

Real Exchange Rates in the European Price Revolution

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Abstract

The European price revolution that began in the 16th century is one of the earliest episodes of widely-recognized inflation. I use inter-city real exchange rates (measured in silver) to assess how extensively it spread and how rapidly long-run PPP was restored. Real exchange rates appear to be stationary with half-lives similar to those in modern data: Thus there was a shared price revolution. But I also find that debasements contributed significantly to local-currency inflation. Tracking these relative prices also reveals some notable differences across cities that are worthy of further study.

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1. Introduction

The term ‘price revolution’ refers to a period of relatively high inflation beginning in the 16th century in Europe. Historians have long attributed this episode to the flow of precious metals (especially silver) from the Americas. In studying modern data 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 from 1503 to 1765. The cities are 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 address three main questions. First, the price level in Valencia roughly followed the shape of a logistic curve, rising in the 1500s before levelling off in the 1600s. I document that real exchange rates are stationary, so that price indexes in the other cities share this general pattern. That finding shows that the price revolution occurred in all these cities and extended as far northeast as Krakow.

Second, I then estimate the half-life of departures from long-run purchasing power parity (PPP) for each city. The average half-life is 3.75 years, which is comparable to estimates found in modern studies. However, there is considerable variation across cities. The variation is related to distance from Spain, with nearby cities displaying lower half-lives, but only weakly so. These statistics thus reveal some interesting geographical heterogeneity in the price revolution.

Third, the price indexes are available in silver and in local currency (though not for Krakow and Strasbourg). Comparing average inflation rates in silver and in local currency

shows that the local-currency component was important and varied across cities. I use these measures to construct relative prices of silver, which are the synthetic analogues to modern nominal exchange rates. Nominal depreciations reflect differential currency debasements across cities. I then assess whether these depreciations were predictable from real exchange rates. For Amsterdam, Augsburg, and Vienna they were not. But for London, and to a lesser extent Antwerp, Munich, and Paris, there is an early modern Big Mac effect, whereby the nominal exchange rate appears to respond to departures from PPP.

As background, Munro (2008) offers a concise introduction to research on the price revolution. Hamilton’s (1934, 1947) books are classic sources. The chapter by Braudel and Spooner (1967) is a further 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) documents and discusses the diverse impacts of American precious metals on European countries. I focus instead on the differential inflation experiences of a group 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 study ten cities for which there are data over 1503–1765, as listed in the introduction. 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 residential rent and there is no panel of price indexes that includes rent on a consistent basis for many cities. The online appendix cites the studies of price indexes that do include rent for a small number of cities. It is possible that mean reversion would be slower in indexes using these shelter costs (and hence slower than

in modern data), because shelter is a nontraded good. However, other components of the baskets, such as perishable foods, also 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.

The prices are recorded both in silver and in local currency for each city (except for Krakow and Strasbourg where they are only in silver). 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 physical wear on coins. For simplicity I use the term debasement to refer to any erosion of the value of the local currency in silver, whether sudden or gradual.

Figure 1 shows log CPIs in silver (in black) and in local currency (in grey) for three example cities: Amsterdam, Paris, and Valencia. (Graphs for the other cities are in the online appendix.) The time span in the figure is 1450–1765 (where data are available) to put the price revolution in context by including data before it began. The general upward trend over this period of course constitutes the price revolution. But notice that inflation rates tended to fall after 1600: That feature should help assess whether price levels were linked across cities.

For each city the log prices in silver (black line, left axis) and in local currency (grey line, right 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 grey line rises more than the black line reflecting debasements of the local currency. Cumulative inflation in local currency exceeds that in silver for each city in the sample. I study the predictability of these depreciations below in section 7.

For several cities there are some not-so-gradual changes. For example, Figure 1 shows a spike in the price level for Paris in the late 1500s. For Augsburg such outliers appear in 1621–1623 and for Vienna in 1621–1624. These are the early years of the Thirty Years’ War which brought cataclysmic events to both cities. For London and Munich 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 tests for stationarity and estimates of mean-reversion that are robust to outliers.

