation (3) but now at several horizons:

$$q_{j,t+h} - q_{j,t} = \alpha_{j,h} + \beta_h q_{j,t} + \epsilon_{j,t+h}, \tag{9}$$

so that $\hat{\beta}_h$ measures the total mean-reversion in the real exchange rates at horizon h.

Table 6 contains the results, with horizons of $h = 1, 3, 5$ years. In the nominal exchange rate equation (8) the coefficients $\hat{\theta}_h$ take small negative values. The impacts become larger in absolute value as the horizon increases and are estimated with some precision. For this group of cities then, the synthetic, nominal exchange rate plays a role in adjustment towards PPP. On average, a city with low relative prices will tend to experience a small, nominal appreciation in subsequent years.

The adjacent column of Table 6 reports results from pooled estimation of the adjustment in the real exchange rate (9) for the same group of cities. The estimates β_j are negative, tending to restore PPP, and they too rise in scale as the horizon h increases. The responses are estimated with more precision than those of the nominal exchange rate. Most notably, the ratio $\hat{\theta}_h/\hat{\beta}_h$ takes values 0.9%, 1.8%, and 3.2% across the horizons $h = 1, 3, 5$, for the span 1503–1650. Thus the response of the nominal exchange rate is very small as a share of the adjustment in the real exchange rate.

Using the study of Eichenbaum, Johannsen, and Rebelo as a benchmark, the results in Table 6 are most similar to their modern findings for countries in the Euro area relative to Germany. Given this similarity, one might wonder whether projections (8) and (9) really can discriminate among exchange-rate regimes, in other words about test power. But those authors find that there is power. For example, for modern, inflation-targetting countries a majority of the adjustment is done by the nominal exchange rate rather than by the inflation rates. And there is no clear pattern to $\{\hat{\theta}_h, \hat{\beta}_h\}$ for a range of countries with floating exchange rates that do not target inflation. They thus conclude that these statistics do reflect the monetary policy regime.

As the introduction mentioned, the goal here is to uncover some characteristics of the early modern economy of Europe by studying the price revolution. In this case, the statistical evidence suggests the city-level price indexes evolved as if the cities were in a fixed exchange rate system or monetary union, and specifically a silver standard.

Table 5 showed that all of the cities experienced nominal depreciations relative to Valencia, ranging from $54.6 - 4.3 = 50.3\%$ for London to $236.7 - 4.3 = 232.4\%$ for Amsterdam, over the 1503–1650 period. The results of this section show that these generally were not associated with movements in the real exchange rate but instead were largely offset by increases in prices, when measured in local currency.

What do I mean by saying that a silver standard applied? Clearly all of these jurisdictions of course were on silver during this period and just as clearly they almost all experienced changes (usually decreases) in the value of local currency in terms of silver. Here I mean specifically that the depreciations and debasements did not appear to affect real exchange rates. This conclusion is complementary to that of Karaman, Pamuk, and Yıldırım-Karaman (2020) who show there is no causation in the other direction, from economic factors like prices to debasements.

4. Diffusion of Spanish Inflation

As the introduction noted, historians have long concluded that the price revolution arose from imports of precious metals to Spain. This section studies the diffusion of inflation from Spain by ordering inflation in Spanish cities first in a system of local projections (LPs). That allows one to study the impact on other cities, without a complete identification of their structural innovations. The idea is simply to assume that an important shock originated in Spain and affected inflation there first. Inflation then radiated out from Spain, possibly via (a) the use of Spanish coins, (a) the spread and minting of silver or (b) a demand shock as output rose in Spain and other countries, as documented by Palma (2022).

I consider two origin cities, Valencia and Madrid. Treating Valencia as the city of origin allows a panel with a long time series dimension, including the entire sixteenth century. Treating Madrid as the city of origin locates it in Castile, the centre of Spansih silver imports.

So far I have not found a valid instrument that would allow estimation by LP-IV. A measure of metal inflows to Spain is *not* an instrument for inflation in Valencia or Madrid, at least in this context. Spanish coins minted in America circulated widely in Europe and so this inflow may have directly affected $p_{j,t}$ in non-Spanish cities. Moreover, Palma (2022) argues that Spanish silver was minted in other European countries. For example, chemical tests show English coins contained silver from the Americas. Thus this variable does not satisfy the exclusion restriction of instrumental-variables design.

For detailed evidence on the effects of American metal, treated instead as an exogenous variable, I refer the reader to Palma (2022) and Brezinski et al. (2024). The former finds a positive effect on prices, pooled across six countries. The latter find insignificant effects of such imports on the aggregate Spanish price level.

Let the subscript *esp* refer to either Valencia (*val*) or Madrid (*mad*). To measure diffusion I first study the real exchange rate, for city j and horizon h:

$$q_{j,t+h} - q_{j,t-1} = \alpha_{q,j,h} + \rho_h \Delta p_{esp,t} + \sum_{i=1}^{2} \kappa_{q,j,h,i} \Delta p_{esp,t-i} + \sum_{i=1}^{3} \eta_{q,j,h,i} q_{j,t-i} + \epsilon_{q,j,t+h}. \quad (10)$$

I then also record the impact on city-specific h-step inflation:

$$p_{j,t+h} - p_{j,t-1} = \alpha_{p,j,h} + \lambda_h \Delta p_{esp,t} + \sum_{i=1}^{2} \kappa_{p,j,h,i} \Delta p_{esp,t-i} + \sum_{i=1}^{3} \eta_{p,j,h,i} q_{j,t-i} + \epsilon_{p,j,t+h}. \quad (11)$$

In the rest of the paper I report results based only on pooled estimates, using data from all cities, for two reasons. First, LPs like systems (10) and (11) involve more parameters than the tests in section 3 do. I base inference on estimates restricted to hold over the $J \times T$ city/year observations to aid statistical precision. Second, although Figures 1 and 2 and Tables 2 and 3 shows differences across cities in the properties of real exchange rates, I have not found evidence that these are related to distance from Spain. Figure 3 shows that a number of cities are approximately the same distance from Valencia and section 3.3 showed no distance effect in the speed of mean-reversion in $q_{j,t}$. However, I report tests of the hypothesis of common parameters across cities so that the reader can see where further investigation is needed.

