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DWELLING PRICE DYNAMICS IN PARIS, FRANCE

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Dwelling Price Dynamics in Paris, France

by

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Abstract: Using transaction level data for dwellings in Paris, France over the period
1986-92, we find evidence consistent with the hypothesis that economic fundamentals
constrain movements in Parisian dwelling prices over longer term horizons. The
conclusion is based on the results of two different procedures for estimating an error
correction model of housing prices based on supply and demand fundamentals. The error
correction models suggest that the speed of adjustment in the Paris dwelling market is
about 30% per month. The paper also introduces a new econometric methodology that
permits simultaneous estimation of the parameters of a dynamic hedonic price model, the
price index, and the parameters of a structural model for housing prices. The new
methodology is compared with the more traditional two-step procedure of first estimating
a price index, and then using the estimated index in subsequent structural modeling.

1. Introduction

The French housing market went into recession during the middle of 1990 following a boom in residential building activity, home-sales, and housing price appreciation that began in 1986. The depth and duration of this recession and the lack of consensus about the underlying causes of the previous boom are a source of great concern to French economic policy makers. Another source of concern is a surprising lack of comprehensive and accurate statistical series available to monitor residential real estate markets in France. Without price indices, it has been difficult to determine whether the 1990 recession is a business cycle adjustment, the collapse of a "speculative bubble," or is the result of structural changes in housing supply and demand.

The paper has three objectives. First, we develop Fisher Ideal price indices using individual transaction data for dwelling unit sales between January 1987 through December 1992 in Paris.[1] Second, we apply a conventional two-step procedure in which the estimated price index is used in a second stage structural model of housing supply and demand. Third, we introduce a new methodology that allows for the simultaneous estimation of the parameters of a dynamic hedonic price model, the price index, and the parameters of a structural model for housing prices. This new estimation procedure exploits the panel nature of the transaction level data set, and accounts for market adjustments in the estimation of the index. Since the simultaneous estimation procedure we suggest is quite computer intensive, the conventional two-step procedure provides a useful comparison.

We find that the two estimation strategies lead to similar conclusions about housing price dynamics in the Paris. Our results indicate that economic fundamentals do constrain movements in housing prices over our six-year sample period. The period of analysis covers four years of a boom and two years of a downturn in the Paris housing market. We conclude that although the conventional two-step procedure is less efficient, its relative ease of application and overall similarity of results suggest that it is the more cost effective estimation strategy.

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[1] Our Paris data set contains 87,242 usable transactions. These data were obtained from the National Council of Notaries (Conseil Supérieur du Notariat).

The paper is organized as follows. Section 2 describes the nonparametric estimation of a Fisher Ideal dwelling price index for the city of Paris. Section 3 describes the conventional two-stage estimation strategy for tests of housing market dynamics and presents estimation results. Section 4 introduces a new state-space formulation of an error correction model for Paris dwelling prices, and then compares the estimation results from the two procedures.   Section 5 summarizes our findings and concludes.

2. Construction of Dwelling Price Indices and the Fundamental Data Series

At the present time, there are no price indices for existing or new dwellings in France as a whole. The database of the central taxation agency (Direction Générale des Impôts) keeps information on sales prices in the largest French cities, however, these sale numbers are unweighted means provided by "experts." The other source of information on overall trends in dwelling prices is provided by a survey of households every three to four years by the Institut National de la Statistique et Etudes Economiques (INSEE) (Taffin, 1992a).   The documentation and appraisal services provided by the notary system in France gives them access to detailed information on all real estate transactions in the country. Recently, the Chamber of Notaries of Paris and the National Council of Notaries initiated important projects to assemble their transaction information into sophisticated data banks containing accurate information on property characteristics and prices for several residential real estate markets most notably Lyon and Paris. These databases are well suited to price index development for the residential housing market and offer a first chance to study the possible determinants of housing market dynamics based on transaction information. One particularly interesting feature of these databases is that they have been linked to geo-coded socio-economic information specific to the city block.

We do not have access to the street address of properties in the Paris transaction database, so repeat sales estimation methods cannot be used. Thus, we only consider hedonic regression methods, and use a nonparametric technique to estimate the regression function. This technique is called locally weighted regression, abbreviated "loess" by

Cleveland and Devlin (1988), and Cleveland et al. (1988). Loess can successfully approximate a wide range of smooth functions. In our context it is used to estimate the attribute function G(x) in the equation

$$p_{i(t)} - \bar{p}_t = G(x_{i(t)}) + u_{i(t)},$$

where   (1)

$$\bar{p}_t = \frac{1}{N(t)} \sum_{i(t)=1}^{N(t)} p_{i(t)},$$

and $p_{i(t)}$ is the log price of the ith dwelling unit in period t, $x_{i(t)}$ is a set of property attributes, and $N(t)$ is the total number of dwelling units sold in period t. Loess requires stationary dependent and independent variables, so we remove the trend in $p_{i(t)}$ by subtracting the monthly mean of the dependent variable. The function G is estimated by running a weighted least squares regression for **each** time period t in the sample, using a subset or "window" of sales observations for **all** time periods. The observations in the subset are selected to be those most like the dwelling unit with the median set of hedonic characteristics in time t. The weights are computed as the inverse function of the Euclidean distances between the median dwelling unit attributes at time t, and the attributes all other dwelling units in the observation subset; see Meese and Wallace (1991).[2] Since Euclidean distance is sensitive to the measurement of the hedonic attributes (i.e. dwelling floor space might be measured in square meters or square feet), all attribute data is standardized to remove the effect of units of measure. This is done in the usual way, by taking the attributes and subtracting off their sample mean and dividing by their sample standard deviation.