The focus here is on *relative* price revolutions. To begin to see how these price indexes 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 Amsterdam and Paris. Now the graphs apply to 1503–1765, a span which provides the largest balanced panel. (Again the online appendix provides graphs for the remaining cities.) These real exchange rates 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.

I study several 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) how correlated are real and nominal exchange rates; and (5) whether currency depreciations relative to silver respond to real exchange rates. One of the goals is to compare these properties with those found for modern real exchange rates.

3. 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. The span of 242 years should also 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 1 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.

However, section 2 referred to several city-specific outliers in the price index series and those are inherited by the real exchange rates because Valencia did not experience the same historical events. When there is an additive outlier (AO), tests for a unit root can over-reject and so incorrectly suggest that the real exchange rate is stationary. I therefore also calculated the test statistics recommended by Perron (2006), which test for an additive outlier and then control for it in constructing augmented Dickey-Fuller (ADF) statistics. The lag length is found using the BIC just as with the Elliott tests.

The Perron procedure reports outliers in 1590 for Paris, 1620 for Augsburg, 1638 for Strasbourg, 1621 for Vienna, and 1550 for London and then controls for them. The last column of Table 1 shows the resulting ADF t -statistics, which range from -5.87 (Strasbourg) to -9.15 (Vienna). These are well below the 1% asymptotic critical value (which is -3.43)

and also well below the finite-sample critical values based on the lag length and sample size from the response surfaces of Cheung and Lai (1995). Allowing for an additive outlier leads thus to the same conclusion (as does allowing for two outliers). Based on these results, I therefore treat the $q_{j,t}$ as stationary and use standard tools of inference in studying them below.

Historically, these findings suggest that the price revolution applied to all 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 the unit-root tests or Figure 2. In fact, graphs of $\{q_{j,t}\}$ like those in Figure 2 provide 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 the graphs do not show such a general pattern in the real exchange rates.

4. Mean Reversion is Comparable to that in Modern Data

To characterize the mean reversion in each real exchange rate, I use this statistical model:

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

Here mean reversion will be reflected in a negative coefficient $\hat{\beta}_j$.

Table 2 contains OLS estimates $\hat{\beta}_j$ and their HAC standard errors and p -values. All the estimates are negative and most are estimated quite precisely. In absolute value the smallest values are for Antwerp (-0.087) and Krakow (-0.105), 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 (-0.341), Augsburg (-0.256), and Amsterdam (-0.228). Overall, cheap (expensive) cities experience real appreciation (depreciation), restoring long-run PPP. (The results are very similar if I zoom in on the span 1503–1650, as in Hamilton’s 1934 study.)

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 modified by West and Cho (1995) to be robust to heteroskedasticity. The p -values show there is little evidence of remaining dynamics in $\hat{\epsilon}_{j,t}$. I also tested for nonlinearity or threshold effects using a variety of LM tests and visual inspection of scatter plots corresponding to equation (2). Those tests did not reveal any such patterns.

Recall from section 2 that there are outliers for several 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 -0.209 (-0.256), for London -0.112 (-0.168), and for Vienna -0.131 (-0.202). The LAD estimates are negative and generally comparable in scale to those from OLS. (The online appendix contains the LAD estimates for each city.) I also estimated the adjustment speeds treating the additive outliers selected in section 3 with dummy variables. The results (not shown) again were similar to those in Table 2.

The half-life of departures from PPP, denoted h_j for city j , is given by:

$$h_j = \frac{\ln(0.5)}{\ln(1 + \beta_j)}. \quad (3)$$

Table 2 also shows half-life estimates $\hat{h}_j$ along with standard errors and p -values. Point estimates of half-lives range from 1.66 years for Paris to 7.61 years for Antwerp. The average value across the nine cities is 3.75 years.

To put these findings 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 six 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.

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 half-lives found in Table 2 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 $T = 262$ years and the point estimates of persistence ($1 + \hat{\beta}_j$) are not close to one.

Treating Valencia as the base city allows study of a long time span, including the entire sixteenth century. However, 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). The Castilian city in the Allen data is Madrid. In the Allen data its price index begins in 1551, giving $T = 214$ observations until 1765. As a check on the findings, I calculated real exchange rates with Madrid as the base city. I included Valencia among the other J cities. The results (not shown) are similar to those reported above.