These LP systems (10) and (11) involve long differences on the left-hand side, as recommended by Jordá and Taylor (2025) and Piger and Stockwell (2025). The intercepts $\alpha_{q,j,h}$ and $\alpha_{p,j,h}$ are specific to the city and horizon. The right-hand side includes lagged real exchange rates for city j, as we know from sub-section 3.4 that these help forecast inflation. Notice that the parameters on lagged variables can vary by horizon, lag, and city (as subscripted above). At horizon $h = 0$ I calculated $Q(4)$, the Ljung-Box test statistic for residual autocorrelation, as modifed by West and Cho (1995) to be robust to heteroskedasticity. The p-values show there is little evidence of remaining dynamics in the residuals. At further horizons, local projection h steps ahead induces a moving average error of order $h - 1$. Systems (10) and (11) also include extra lags of $\Delta p_{esp,t-i}$ and $q_{j,t-i}$ so that standard errors can be calculated as heteroskedasticity-consistent (HC) with lag augmentation. Jordá and Taylor (2025) and Montiel Olea, Plagborg-Møller, Qian, and Wolf (2025) recommend calculating standard errors in this way instead of using HAC standard errors.

Montiel Olea, Plagborg-Møller, Qian, and Wolf (2025, Lesson 1) note that including the control variables isolates a shock to $\Delta p_{esp,t}$ relative to this information set, by the Frisch-Waugh-Lovell Theorem. The sequences $\{\hat{\rho}_h\}$ and $\{\hat{\lambda}_h\}$ thus give the impulse response functions (IRFs) for real exchange rates and inflation respectively in response to a shock to inflation in either Valencia or Madrid. They are common across cities, to aid precision. But at each horizon I also calculate likelihood ratio tests of the hypotheses that these parameters are equal across cities: $H_0 : \rho_{h,j} = \rho_h$ and $H_0 : \lambda_{h,j} = \lambda_h$.

4.1 Valencia Inflation Shock: 1503–1650

Valencia (in Aragon) was not as central to the expansion of the Spanish money supply via imports of silver as were cities like Seville and Toledo (in Castile). Hamilton (1934, p 128) reported that silver arrived in Seville and was minted in Castile. But he also noted that Castilian reals were the main currency in Valencia during the 1500s. Moreover, treating Valencia as the origin city allows one to study long time spans. There are 148 annual observations in the central period of the price revolution (1503–1650). I base inference on estimates restricted to hold over the 9×148 city/year observations to aid statistical precision.

Table 7 contains the results. The second and fourth columns trace out the IRFs for the real exchange rates and inflation rates respectively. In point estimates, a 1% shock to Valencia inflation leads to a 0.84% real appreciation in Valencia and 0.16% increase in inflation in other cities within the year. These values then decline over time. Note that $\hat{\rho}_0 + \hat{\lambda}_0 = 1$ automatically. But these coefficients do not add up to 1 at $h > 1$ because of the response of $\Delta p_{esp,t+h}$ to the shock.

The test statistics for common impacts ρ_0 and λ_0 across cities have very low p-values, so that the pooling is unlikely to hold. In that case these two impact coefficients are best thought of as an average over the cities (weighted by the fit for each city) rather than being representative of each city. But at each further horizon the p-values are much larger, suggesting that the pooling is valid.

The IRF coefficients $\{\hat{\rho}_h\}$ in Table 7 show the persistence in the real exchange rates now conditional on a shock to inflation in Valencia. That pattern is roughly comparable to the unconditional persistence reported in the lower panel of Table 3, where the first-order autocorrelation coefficient for $\{q_{j,t}\}$ is approximately 0.8.

The last panel of Figure 1 shows that the inflation rate (measured in silver) fell in Valencia after 1600 and that the price level itself fell after 1650. So it is interesting to see whether diffusion continued throughout the 17th century. Updating Table 7 to apply to 1503–1700 (not shown) suggests that it did continue: The results are largely unchanged.

I next see whether some of the heterogeneity in the initial impact of a Valencia inflation shock can be statistically explained by distance from Valencia. It might seem plausible that the impacts would vary from city to city while the effects were similar after the passage of several years, hence at larger horizons h. To assess this possibility I augment the impact parameters at $h = 0$ by writing them as linear in distance d_j or rank distance r_j, as in sub-section 3.3. However, there is no evidence that the resulting interaction terms play a role, as their p-values are large.

I also looked at recursive chains or rays emanating from Valencia, such as Valencia → Strasbourg → Vienna and Valencia → Paris → London, in which the inflation rate in the third city was also influenced by that in the second city. These chains did not reveal a consistent pattern. Finally, I also distinguished between cities that were governed by the Habsburgs (Antwerp, Vienna, and Valencia after 1516), imperial cities (Augsburg, Munich, Strasbourg), and cities with no Habsburg connection (Krakow, London, Paris, and Amsterdam except during 1556–1581). Those distinctions did not explain the heterogeneity either.

4.2 Madrid Inflation Shock: 1551–1650

Treating the inflation shock as originating in Madrid is appealing because Castile was more central to the Spanish economy. However, Madrid was a relatively small city until Philip II made it the capital in 1561. In the Allen data its price index begins in 1551, giving $T = 100$ observations until 1650. I also add Valencia to the panel, treating it like other cities impacted by the inflation shock, so there are now 10 such cities. And real exchange rates $q_{j,t}$ are defined with Madrid as the base city.

Table 8 contains the results of estimating the LP systems (10) and (11). The results for persistence in real exchange rates are generally similar to those in Table 7. An inflation shock leads to a real appreciation in Madrid (with an impact at $h = 0$ of 0.83) and to inflation in other cities (with an impact of 0.17). But there are two notable differences. First, there now is a hump-shaped inflation response, with the impact on inflation in other cities peaking at 0.29 at horizon $h = 2$ years. Second, there now is much less evidence of heterogeneity on impact at $h = 0$: The p-values are 0.16 for both real exchange rates and inflation rates. This finding does not stem from there being less precision in the impact coefficients $\hat{\rho}_0$ and $\hat{\lambda}_0$, which are estimated quite precisely just as in Table 7. Thus the overall message from treating Madrid as the origin city is that there was a gradual but uniform impact on real exchange rates and a delayed impact on inflation rates. (Again, the difference reflecting the response of Madrid inflation to its own shock.)