The $x_{i(t)}$ variables in the Paris data set include the area of the dwelling unit in square meters, the floor level where the dwelling unit is located, three socio-economic

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[2] The choice of the observation window size is an important operational issue in loess. In practice the choice of the percentage of observations in the window is selected as a trade-off between bias and sampling error. In the results reported below, the window size is equal to 33%. Because we have a large number of observations, increasing the window size beyond 33% has little effect on the generated price

indicators of the quality of the location, and four composite geographic indicators. The indicators for the quality of the location are measures of the social and economic standing of the neighborhood in which the property is located. The indicator, called ilotype (îlot is a city block) was developed from a principal components analysis of thirty-five socio-economic indicators which yielded ten factors.[3] We combined these into three dummy variables.   The first level of ilotype (Ilotype dummy 1) is for middle to upper income residential neighborhoods, the second (Ilotype dummy 2) is for mixed economic use neighborhoods, and the third excluded dummy is for blue-collar working-class neighborhoods.

The twenty arrondissements in Paris are administrative jurisdictions. We include them as proxies for the geographic and infrastructure amenities of a dwelling unit's location in Paris. To conserve degrees of freedom for our monthly samples, we consolidated the twenty arrondissements into four locational dummies. We grouped the 1st, 4th through 8th, the city center arrondissements, into the first arrondissement dummy variable. The second grouping is for the "beaux quartier" residential areas of the 15th through 17th arrondissements. The third grouping is roughly the southwestern periphery of the city and includes the 2nd, 3rd, 9th through 11th, and 13th arrondissements. The omitted locational dummy is for the remaining arrondissements; they are roughly located in the northeastern periphery of the city.

In Table 1, we present the mean and the standard deviation of the loess coefficients for the attributes in standardized units. As expected we find that the squared metric area of a dwelling has a large positive effect on price and that it exhibits little variation over the sample of seventy-two months. The upper-income ilotype and the city center and "beaux quartiers" arrondissement measures also produce positive mean implicit prices, again with little variation over the seventy two sample period.

We construct our price index, by adjusting the monthly average dwelling price by the loess estimates of the implicit attribute prices for each month. The initial year and last year median attributes are used to form the Paasche and Laspeyres price indices

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index (our major interest), and on the distribution of the hedonic slope coefficients for each time period.
[3] The measure was developed at the Chambre de Notaires de Paris. We did not have access to the raw

respectively. The geometric average of these two indices is the Fisher Ideal index. Figure 1 presents a graphical comparison of the Fisher Ideal index with indices constructed from the simple mean and median prices for each month's transactions. The mean and the median indices are higher than the Fisher Ideal price index because they do not accurately account for changes in the attribute characteristics of dwelling units sold. However, all three indices demonstrate the run-up in prices from early in 1988 through May of 1991. Over the whole sample period the nominal cost of dwelling units increased about 60% .

The trend in all three of the indices (mean, median, and Fisher Ideal) appears to be stochastic, as there are obvious breaks in the movement of dwelling prices over time. The first appears to occur after the bank liberalization policies in 1987, where the price change begins to increase rapidly. A downward trend is evident at the beginning of 1992. Application of conventional unit root tests confirms the presence of a stochastic trend in the Fisher Ideal index. Table 2, presents results from an augmented Dickey-Fuller (1979) test statistic for a unit root in the Fisher Ideal price index. The test statistic is consistent with a unit root null at all conventional significance levels. Application of nonparametric tests of Phillips-Perron (1988) confirms the findings in Table 2. Both tests are conducted with an estimated time trend to account for nonzero drift in the growth rate of the indices. We include two lags in all our tests and find that a third lag is insignificant in all cases.

There are a number of additional series that will be used in our fundamental model of the Parisian dwelling market. These include a rental index for the city of Paris, household revenue (an income proxy), employment, a residential construction cost index, and two cost-of-capital series. Since evidence on the nature of the trend in these additional series will be important to the model development of the next section, Table 2 also presents augmented Dickey-Fuller tests for the additional macro-economic series. In addition, the Paris rental, revenue, employment and construction cost series are available only on a quarterly basis. We first apply econometric interpolation procedures, as discussed in the appendix, to construct the monthly versions of these fundamental series. We do so by using related series that are available on a monthly basis.

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data on socio-economic indicator variables.

There are two different cost-of-capital measures. The first is a dwelling-owner cost-of-capital series. It is measured by the long-term private signature rate adjusted by time varying regional and city property tax rates[4] using Kearl (1979) and Dougherty and Van Order (1982) methods. We do not adjust for income taxes because there is very limited use of interest deductibility[5] (Bosvieux and LeLaidier, 1994 and Riou, 1994) nor do we adjust for the depreciation rate. It can be assumed to be constant in our log specification and relative short sample. The second cost-of-capital measure does not adjust for tax rates and is simply the long-term private signature rate. The test results indicate that the rental series and the dwelling-owner cost-of-capital series appear to be integrated of order one, however, the non-tax adjusted cost-of-capital measure fails to reject the null of a no unit root.[6] The unit root test statistics are statistically significant at conventional levels for the household revenue and the employment per household series, however, the residential construction cost index is only marginally significant.