5. Evidence is Weak for Distance Effects in Mean Reversion

Section 3 concludes that real exchange rates are stationary, so that the price revolution in log price levels is common across cities. But section 4 then shows that the speed of adjustment to this pattern, benchmarked to Valencia, is by no means uniform. For example, half-lives are measured quite precisely for Paris and Antwerp and the two are quite different. I next study whether the heterogeneity of adjustment speeds or half-lives is related to distance from Valencia. This approach fits with the traditional view of the price revolution as originating in Spain.

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) and that in Valencia. The online appendix lists the church or cathedral in each city, its latitude and longitude, and the associated distance from Valencia Cathedral.

Call d_j the distance to city j from Valencia. 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 replace distance d_j with its rank r_j . That replacement may resolve the first issue and does resolve the second one, since rank regression is linear for any monotone function.

The rank correlation between half-life and distance is 0.32. Thus, the further from Valencia is a city the greater is the half-life of departures from PPP. But the Bonett standard error for this correlation is 0.37. Thus there is only very weak evidence of a distance effect. Direct inspection of Table 2 of course gives the same result. Paris is closer to Valencia than London is and much closer than Krakow is, and their half-lives (1.66, 3.76, 6.28 years) are ranked accordingly. But Antwerp is closer to Valencia than Amsterdam

is or Vienna is yet the Valencia-Antwerp real exchange rate has the largest half life (7.61 years).

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.

6. Debasement Contributed Significantly to Local-Currency Inflation

It has long been common to describe the price revolution by saying that average inflation rates were not high by modern standards but that they persisted over more than a century and so led to large changes in price levels. In keeping with that tradition, I next record average inflation rates for each city. They are measured in percent per year and calculated from cumulative changes in levels over 1503–1765 (i.e. without log approximation). Table 3 gives these average inflation rates for prices measured in silver, labelled $\overline{\Delta p_j}$. For most cities these rates are close to the rate in Valencia (0.33% per year). The exceptions are Antwerp and Krakow which had lower average inflation rates (0.15% and 0.18% respectively).

Table 3 then gives the average rate of real depreciation for each city, labelled $\overline{\Delta q_j}$. To two decimal places, these growth rates satisfy the differenced version of the definition of the log real exchange rate:

$$\Delta q_{j,t} = \Delta p_{val,t} - \Delta p_{j,t}, \quad (4)$$

because the annual growth rates are relatively small. These rates of real depreciation are generally small, which corresponds with the findings on stationarity of real exchange rates in Table 1. There the evidence of stationarity was weakest for Antwerp and Krakow. But notice that Amsterdam, London, and Paris all had higher inflation rates in silver than Valencia did, and so had modest rates of real appreciation.

Reconstructing prices in local currency begins with Allen’s (2001) data on values of local currencies in terms of silver (for example grams per guilder in Amsterdam), 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. I omit Krakow and Strasbourg from this list because the prices of their currencies in silver are not available for the full sample period. Denote the log value of the currency in silver (i.e. the inverse of the local price of silver) by $a_{j,t}$. These values tend to fall over time, reflecting debasements for example. Debasements occurred as a sovereign reduced the amount of silver in coins 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. (The online appendix provides further historical references on debasements in early modern Europe.)

Call the log price index in local currency $r_{j,t}$. The value of local currency in silver, the price index in silver, and the price index in local currency are linked by this identity:

$$a_{j,t} \equiv p_{j,t} - r_{j,t}. \quad (5)$$

Thus, a decrease in the value of local currency in terms of silver corresponds with an increase in prices of goods in local currency relative to prices of goods in silver.

The fourth column of Table 3 gives the average rate of inflation in local currency $\overline{\Delta r_j}$. The highest rates are for Paris (0.96% per year) and Augsburg (0.75% per year). The average, local-currency component of inflation is $\overline{\Delta r_j} - \overline{\Delta p_j}$. This component is positive in every city: Debasements caused inflation rates in local currency to exceed those in silver. The average local-currency component of inflation is (a) comparable in scale to the average inflation rate in silver across cities and (b) more variable across cities. Thus, these components seem just as worthy of study as are inflation rates in silver, which have traditionally been attributed to metal imports from the Americas.