5. Conclusion

The European price revolution has long been viewed as an important episode in monetary history. This paper applies modern statistical tests to a series of questions about how it unfolded differentially across cities. There are five main findings:

1. Real exchange rates are stationary, consistent with long-run PPP. That statistical evidence suggests or confirms that the price revolution occurred across European cities, possibly as far east as Krakow.

2. The half-life of departures from long-run PPP is 3.21 years with a standard error of 0.30 years. This speed is comparable to those in many modern studies of exchange rates.

3. Periodic debasements did not appear to affect real exchange rates, so European price levels continued to be linked as if on a silver standard.

4. Shocks to inflation originating in Valencia or Madrid led to real appreciation there but then gradually diffused across cities.

5. There is no evidence of distance effects relative to Valencia, either in the speed of adjustment of real exchange rates or in the diffusion of inflation.

Inspecting this list leads one to wonder what other common shocks (possibly originating outside Spain) may have affected these inflation rates. Meanwhile, though, these characteristics for the 16th and 17th centuries—long-run PPP, European currencies on a common (though commodity) standard, inflation gradually diffusing across countries—are after all not very different from those observed for much of the 20th and 21st centuries.

References

Allen, Robert C. (2001) The great divergence in European wages and prices from the Middle Ages to the First World War. *Explorations in Economic History* 38, 411–447.

Bergin, Paul R., Reuven Glick, and Jyh-Lin Wu (2014) Mussa redux and conditional PPP. *Journal of Monetary Economics* 68, 101-114.

Braudel, Fernand P. and Frank Spooner (1967) Prices in Europe from 1450 to 1750. *The Cambridge Economic History of Europe* 4, 378–486.

Brzezinski, Adam, Yao Chen, Nuno Palma, and Felix Ward (2024) The vagaries of the sea: New evidence on the real effects of money from maritime disasters in the Spanish empire. *Review of Economics and Statistics* 106, 1220–1235.

Challis, C. E. (1989) *Currency and the Economy in Tudor and Early Stuart England.* London: The Historical Association.

Cheung, Yin-Wong, Kon S. Lai, and Michael Bergman (2004) Dissecting the PPP puzzle: the unconventional roles of nominal exchange rate and price adjustments. *Journal of International Economics* 64, 135–150.

Çürük, Malik, K. Kıvanç Karaman, and Seçil Yıldırım-Karaman (2022) Real exchange rate persistence under the silver standard over three centuries. mimeo

Diebold, Francis X., Steven Husted, and Mark Rush (1991) Real exchange rates under the gold standard. *Journal of Political Economy* 99, 1252–1271.

Drelichman, Mauricio and David González Agudo (2015) Housing and the cost of living in early modern Toledo. *Explorations in Economic History* 54, 27–47.

Driscoll, John C. and Aart C. Kraay (1998) Consistent covariance matrix estimation with spatially dependent panel data. *Review of Economics and Statistics* 80, 549–560.

Eichenbaum, Martin S., Benjamin K. Johannsen, and Sergio T. Rebelo (2021) Monetary policy and the predictability of nominal exchange rates. *Review of Economic Studies* 88, 192-228.

Elliott, Graham (1999) Efficient tests for a unit root when the initial observation is drawn from its unconditional distribution. *International Economic Review* 40, 767–783.

Feavearyear, A. E. (1931) *The Pound Sterling: A History of English Money.* Oxford: Clarendon Press.

González Agudo, David (2019) Prices in Toldeo (Spain). *Social Science History* 43, 269–295.

Gould, J. D. (1970) *The Great Debasement.* Oxford: Oxford University Press.

Hamilton, Earl J. (1934) *American Treasure and the Price Revolution in Spain, 1501–1650.* Cambridge, MA: Harvard University Press.

Hamilton, Earl J. (1947) *War and Prices in Spain 1651–1800.* Cambridge, MA: Harvard University Press.

Holterman, Bart, A.B. Maartje, Kasper H. Andersen, Maria C. Dengg, and Niels Petersen, (2022) Viabundus: Map of premodern European transport and mobility. *Research Data Journal for the Humanities and Social Sciences* 7, 1–13.

Jordà, Òscar and Alan M. Taylor (2025) Local projections. *Journal of Economic Literature* 63, 59–110.

Karaman, K. Kıvanç, Şevket Pamuk, and Seçil Yıldırım-Karaman (2020) Money and monetary stability in Europe, 1300–1914. *Journal of Monetary Economics* 115, 279–300.

Kindleberger, Charles P. (1991) The economic crisis of 1619 to 1623. *Journal of Economic History* 51, 149–175.

Lothian, James R. and Mark P. Taylor (1996) Real exchange rate behavior: The recent float from the perspective of the past two centuries. *Journal of Political Economy* 104, 488–509.

Montiel Olea, José Luis, Mikkel Plagborg-Møller, Eric Qian, and Christian K. Wolf (2025) Local projections or VARs? A primer for macroeconomists. Chapter 3 in *NBER Macroeconomics Annual 2025* eds. John V. Leahy and Valerie A. Ramey.

Munro, John (2008) Price Revolution. In: Durlauf, S.N., Blume, L.E. (eds) *The New Palgrave Dictionary of Economics.* Palgrave Macmillan: London.

Outhwaite, R. B. (1982) *Inflation in Tudor and Early Stuart England,* 2nd ed. London: Macmillan.

Palma, Nuno (2020) American precious metals and their consequences for early modern Europe. In: *Handbook of the History of Money and Currency*, edited by Stefano Battilossi, Youssef Cassis, and Kazuhiko Yago. Springer.

Palma, Nuno (2022) The real effects of monetary expansion: Evidence from a large-scale historical experiment. *Review of Economic Studies* 89, 1593–1627.

Piger, Jeremy and Thomas Stockwell (2025) Differences from differencing: Should local projections with observed shocks be estimated in levels or differences? *Journal of Applied Econometrics* `https://doi.org/10.1002/jae.70003`

Rogoff, Kenneth (1996) The purchasing power parity puzzle. *Journal of Economic Literature* 34, 647–668.

Rossi, Barbara (2005) Confidence intervals for half-life deviations from purchasing power parity. *Journal of Business and Economic Statistics* 23, 432–442.

Taylor, Alan M. (2002) A century of purchasing-power parity. *Review of Economics and Statistics* 84, 139–150.

Velde, François R. (2007) John Law's system. *American Economic Review (P)* 97, 276–279.