3. Fundamental Determinants of Dwelling Prices

There are a number of competing theories about the causes of the increase in dwelling prices as shown in Figure 1 from January 1988 through about January 1991, and the slowdown appearing throughout 1992. French demographers have identified significant increases in "décohabitation" during the last twenty years as traditional households break into smaller units. Increases in divorce rates, numbers of individuals living alone or in single parent households and decreases in marriage rates and intergenerational living arrangements have all contributed to a fall in the size of households from 2.70 people in 1982 to 2.46 people in 1995 (Louvet, 1989). Although the demographic indicators suggest sustained demand-side pressure on French housing markets, other indicators suggest that the rapid increases in home ownership rates in

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[4] We use the property tax rate (taxe sur le foncier bâti) including garbage collection taxes which are applied against 50% of the assessed value of the property. We also add in a household tax (taxe d'habitation) which is computed as a function of property value.

[5] In the subsidized part of the French mortgage market, there is a limited form of interest rate deductibility. The deductions are available for the mortgage interest charges but they are limited by ceilings and can only be used for the first five years of the loan. The importance of this deduction is further reduced because the loan programs that allow them account for a small proportion of the overall French mortgage market.

[6] These results are again reinforced by the nonparametric tests of Phillips and Perron (1988).

France (40% rate in 1965 to a 55% rate in 1992 (CIEC, 1992)) cannot be sustained. One important constraint has been the relatively high levels of long term real interest rates from 1986 through 1992 and a progressive and increasingly effective policy of salary de-indexation in France. In addition, government subsidies in the form of public construction of housing and mortgage subsidies have steadily declined.[7] The combined effect of these changes has decreased the purchasing power of French households.

With the exception of 1985-1989, housing construction levels have steadily declined from their 1974 high of 550,000 new projects in one year. In the context of the long term evolution of the French housing market, the relatively high production levels in 1989 (350,000 new projects) only brought production back to its 1982 level (CIEC, 1992). A frequently cited reason for the upsurge in residential construction from 1985-1990 is the French banking liberalization policies initiated in 1987 which led to a strategic deployment of funds into the French real estate market (Nappi, 1993).[8]

The French banking liberalization measures also coincide with banking liberalization policies in other countries, in particular in Sweden and Japan. From 1989 though 1990 foreign investment in real estate (primarily office and commercial) doubled in France (Nappi, 1993). Finally, many market participants now claim that the lack of reliable pricing information at all levels of French real estate markets was another important influence on housing price dynamics from 1985 through 1992.

In the next section, we postulate a quasi reduced-form equilibrium model of the supply and demand for housing services. Using the Fisher Ideal price indices, we estimate an autoregressive distributed lag model for the price index as a function of our supply and demand fundamental variables. We also test for the number of cointegrating vectors and then estimate an error correction model for housing prices. This benchmark estimation strategy provides a comparison for the dynamic hedonic model in state space form presented in Section 4.

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[7] The reduction in the subsidized sector has been particularly important in the new housing sector. In 1984, the subsidized loans for the construction of rental social housing (PLA-Prêts du secteur locatif aidé) and the subsidized (PAP - prêts pour l'accession à la propriété) and regulated loans for new homeowners (PC - prêts conventionnés) accounted for 50% of housing and 46% of the investment amount. In 1990, its share was 25% and 19% respectively (Bosvieux and Le Laidier, 1994).

[8] The lending levels of French banks to real estate developers and real estate syndicators increased sixfold

3.1 Tests of Long-run Demand and Supply Determinants of Housing Prices

We assume that the "long run" demand for the stock of housing services can be written as:

$$Q_d(t) = W(t)' \boldsymbol{g}_d + u_d(t), \tag{2}$$

where $W(t)$ denotes the vector of demand determinants, including the housing price. For the non-price fundamentals in $W(t)$ we use data on per household revenue, employment per household, and the homeowner cost-of-capital.[9] We expect all variables except price and the cost-of-capital to have positive coefficients; these coefficients are denoted by the parameter vector $\boldsymbol{g}_d$. The term $u_d(t)$ is the structural error in the demand schedule.

The long run supply equation for each municipality has the form:

$$Q_s(t) = Z(t)' \boldsymbol{g}_s + u_s(t), \tag{3}$$

where $Z(t)$ is the vector of supply determinants. We assume that the nonprice supply fundamentals include an index of construction costs and a proxy for the opportunity cost-of-capital. We expect all components of $Z(t)$ except price to shift the supply schedule inward and thus we anticipate negative coefficients, $\boldsymbol{g}_s$. The term $u_s(t)$ denotes the supply schedule disturbance. Our assumptions allow us to sign all coefficients in the quasi-reduced form for housing price (all are positive), with the exception of the interest rate variable.[10]

Our strategy is to estimate an autoregressive distributed lag (ARDL) model for price as a function of construction costs, the interest rate variable, employment, and the

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in constant francs from its 1988 levels to its 1990 levels (Nappi, 1993)

[9] We were unable to obtain demographic variables or a series on household formation.