Karaman, Pamuk, and Yildirim-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, but not by economic factors involving the prices of commodities or precious metals. Palma (2022, p 1609) finds that debasements were not

predicted by inflows of precious metals to Europe. The next section investigates properties of *relative* debasements.

7. Predictability of Nominal Exchange Rates Varies Across Cities

The analogue to a log nominal exchange rate is the relative value of local currency in silver:

$$s_{j,t} = a_{val,t} - a_{j,t}. \quad (6)$$

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. Combining this definition with the identity (5) gives an alternate expression:

$$s_{j,t} = (p_{val,t} - r_{val,t}) - (p_{j,t} - r_{j,t}). \quad (7)$$

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.

The rate of nominal depreciation thus is:

$$\Delta s_{j,t} = (\Delta p_{val,t} - \Delta r_{val,t}) - (\Delta p_{j,t} - \Delta r_{j,t}). \quad (8)$$

Time-averages of these rates are shown in the last column of Table 3, labelled $\overline{\Delta s_j}$. There were on average nominal depreciations, relative to Valencia, in each of the seven cities for which one can construct the nominal exchange rate in the Allen data. These ranged from 0.02% per year for London to 0.44% per year for Paris, over the 1503–1765 period.

I tested for unit roots in the nominal exchange rates again using the Perron procedure with AO adjustment. The results suggested that the $\{s_{j,t}\}$ are stationary. Modern studies of international macroeconomics often find a correlation between the nominal and real exchange rates. I also calculated the sample correlations between $q_{j,t}$ and $s_{j,t}$. Table 4 contains the results. The correlations are positive for each city, but they vary greatly across

cities, from 0.01 (Augsburg) and 0.09 (Paris) to 0.74 (London) and 0.84 (Antwerp). (The table also shows the correlations between $\Delta q_{j,t}$ and $\Delta s_{j,t}$ for comparability with modern studies.) Overall, then, these correlations are lower and more heterogeneous than those for modern, floating exchange-rate regimes.

I next study whether nominal depreciations are predicted by real exchange rates. This too is a standard question asked in modern data. The statistical model is:

$$s_{j,t+h} - s_{j,t} = \xi_{j,h} + \theta_{j,h} q_{j,t} + \epsilon_{j,t+h}, \quad (9)$$

where the new subscript h measures the horizon in years. From section 4 we know that cities with high prices tend to experience real depreciations ($\hat{\beta}_j < 0$). We now see whether they also experience nominal depreciations. 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*.

Estimating equation (9) also has two appealing statistical features. First, even if the nominal exchange rate does have a unit root the h -step difference would induce stationarity, leading to a balanced regression. Second, if there is classical measurement error in nominal exchange rates that should not affect consistency of $\hat{\theta}_{j,h}$ because those exchange rates are on the left-hand side.

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 (9) 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.

Table 5 contains the results, with horizons of $h = 1, 3, 5, 7$ years. For a first group of cities—Amsterdam, Augsburg, and Vienna—there is essentially no predictable pattern

in the nominal depreciations. The estimates $\hat{\theta}_{j,h}$ are small and their p -values are large. For a second group of cities—Antwerp, London, Munich, and Paris—the coefficients $\hat{\theta}_h$ generally take small negative values. The impacts become larger in absolute value as the horizon increases and are estimated with some precision. For this second group of cities then, the synthetic, nominal exchange rate is predictable given departures from PPP. On average, a city in this group with low relative prices will tend to experience a small, nominal appreciation in subsequent years.