West, Kenneth D. and Dongchul Cho (1995) The predictive ability of several models of exchange rate volatility. *Journal of Econometrics* 69, 367–391.

Zellner, Arnold (1962) An efficient method of estimating seemingly unrelated regressions and tests of aggregation bias. *Journal of the American Statistical Association* 57, 348–368.

Table 1: Cities and Time Spans

City	Label	Time Span	Currency
Amsterdam	ams	1500–1910	guilder
Antwerp	ant	1399–1913	guilder
Augsburg	aug	1502–1800	pfennig
Krakow	kra	1409–1795	grosz
London	lon	1264–1913	pence
Madrid	mad	1551–1913	maravedi
Munich	mun	1427–1765	denar
Paris	par	1431–1786	livre tournois
Strasbourg	str	1386–1875	franc
Valencia	val	1413–1785	pence
Vienna	vie	1440–1800	kreutzer

Notes: Consumer price index spans and currency units from Allen (2001).

Table 2: Unit Root Tests in Log Real Exchange Rates

$$q_{j,t} = p_{val,t} - p_{j,t}$$

City	Q_T	$DF\text{-}GLSu$
Amsterdam	1.128	-5.812
Antwerp	4.328	-2.565
Augsburg	0.879	-5.147
Krakow	4.931	-2.946
London	1.837	-3.919
Munich	1.086	-4.853
Paris	0.838	-7.149
Strasbourg	1.249	-5.107
Vienna	1.090	-5.411

Notes: The table shows Elliott (1999) $DF{-}GLSu$ and Q_T test statistics for a unit root in the log real exchange rate of city j relative to Valencia. The lag length is selected by the BIC. A constant is included. The lag lengths, test statistics, and critical values are obtained using the RATS-Estima procedure erstest.src. For this table the time span is 1503–1765 to allow the longest balanced panel and enhance test power. Evidence against the unit root appears in the left tail of each distribution. For Q_T some asymptotic critical values are 3.06 (1%), 3.80 (2.5%), 4.65 (5%), and 5.94 (10%). For $DF{-}GLSu$ these are -3.28 (1%), -2.98 (2.5%), -2.73 (5%), and -2.46 (10%). For Antwerp and Krakow one statistic has a p-value below 10% and the other has a p-value below 5%. For the other 7 cities both statistics have p-values below 1%.

Table 3: Mean Reversion and Home Inflation

$$q_{j,t} \equiv p_{val,t} - p_{j,t}$$
$$\Delta q_{j,t} = \alpha_j + \beta_j q_{j,t-1} + \epsilon_{j,t}$$
$$\Delta p_{j,t} = \gamma_j + \omega_j q_{j,t-1} + \epsilon_{j,t}$$

City	$\hat{\beta}_j$ (se)	$\hat{\omega}_j$ (se)
Amsterdam	-0.334	0.210
	(0.090)	(0.109)
Antwerp	-0.100	0.056
	(0.039)	(0.032)
Augsburg	-0.312	0.224
	(0.161)	(0.169)
Krakow	-0.117	0.044
	(0.035)	(0.020)
London	-0.294	0.247
	(0.149)	(0.159)
Munich	-0.263	0.151
	(0.059)	(0.054)
Paris	-0.358	0.224
	(0.079)	(0.074)
Strasbourg	-0.207	0.129
	(0.064)	(0.060)
Vienna	-0.191	0.164
	(0.156)	(0.186)
Pooled	-0.194	0.137
	(0.016)	(0.029)
$\chi^2(8)$	24.62	28.02
(p)	(0.002)	(0.001)

Notes: Standard errors are HAC. There are 148 observations from 1503 to 1650. $\chi^2(8)$ is the likelihood ratio test statistic for the cross-equation restrictions $\beta_j = \beta$ or $\omega_j = \omega$ in the pooled models.

Table 4: Cathedral Locations and Distances

City	Cathedral/Church	Latitude	Longtitude	Distance d_j	Rank r_j
Amsterdam	Oude Kerke	52.374°N	4.898°E	1490	7
Antwerp	Our Lady	51.220°N	4.402°E	1357	5
Augsburg	Dom Maria	48.373°N	10.897°E	1337	4
Krakow	Wawel	50.055°N	19.936°E	1981	9
London	St. Paul's	51.514°N	0.098°W	1337	3
Munich	Frauenkirche	48.139°N	11.573°E	1358	6
Paris	Notre-Dame	48.853°N	2.350°E	1064	1
Strasbourg	Notre-Dame	48.582°N	7.751°E	1202	2
Valencia	St. Mary's	39.475°N	0.376°W	—	—
Vienna	St. Stephen's	48.208°N	16.374°E	1654	8

Notes: Latitude and longtitude are in decimal degrees. Distance is the great circle distance in kilometres from Valencia. It is calculated using the Vincenty formula from Adam Schneider's `gpsvisualizer.com/calculator`.

Table 5. Local Currency Values in Terms of Silver
Cumulative Decreases

City	% Decrease
Amsterdam	-236.7
Antwerp	-93.1
Augsburg	-81.1
London	-54.6
Munich	-73.4
Paris	-142.2
Valencia	-4.3
Vienna	-75.4

Entries are cumulative percent decreases in the value of local currency in terms of silver (i.e. depreciations) for 1503–1650 from Allen (2001). Notice that the rate is least for Valencia, so that there are nominal depreciations relative to Valencia as measured by synthetic nominal exchange rates.

Table 6. Adjustment in Nominal Exchange Rates

$$s_{j,t+h} - s_{j,t} = \xi_{j,h} + \theta_h q_{j,t} + \epsilon_{t+h}$$
$$q_{j,t+h} - q_{j,t} = \alpha_{j,h} + \beta_h q_{j,t} + \epsilon_{t+h}$$

Horizon	$\hat{\theta}_h$ (se)	$\hat{\beta}_h$ (se)
1	-0.002	-0.214
	(0.001)	(0.043)
3	-0.009	-0.490
	(0.003)	(0.050)
5	-0.020	-0.630
	(0.005)	(0.044)

Notes: h is the horizon in years. Standard errors are HAC. Krakow and Strasbourg are omitted because the Allen data do not show variation in their silver prices. There are 148 observations from 1503 to 1650.