[10] We were unable to obtain a separate cost-of-capital measure for both the supply and demand equations. Thus, we cannot differentiate supply and demand equation interest rate effects from our reduced form estimates.

real income proxy. The results are reported in Table 3, and are based on the following equation:[11]

$$P_t = a_0 + a_1 P_{t-1} + a_2 P_{t-2} + a_3 C_t + a_4 C_{t-1} + a_5 r_t + a_6 r_{t-1}$$
$$+ a_7 E_t + a_8 E_{t-1} + a_9 Y_t + a_{10} Y_{t-1} + \varepsilon_t \tag{4}$$

The ARDL representation above can be rewritten as an error correction model (ECM) in differences, where the "equilibrium" long run relation between the price index P_t, the construction cost variable C_t, the cost-of-capital r_t, employment E_t, and the real income proxy Y_t can be inferred from the coefficients in (4) after rearranging terms:

$$P_t - P_{t-1} = a_0 - a_2(P_{t-1} - P_{t-2}) + a_3(C_t - C_{t-1}) + a_5(r_t - r_{t-1})$$
$$+ a_7(E_t - E_{t-1}) + a_9(Y_t - Y_{t-1}) +$$
$$b[P_{t-1} + (a_4 + a_3)C_{t-1}/b + (a_6 + a_5)r_{t-1}/b$$
$$+ (a_8 + a_7)E_{t-1}/b + (a_{10} + a_9)Y_{t-1}/b] + \mathbf{e}_t, \tag{5}$$

where $b = (a_2 + a_1 - 1) < 0$ is the speed of adjustment to disequilibrium in the housing market.

The OLS point estimates reported in Table 3 give rise to an error correction parameter $b = (.187 + .497 - 1) = -.316$. Plugging the OLS point estimates into (5) yields the long run, quasi-reduced form relation: $P = 6.57\ C + 7.07\ r + 5.48\ E + .646\ Y$. While the signs in this relation are consistent with our priors (and with a dominant interest rate effect from the supply equation), it is difficult to compare coefficient magnitudes because the quasi-reduced form model that we estimate does not identify the individual supply and demand elasticities in equations (2) and (3).

We now test the assumption that there is only one cointegrating relation between price and its macroeconomic fundamentals,[12] using the test of Johansen and Juselius (1990). We find (using a 5% significance level) a single cointegrating vector for the five series P, C, r, E, and Y. In contrast to results for equation (4), the Johansen and Juselius normalized cointegrating coefficient on the real income proxy is negative, and the building cost coefficient is four times larger. All of the cointegrating coefficients are at least twice

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[11] We fit the model in logarithmic form so that all coefficients are interpretable as elasticities and variable scaling is irrelevant. Interest rate variables are constructed as the logarithm of one plus the rate. Two lags of price are included in (4) to ensure a serially uncorrelated disturbance term.

[12] Indeed, if we had constructed a quantity index as well as price a index, then we might expect to find

their asymptotic standard errors, with the exception of the interest rate variable. This

latter finding is consistent with our earlier discussion.

On economic grounds we prefer the coefficients of the quasi-reduced form for the

price index P estimated from equation (4), as all coefficient signs are consistent with our

priors, and the magnitudes of the coefficients are more uniform. An augmented Dickey-

Fuller unit root test confirms that this linear combination of price and its fundamentals is

stationary, as we can reject the null of a unit root at a 1% significance level. This lends

further credence to our long run supply and demand model for the Parisian dwelling

market, as in general, a linear combination of variables with some unit roots need not be

stationary. Define the disequilibrium error $v = (P - 6.57\ C - 7.07\ r\ - 5.48\ E - .646\ Y\ )$.

We now proceed to directly estimate an error correction version of (4) using our preferred

proxy for the disequilibrium error: [13]

$$P_t - P_{t-1} = b_0 + b_1(P_{t-1} - P_{t-2}) + b_2(C_t - C_{t-1}) + b_3(r_t - r_{t-1})$$
$$+ b_4(E_t - E_{t-1}) + b_5(Y_t - Y_{t-1}) + b_6 v_{t-1} + u_t \tag{6}$$

Before doing so we must deal with one additional complication. The price index is a

generated regressor, and its appearance on the RHS of (4), (5), and (6) can complicate

statistical inference.[14] To account for estimation error in the price index, we estimate the

ECM (6) by both ordinary least squares (which ignores the errors-in-variables problem),

and by an instrumental variables (IV) procedure that uses contemporaneous and a single

lag of differences in fundamentals as instruments for $(P_{t-1} - P_{t-2})$ and v_{t-1} in (6).

Accounting for estimation error in P will induce a second order moving average

error term in the composite disturbance of equation (6).[15] We use a Newey West (1987)

heteroskedasticity and autocorrelation consistent covariance estimator for both the OLS

and IV standard errors of the parameters in (6). These estimated standard errors are thus

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cointegrating vectors corresponding to both the structural demand and supply equations.

[13] See Engle and Granger (1987), Campbell and Perron (1991), and Banerjee, et al. (1993) for a thorough discussion of ECM estimation. Economic theory can be imposed in the equilibrium error term v, while lagged changes in prices and fundamentals account for any short-run dynamics that are not usually subject to theoretical priors. When we estimate (6) using the Johansen and Juselius cointegrating vector, the error correction coefficient drops to $b_6 = -.09$, with a p-value of 11%.