To study the scale of this effect, I also estimated the adjustment in the real exchange rate itself over $h = 7$ years, using the h -step version of equation (2):

$$q_{j,t+h} - q_{j,t} = \alpha_{j,h} + \beta_{j,h} q_{j,t-1} + \epsilon_{j,t+h}. \quad (10)$$

The ratio $\hat{\theta}_h/\hat{\beta}_h$ then compares the relative responses of nominal and real exchange rates to departures from PPP. For the first group of cities that ratio is essentially zero. For the second group of cities the values for $\hat{\beta}_{j,7}$ are estimated with precision and the ratios are 21% for Antwerp, 43% for London, 29% for Munich, and 16% for Paris. Eichenbaum, Johannsen, and Rebelo show that regressions like (9) and (10) can discriminate among nominal exchange rate regimes, with $\hat{\theta}_h/\hat{\beta}_h$ very small for fixed exchange rate systems and large for floating exchange rate regimes with inflation targeting. For six of these cities the ratio here is similar to the ratios that they find for managed exchange rates. The exception is London, for which there is a larger role for adjustment in the nominal exchange rate, a role that is comparable to those in some modern floating countries.

As the introduction mentioned, the goal here is to uncover some characteristics of the early modern economy of Europe by studying the price revolution. This statistical window reveals some interesting heterogeneity. For Amsterdam, Augsburg, and Vienna 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. 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 be related to real exchange rates.

This conclusion is complementary to that of Karaman, Pamuk, and Yıldırım-Karaman (2020) who show that local economic indicators do not predict debasements. The results for Antwerp, Munich, and Vienna more closely resemble findings for managed floating regimes, while the results for London are similar to those for some modern floats (as in Tables 2 and 3 of Eichenbaum, Johannsen, and Rebelo’s study).

Finally, as in section 4, I also estimated both equations (9) and (10) by LAD to allow for outliers. The results (not shown) give virtually no evidence of significant values $\hat{\theta}_{j,h}$. Thus the data for all cities appear to behave like those in a fixed exchange rate system or common currency area. Reporting the OLS results in Table 5 thus presents a more agnostic conclusion.

8. Conclusion

The European price revolution has long been viewed as an important episode in monetary history. This paper uses real exchange rates to study how it unfolded across cities. There are four 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 averages 3.75 years. This speed is comparable to those in many modern studies of exchange rates. Half-lives vary across cities but are only weakly correlated with distance from Spain, so their heterogeneity remains to be explained.
3. Averaging across time and cities, the inflation rate in local currency is roughly twice that in silver. Although there again is much heterogeneity across cities in this ratio, this finding suggests that local debasements are as worthy of study as silver imports in understanding early modern inflation.
4. For several cities, periodic debasements were not predictable from real exchange rates, so European price levels continued to be linked as if on a silver standard. But several

other cities (and especially London) display a proto-Big Mac effect in which departures from PPP help forecast nominal depreciations.

Overall, studying real exchange rates provides support for the traditional, historical view of a Europe-wide price revolution in the early modern period. But Tables 2–5 also provide new evidence of differences across cities in several dimensions. Those differences seem worthy of further research.

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Table 1: Unit Root Tests in Log Real Exchange Rates

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

City	Q_T	$DF-GLSu$	ADF t (AO-adjusted)
Amsterdam	1.128	-5.812	-7.15
Antwerp	4.328	-2.565	-8.48
Augsburg	0.879	-6.147	-8.37
Krakow	4.931	-2.946	-6.56
London	1.837	-3.919	-6.87
Munich	1.086	-4.853	-7.03
Paris	0.838	-7.149	-7.88
Strasbourg	1.249	-5.107	-5.87
Vienna	1.090	-5.411	-9.16

Notes: There are 262 observations from 1503 to 1765. The table first shows Elliott (1999) Q_T and $DF-GLSu$ 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. It then shows augmented Dickey-Fuller t -statistics adjusted for an additive outlier (AO) using the method of Perron (2006). The test statistics are found with the RATS-Estima procedure @perronbreaks.src. 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%. Each of the ADF t -statistics is significant at the 1% level using either asymptotic critical values or finite-sample values from Cheung and Lai (1995).