Table 7. IRFs for Real Exchange Rates and Inflation Valencia Inflation Shock: 1503–1650

	$q_{j,t+h} - q_{j,t-1}$		$p_{j,t+h} - p_{j,t-1}$	
Horizon	$\hat{\rho}_h$ (se)	$\chi^2(8)$ (p)	$\hat{\lambda}_h$ (se)	$\chi^2(8)$ (p)
0	0.84 (0.05)	25.5 (0.00)	0.16 (0.04)	25.6 (0.00)
1	0.54 (0.10)	11.0 (0.20)	0.11 (0.06)	11.2 (0.19)
2	0.31 (0.09)	4.3 (0.82)	0.10 (0.06)	4.4 (0.82)
3	0.22 (0.10)	6.0 (0.64)	0.08 (0.06)	6.6 (0.58)
4	0.13 (0.09)	9.7 (0.29)	0.11 (0.06)	10.2 (0.25)
5	0.17 (0.09)	5.0 (0.76)	0.11 (0.07)	5.3 (0.74)

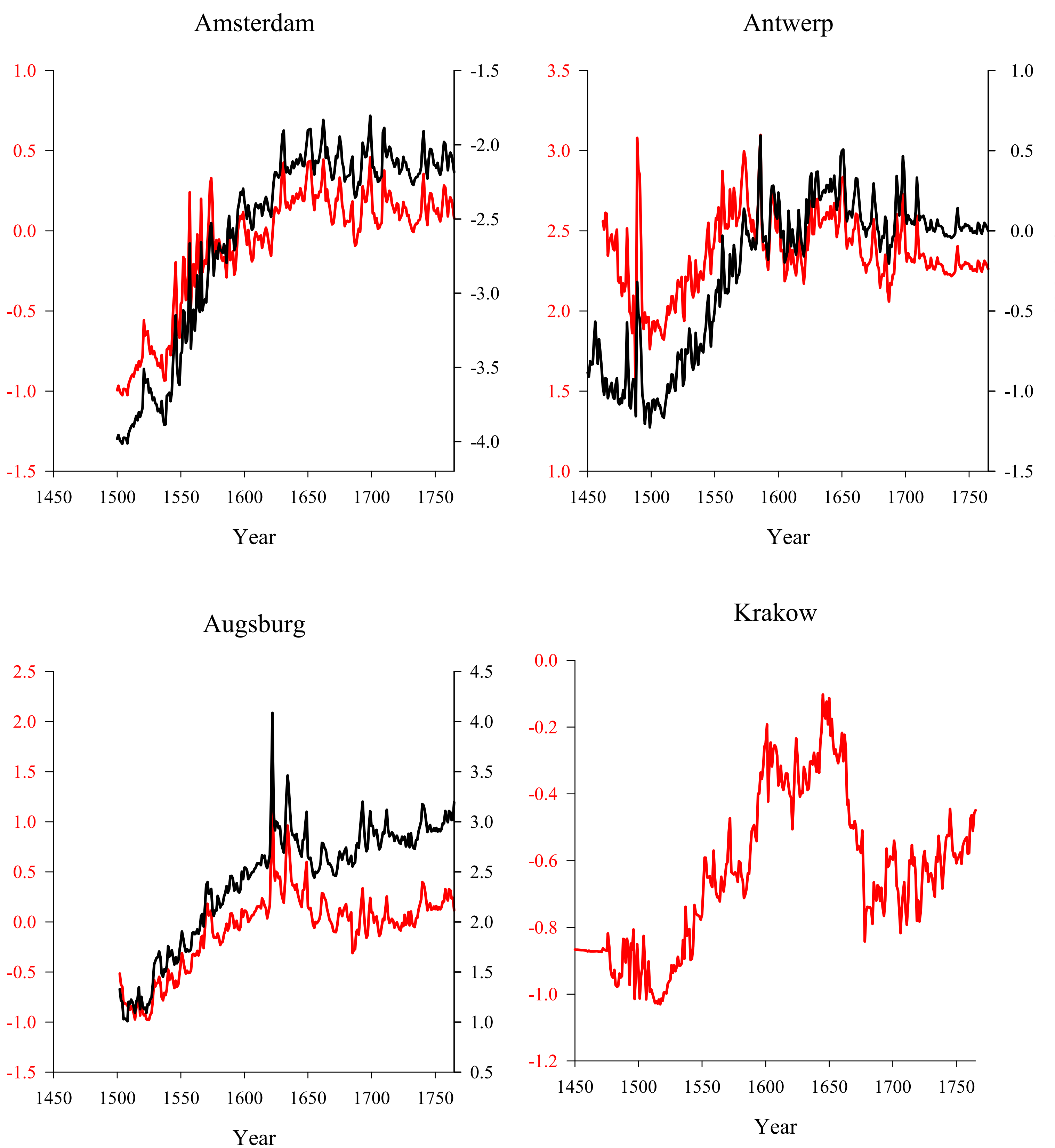
Notes: Entries are IRFs for the responses of the real exchange rates and the city inflation rates to a shock to Valencia inflation. Standard errors are HC and lag-augmented. $\chi^2(8)$ is the likelihood ratio test of the restriction that impact parameters are the same across the 9 cities.