[14] Pagan (1984) is the standard reference for econometric issues in regressions with generated regressors.

[15] Substitute $P = P^* + z$, where P is the actual price index, P^* the estimated index, and z the estimation error, into the ECM and collect values of z and v into a composite disturbance term.

robust to serial correlation and potential heteroskedasticity in the composite disturbance term. The results for both OLS and IV with robust standard errors estimation of (6) are reported in Table 4a and 4b respectively.

The estimated error correction coefficient is $b_6 = -.32$ (-.38) with a marginal p-value of 0.14% (35%) when (6) is estimated by OLS (IV with robust standard errors), respectively. Either point estimates suggests that about one third of the discrepancy between actual and fundamental price is removed each month. Contemporaneous changes in fundamentals have marginal explanatory power in (6) as a joint Wald test for zero coefficients (b_{2-5} are jointly zero) on these four terms has a p-value of 1.3% (10%) for the two estimation procedures. The R^2 is reasonable for a model fit in differences (35%), and the equation passes all standard residual diagnostic tests (normality, lack of serial correlation, no conditional heteroskedasticity, coefficient stability, and functional form). We also fit (6) with a dummy variable to allow for an asymmetric response in the speed of adjustment parameter b_6 and find that this complication is unnecessary.[16]

We conclude this section with one last comment on our implicit assumption that the error terms in equations (4) or (6) are uncorrelated with contemporaneous values of the regresssors. This assumption seems especially suspect for both the time t employment and income variables. Construction costs and interest rates are more likely to be uncorrelated with the error term in our Parisian real estate model, while the local employment and revenue (income proxy) variables are likely to be jointly determined with housing prices. When we attempt to account for the joint determination of employment and income using an IV procedure, the results in Table 4a are not robust. A better way to proceed may be to jointly estimate the error correction models for all the jointly determined variables, this we leave as a future research question.

4.   A Dynamic Hedonic Model in State Space Form

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[16]More specifically, we define a dummy variable equal to 1 when v is above its mean and zero below. When a lagged interaction term consisting of the dummy multiplied by v is included in (11), its coefficient is zero to three decimal places and has a p-value of 31%.

In this section we propose a new method for simultaneously (i) estimating a dynamic hedonic house price model, (ii) generating a housing price index, and (iii) fitting an ECM using the generated price index. Treating the price index as the unobserved state variable while letting it evolve as an ECM exploits the panel nature of the data, and solves the errors-in-variables problem noted in the previous section. In much of the published work in real estate economics, a price index is used on either the left and/or right hand sides of structural equations that explain the demand and supply of housing services, without any adjustment for the estimation error in the price index.

The first equation of our composite model can be written as:

$$p_{i,t} = a_t + \sum_{k=1}^{K} b_{t,k} z_{i,t,k} + e_{i,t}, \tag{7}$$

where $p_{i,t}$ is the log price of the i-th dwelling in period t; a_t is the price index in period t; $b_{t,k}$ indicates hedonic price coefficient k in period t; $z_{i,t,k}$ is the k-th characteristic of dwelling i in period t; $e_{i,t}$ is a disturbance term; and $t=1,...,T$.

Equation (7) is the measurement equation in our state space framework. The transition equation (8) describes the evolution of the unobserved state variable (the housing price index in our context) a_t as an ECM:

$$a_t - a_{t-1} = c_0 + c_1(a_{t-1} - P^*_{t-1}) + c_2(a_{t-1} - a_{t-2}) + \upsilon_t, \tag{8}$$

where P^*_{t-1} is a measure of fundamental housing price in period $(t-1)$[17], υ_t is an error term, and c_0, c_1, and c_2 are parameters that govern the process.[18]

The measurement equation error term $\varepsilon_{i,t}$ is assumed to obey:

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[17] Alternatively, think of the difference between the variables a and P* as the disequilibrium error defined in the previous section.

[18] Given the difficulty in maximizing the likelihood of the model (7) and (8), we have excluded contemporaneous values of the change in fundamentals from (8). Recall that the p-value for the joint significance of these variables was around 1% (10%) for the model of Section 3.1 using OLS (IV) estimation procedures. In addition, a single lag of the dependent variable produces a model with serially uncorrelated residuals.

$$E(\boldsymbol{e}_{i,t}) = 0; \quad Var(\boldsymbol{e}_{i,t}) = \boldsymbol{s}_t^2; \quad \text{cov}(\varepsilon_{i,t}, \varepsilon_{j,s}) = \theta_t > 0 \text{ for } t = s, i \neq j;$$

$$\text{cov}(\boldsymbol{e}_{i,t}, \boldsymbol{e}_{j,s}) = 0, \text{ for } i \neq j; t \neq s.$$

Since we do not have information on repeat sales, we cannot estimate a component of variance due to idiosyncratic house elements as in Quigley (1995). Our disturbance specification is still quite general, as it allows for different error variances and contemporaneous covariances over time. We must assume that shocks to dwellings in the same time period t are positively correlated, to ensure positive definiteness of the disturbance covariance matrix of (7). It seems intuitively reasonable to assume that any period t macroeconomic shock has the same effect on all contemporaneous dwelling prices.