Table 2: Mean Reversion in Real Exchange Rates

$$q_{j,t} \equiv p_{val,t} - p_{j,t}$$

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

$$h_j = \frac{\ln(0.5)}{\ln(1 + \beta_j)}$$

City	$\hat{\beta}_j$ (se/ <i>p</i>)	$\hat{h}_j$ (se/ <i>p</i>)
Amsterdam	-0.228 (0.049/0.00)	2.68 (0.66/0.00)
Antwerp	-0.087 (0.030/0.00)	7.61 (2.75/0.01)
Augsburg	-0.256 (0.108/0.02)	2.34 (1.15/0.04)
Krakow	-0.105 (0.027/0.00)	6.28 (1.71/0.00)
London	-0.168 (0.094/0.07)	3.76 (2.30/0.10)
Munich	-0.209 (0.038/0.00)	2.96 (0.61/0.00)
Paris	-0.341 (0.058/0.00)	1.66 (0.35/0.00)
Strasbourg	-0.184 (0.043/0.00)	3.41 (0.89/0.00)
Vienna	-0.202 (0.118/0.09)	3.07 (2.02/0.13)

Notes: Brackets contain standard errors then *p*-values. Standard errors are HAC. There are 262 observations from 1503 to 1765. The half-life in years is denoted h_j .

Table 3. Average Rates of Inflation and Depreciation

City	Silver		Local Currency	
	$\overline{\Delta p_j}$ prices	$\overline{\Delta q_j}$ real	$\overline{\Delta r_j}$ prices	$\overline{\Delta s_j}$ nominal
Amsterdam	0.42	-0.09	0.69	0.13
Antwerp	0.15	0.18	0.42	0.13
Augsburg	0.28	0.05	0.75	0.33
Krakow	0.18	0.15	—	—
London	0.46	-0.13	0.63	0.02
Munich	0.29	0.04	0.69	0.25
Paris	0.37	-0.04	0.96	0.44
Strasbourg	0.23	0.10	—	—
Vienna	0.26	0.07	0.60	0.20
Valencia	0.33		0.47	

Notes: Here p is the log price level in silver, q is the log real exchange rate, r is the log price level in local currency, and s is the log nominal exchange rate. Entries are the implicit average growth rates in percent per year, calculated from cumulative changes in levels over 1503–1765. Rates of depreciation are relative to Valencia, so positive entries for $\overline{\Delta q_j}$ or $\overline{\Delta s_j}$ are real or nominal depreciations respectively for city j .

Table 4: Correlations between Real and Nominal Exchange Rates

City	$\rho(s_{j,t}, q_{j,t})$	$\rho(\Delta s_{j,t}, \Delta q_{j,t})$
Amsterdam	0.183	0.185
Antwerp	0.838	0.201
Augsburg	0.011	0.084
London	0.741	0.660
Munich	0.420	0.218
Paris	0.088	0.189
Vienna	0.346	0.026

Notes: Entries are sample correlations between log real and nominal exchange rates and then their differences, over 1503–1765. Krakow and Strasbourg are omitted because the Allen data do not show variation in their silver prices.

Table 5. Adjustment in Nominal Exchange Rates

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

Horizon→	$h = 1$	$h = 3$	$h = 5$	$h = 7$
City	$\hat{\theta}_{j,1}$	$\hat{\theta}_{j,3}$	$\hat{\theta}_{j,5}$	$\hat{\theta}_{j,7}$
↓	(se/ p)	(se/ p)	(se/ p)	(se/ p)
Amsterdam	0.002 (0.006/0.81)	-0.002 (0.016/0.88)	0.010 (0.019/0.60)	0.012 (0.022/0.57)
Antwerp	-0.009 (0.005/0.04)	-0.026 (0.008/0.00)	-0.039 (0.011/0.00)	-0.054 (0.012/0.00)
Augsburg	-0.005 (0.004/0.23)	-0.008 (0.012/0.51)	-0.011 (0.013/0.37)	0.008 (0.018/0.66)
London	-0.067 (0.102/0.51)	-0.079 (0.111/0.48)	-0.132 (0.119/0.27)	-0.210 (0.105/0.05)
Munich	-0.034 (0.034/0.32)	-0.086 (0.075/0.25)	-0.137 (0.078/0.08)	-0.191 (0.072/0.01)
Paris	-0.021 (0.025/0.41)	-0.045 (0.032/0.16)	-0.120 (0.051/0.02)	-0.156 (0.049/0.00)
Vienna	-0.005 (0.003/0.08)	-0.011 (0.006/0.07)	-0.009 (0.010/0.36)	-0.001 (0.013/0.94)

Notes: h is the horizon in years. Brackets contains HAC standard errors followed by p -values. Krakow and Strasbourg are omitted because the Allen data do not show variation in their silver prices.