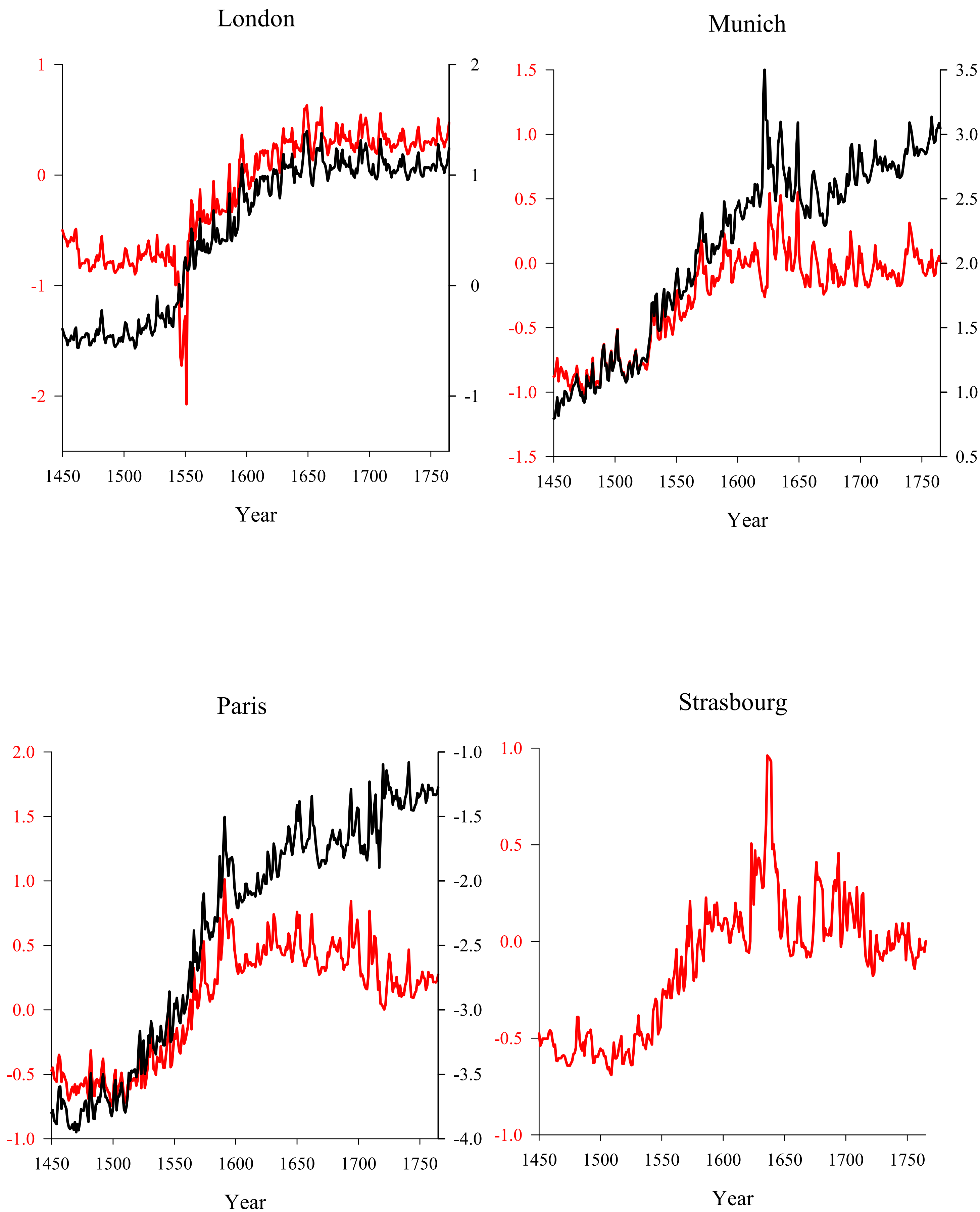
Table 8. IRFs for Real Exchange Rates and Inflation
Madrid Inflation Shock: 1551–1650

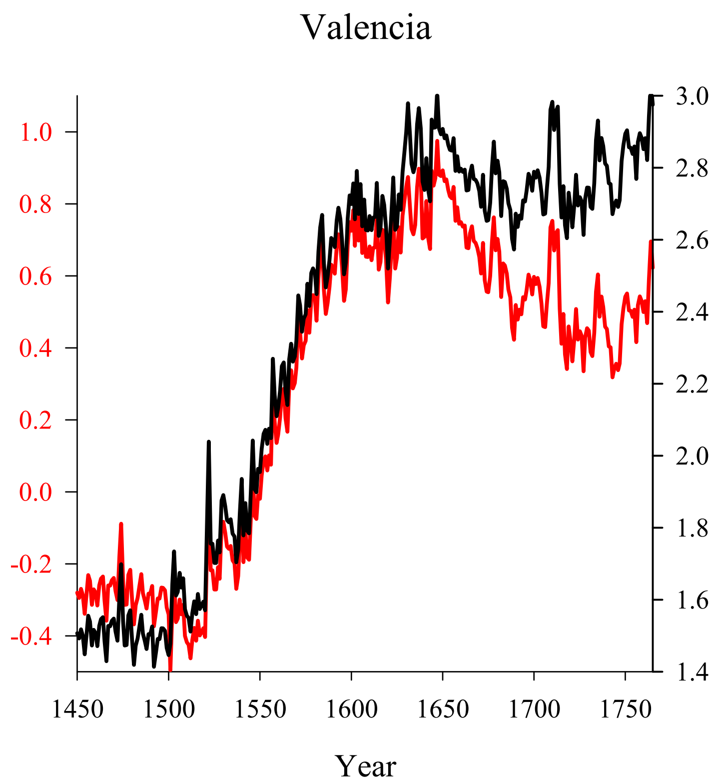
	$q_{j,t+h} - q_{j,t-1}$		$p_{j,t+h} - p_{j,t-1}$	
Horizon	$\hat{\rho}_h$ (se)	$\chi^2(8)$ (p)	$\hat{\lambda}_h$ (se)	$\chi^2(8)$ (p)
0	0.83	13.2	0.17	13.2
	(0.04)	(0.16)	(0.04)	(0.16)
1	0.53	9.8	0.24	8.8
	(0.11)	(0.36)	(0.05)	(0.46)
2	0.39	7.9	0.29	8.9
	(0.11)	(0.54)	(0.06)	(0.44)
3	0.18	3.3	0.15	6.3
	(0.10)	(0.95)	(0.07)	(0.71)
4	0.01	8.5	0.11	10.8
	(0.10)	(0.48)	(0.07)	(0.29)
5	-0.03	10.6	0.17	12.1
	(0.09)	(0.30)	(0.08)	(0.21)

Notes: Entries are IRFs for the responses of the real exchange rates and the city inflation rates to a shock to Madrid inflation. Standard errors are HC and lag-augmented. $\chi^2(8)$ is the likelihood ratio test of the restriction that impact parameters are the same across the 10 cities.