The transition equation (8) error term υ_t is assumed to have classical properties; zero mean, constant variance, and no temporal dependence with itself or the measurement equation disturbance. Using the prediction formulae in Harvey (1994), the system of equations (7) and (8) can be written in terms of the one step ahead prediction errors. Assuming normality of these errors, the likelihood function can be maximized to generate parameter estimates for $b_{t,k}$, c_i, and the disturbance covariance terms. An estimate of the unobservable price index a_t (and its variance) is generated by successive applications of the Kalman Filter. Again following Harvey (1994, pp. 141-144), we use recursive analytic expressions for the likelihood score function and information matrix; both involve only the calculation of first derivatives. The method of scoring is used to find the likelihood maximum, as described by Harvey (1982), chapter 4.

The likelihood function can be concentrated with respect to parameters in the measurement equation (7), given an estimate of the state a_t. The parameters that govern the evolution of the error correction model (7) need to be estimated using the likelihood for the system of equations (7) and (8). Further simplification of the estimation of the parameters in (8) is accomplished by noting that price indices like a_t are typically normalized to a base year. This multiplicative degree of freedom means that we can set the variance of u_t equal to unity, without loss of generality.

In order to proceed with the estimation of the model (7) and (8) we need a proxy for the fundamental price P_t^*. In previous work, Meese and Wallace (1994), we have used capitalized rents as an indicator of fundamental housing price. The advantage of this proxy is that the present value model that links dwelling prices, rents and the cost-of-capital has known coefficients. Since the coefficients are known, we do not need to estimate additional parameters in the system (7) and (8) when using capitalized rents as the proxy for P_t^*. As of this writing, we have been unable to generate results from state space representation (7) and (8), when P_t^* is modeled as the equilibrium solution to our set of supply and demand equations (2) and (3) with estimated coefficients. Thus we have adopted an alternative definition of the equilibrium price series P_t^*.

More specifically, we generate P_t^* from the forward solution of the present value relation for housing prices assuming the cost-of-capital variable (discount rate) is known at the beginning of the period, and extraneous solutions or "bubbles" are ruled out of the analysis. Following Meese and Wallace (1994) we define fundamental price as:

$$P_t^* = \sum_{i=1}^{\infty} [\frac{1}{(1+r_t)}]^i E(R_{t+i}|I_t).$$ (9)

where R_t is the rental cost index in month t, the discount rate is one over one plus the homeowner cost-of-capital, $1/(1+r_t)$, and $E(- | I_t)$ denotes the expectation operator conditional on the information set I_t. Assuming that the expected growth rate of the rental series is a constant q:

$$E(R_t - R_{t-1}|I_{t-1}) = q,$$ (10)

the present value solution to (9), which we define as the fundamental price index P_t^*, is

$$P_t^* = R_t / r_t + q(1 + r_t) / r_t.$$ (11)

In Figure 2, we plot the series for P_t^* and the observed Fisher Ideal Price Index P_t of Section 2. To generate P_t^* we use an appropriately scaled rental index series R_t, its estimated growth rate, and the tax adjusted cost-of-capital series r_t. The plots indicate that the present value prices are greater than the actual prices briefly at the beginning of

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the period and then again during the steep run-up in prices from late 1987 through 1989. As the index reaches its peak, the present value index falls below the observed price index until the beginning of 1992. These results suggest that deviations of present value prices from actual observed dwelling unit prices are followed by adjustments of the observed price back to equilibrium levels, consistent with an ECM framework.

To estimate the state space model (7) and (8) using our generated price series P_t^* we again use the log of dwelling price for N=87,242 individual sales in the city of Paris from January 1987 through December 1992.[19] There are K=7 attributes; these include the log of (1+dwelling floor), the log of living space (measured in square meters), two dummy variables for ilotype, and three dummy variables for geographic grouping of Parisian arrondissements. We use the value of the information matrix at convergence to generate asymptotic standard errors of the estimated parameters, which are reported in Table 5. In order to conserve space we report both the mean and standard deviation of the parametric hedonic coefficient estimates over the 72 months, as in Table 1.

The parameter estimates for the state space model (7) and (8) are reasonable, and very similar to the OLS and IV estimates we reported for equation (6). The speed of adjustment parameter c_1 is estimated to be -.39, a number quite close to the value obtained using the IV estimator of Section 3.1 that also accounts for estimation error in the price index. Standard errors on the speed of adjustment and lagged price change coefficients are smaller than for equation (6), indicating more precise estimation of parameters with the system approach (7) and (8).

Figure 3 compares the estimated state space price index a_t with the Fisher Ideal index of Section 2. Two features of the state space index are striking. First, it is much smoother than the Fisher Ideal index, and second, it closely tracks (but is smoother than) the present value price index of Figure 2, which is based on capitalized rents. Apparently, the state space price index is quite sensitive to the fundamental series used for P_t^*.

The average values of the seven hedonic coefficients (across the 72 months of dwelling sales data) are also reported in Table 5, and are similar to the results in Table 1.

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[19] The maximum number of dwelling sales in any month is 1946 and the minimum is 471. August is always the low sales volume month, given French vacation traditions.