Figure 1: Log Price Indexes in Silver and Local Currency

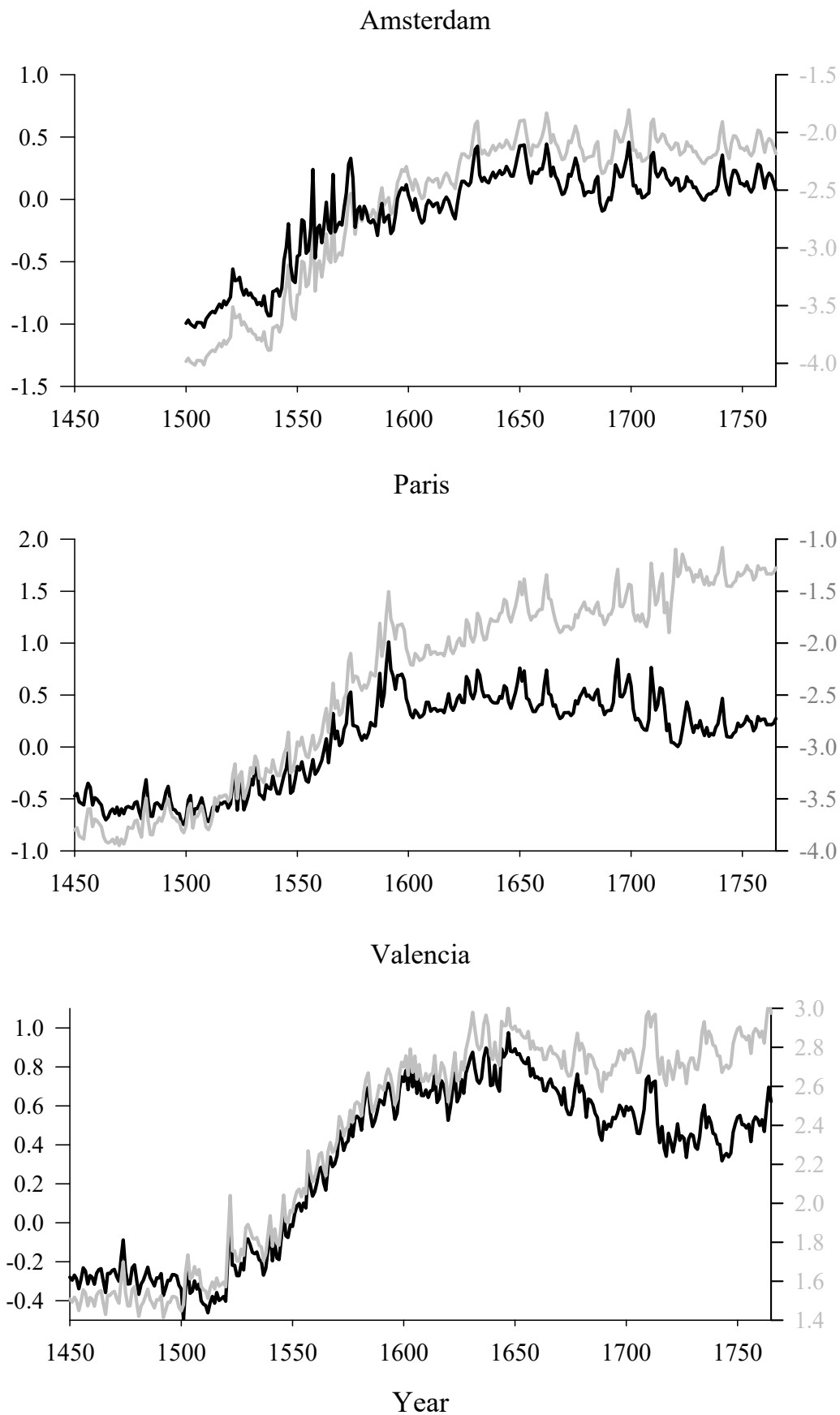


Figure 2: Log Real Exchange Rates 1503-1765