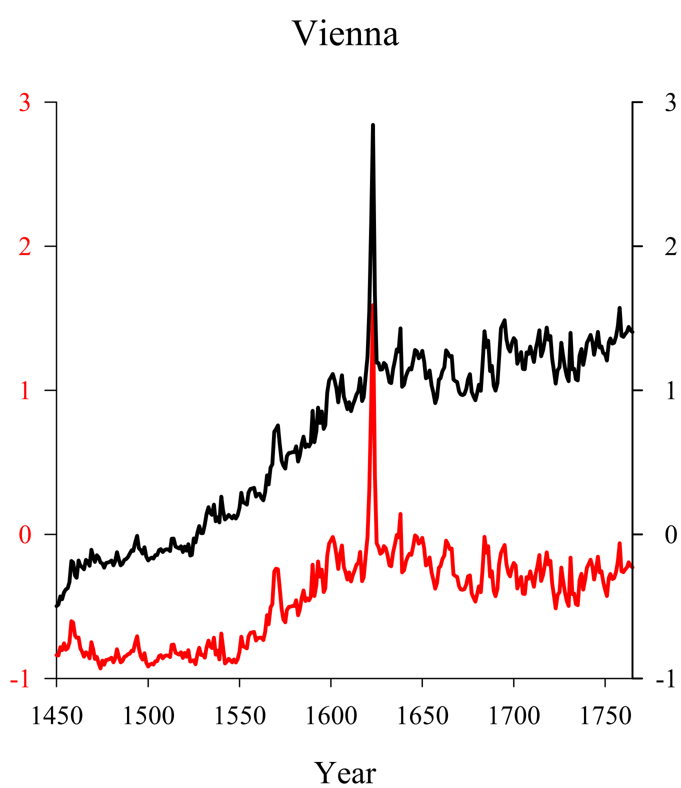
Figure 1: City Log Price Indexes in Local Currency and Silver




London
1
0
-1
-2
2
1
0
-1
1450
1500
1550
1600
1650
1700
1750
Year
Munich
1.5
1.0
0.5
0.0
-0.5
-1.0
-1.5
3.5
3.0
2.5
2.0
1.5
1.0
0.5
1450
1500
1550
1600
1650
1700
1750
Year
Paris
2.0
1.5
1.0
0.5
0.0
-0.5
-1.0
-1.0
-1.5
-2.0
-2.5
-3.0
-3.5
-4.0
1450
1500
1550
1600
1650
1700
1750
Year
Strasbourg
1.0
0.5
0.0
-0.5
-1.0
1450
1500
1550
1600
1650
1700
1750
Year