The notable exception is the average coefficient on the vertical floor level variable, which is now negative but small relative to its standard error. Residual diagnostics for each cross sectional measurement equation indicates a residual distribution with thick tails, just as with the nonparametric approach of Section 2. Given the effort required to estimated the system of equations (7) and (8), the overall similarity of the results for equations (6) and (8), and the correspondence between the estimated state space price index, the less efficient approach of Section 3.1 would appear to be a more cost-effective estimation strategy.

5. Conclusions

We have presented evidence consistent with the hypothesis that economic fundamentals constrain movements in Parisian dwelling prices over a longer-run horizon. The conclusion is based on two different procedures for estimating an error correction model of housing prices based on macroeconomic fundamentals. Our results suggest that the speed of adjustment of Parisian dwelling markets to previous differences between fundamental and actual price is in the neighborhood of 33-40% per month. This speed of adjustment is about three times faster than we found in the San Francisco Bay area housing market over a similar period, although in the case of California, the historical run-up in dwelling prices was much more dramatic.

The graphical analysis presented in Figures 1, 2, and 3 indicates that Parisian nominal dwelling prices increased at most 60% over our sample, and that there are prolonged periods when fundamental price remains above or below actual price. The difference between fundamental and actual prices is greatest during the latter half of 1990, when actual price exceeds fundamental price by about 30%. In the other noteworthy episode, fundamental price lies below actual by as much as 20% during the eighteen-month period January 1988 to August 1989.

Future research using data sets like the one analyzed here should address two unresolved issues. First, the generation of Fisher Ideal quantity indexes would permit separate analysis of both the long run supply and demand equations. Second, in order to fully solve the errors-in-variables problem in the state space framework, we need to extend our methodology to cover simultaneous estimation of the cointegrating vector (fundamental

price), the dynamic hedonic, and the error correction model. Solving the latter problem will be much more difficult than the former, and may fail to pass muster of a cost benefit analysis.
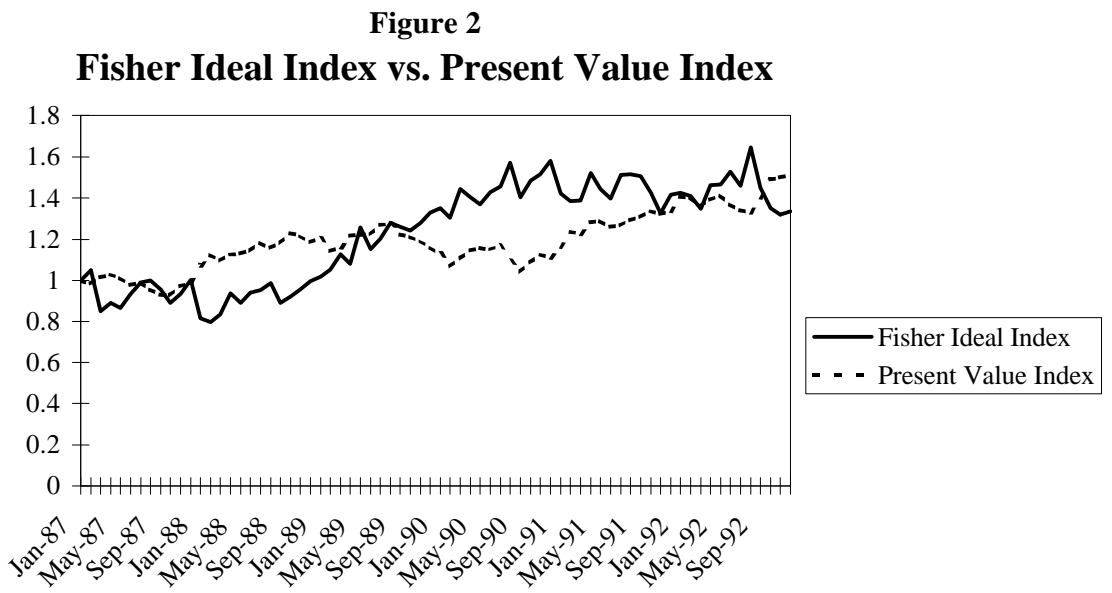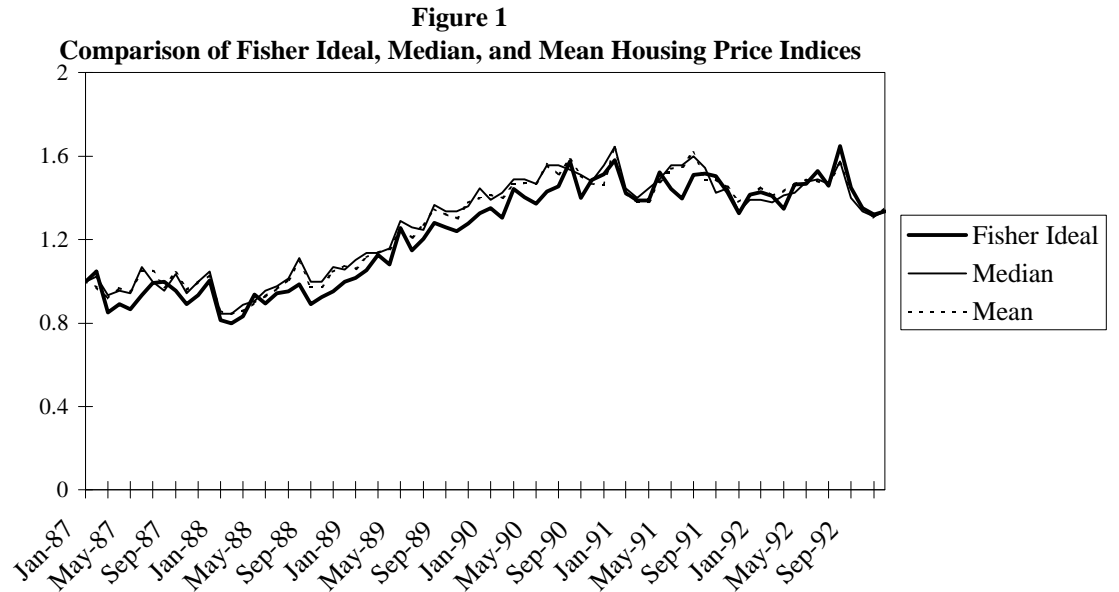
Figure 1
Comparison of Fisher Ideal, Median, and Mean Housing Price Indices