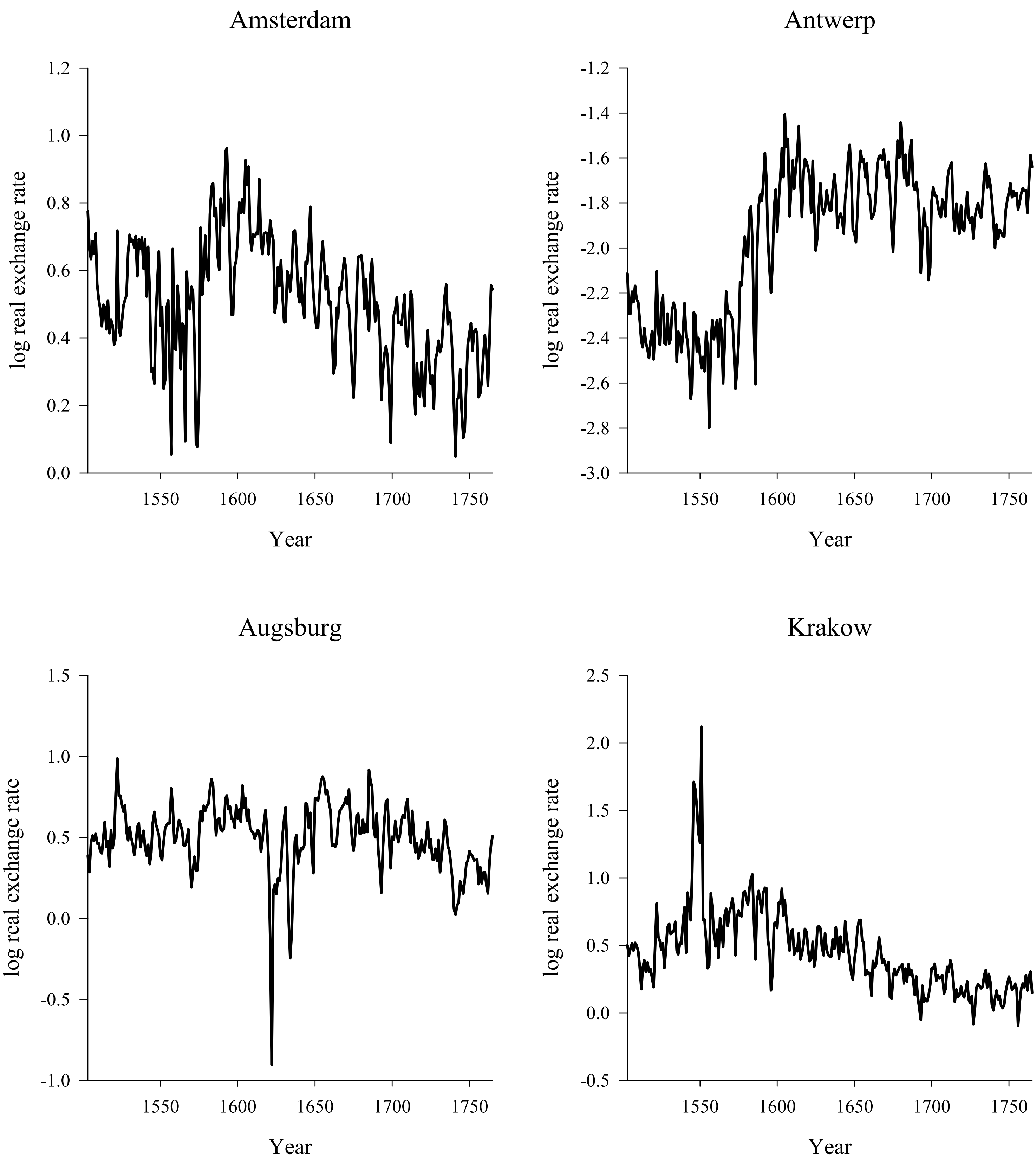


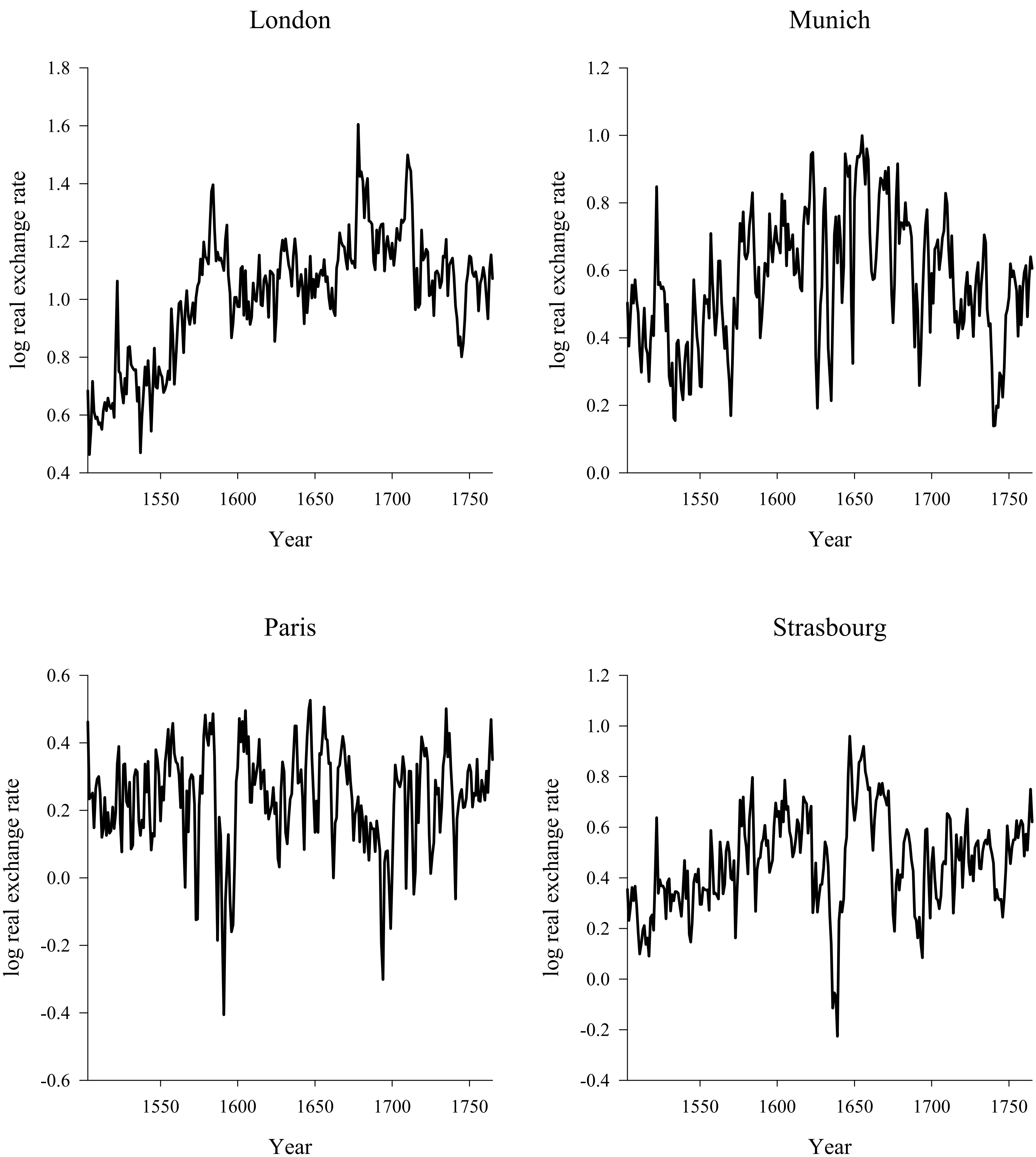
Valencia
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
-0.4
3.0
2.8
2.6
2.4
2.2
2.0
1.8
1.6
1.4
1450
1500
1550
1600
1650
1700
1750
Year

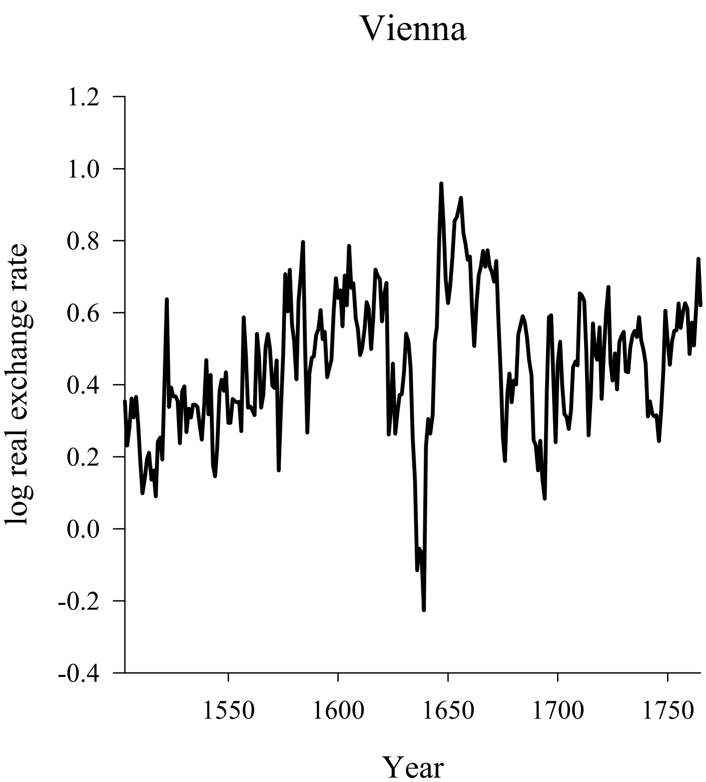


Vienna
3
2
1
0
-1
3
2
1
0
-1
1450
1500
1550
1600
1650
1700
1750
Year

Figure 2: Log Real Exchange Rates 1503-1765




London
log real exchange rate
1.8
1.6
1.4
1.2
1.0
0.8
0.6
0.4
1550
1600
1650
1700
1750
Year
Munich
log real exchange rate
1.2
1.0
0.8
0.6
0.4
0.2
0.0
1550
1600
1650
1700
1750
Year
Paris
log real exchange rate
0.6
0.4
0.2
0.0
-0.2
-0.4
-0.6
1550
1600
1650
1700
1750
Year
Strasbourg
log real exchange rate
1.2
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
-0.4
1550
1600
1650
1700
1750
Year



Vienna
log real exchange rate
1.2
1.0
0.8
0.6
0.4
0.2
0.0
-0.2
-0.4
1550
1600
1650
1700
1750
Year

Figure 3: City Longitudes and Latitudes