Figure 2
Fisher Ideal Index vs. Present Value Index

Figure 3
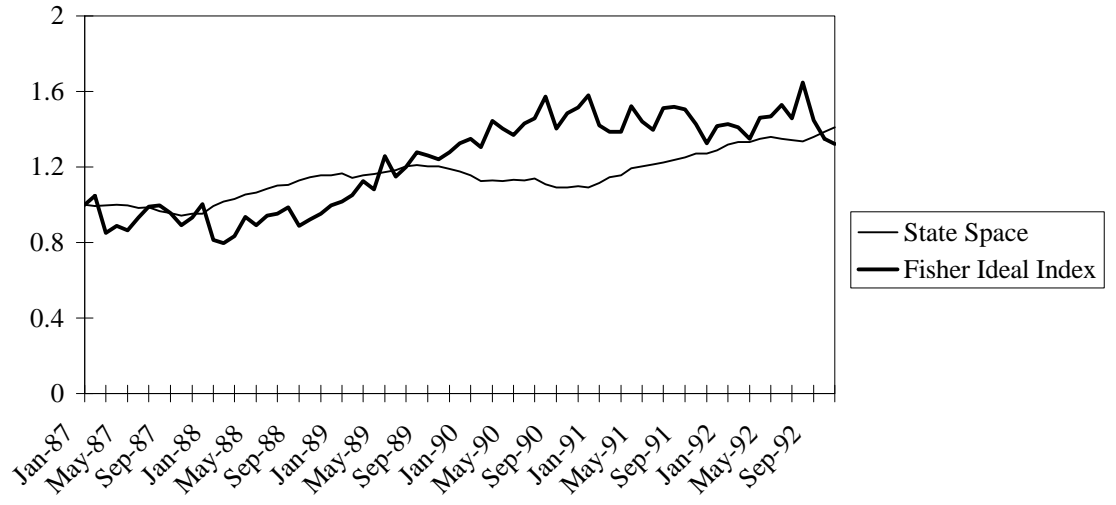**State Space Index vs. Fisher Ideal Index**

Table 1
Averages and Standard Deviations for loess estimated
attributes in Paris: January 1987 - December 1992[a.]

Attributes	Average	Standard Dev.
Square meters of floor space	.745	.028
Vertical floor location	.034	.023
Ilotype dummy 1 - Middle to Upper Income Residential	.162	.049
Ilotype dummy 2 - Mixed economic uses (Residential and commercial)	.074	.045
Arrondissement dummy 1 - Arr. 1, 4-8. (City Center)	.178	.027
Arrondissement dummy 2 - Arr. 15-17. (Beaux Quartiers)	.106	.031
Arrondissement dummy 3 - Arr. 2,3, 9-11, 13. (Southwestern Periphery)	.028	.028

[a.] Statistics based on 87,242 total transactions. Average and Standard Dev. statistics are based on 72 monthly estimates using the loess procedure. The magnitudes are hard to interpret: they are based on standardized (logarithmic) regressors and a dependent variable measured as deviations around the mean of the logarithmic of French Franc dwelling prices. Clearly, a positive coefficient means that attribute is positively priced.

Table 2

Unit root tests for Housing Prices, rents and market fundamentals

Regression Results[a.]		Augmented Dickey-Fuller test statistic	
Fisher Ideal price index for Paris (Trend)		-1.25	
Rental index for Paris (Trend)		-3.41	
Dwelling Cost-of-capital - Property tax adjusted. (No Trend)		-2.21	
Cost-of-capital (No Trend)		-1.55	
Annual real, household revenue (Trend)		-4.31	
Employment per household (Trend)		-4.72	
Residential Construction Cost Index (Trend)		-3.07	
MacKinnon critical values	10%	5%	1%
Trend: T=69	-3.27	-3.47	-4.09
No Trend: T=69	-2.59	-2.90	-3.53

[a.] Tests for the null hypothesis of a unit root in the price, employment, construction cost and income variables include a time trend and two lagged changes, whereas the unit root tests for the cost-of-capital omits the time trend.

Table 3
OLS[a.] Estimation of the ARDL model (4)

Variable	Parameter	Coefficient Estimate	Std. Error	Asymptotic p-value (%)
Intercept	a_0	-14.4	7.64	6.49
P_{t-1}	a_1	.497	.123	.02
P_{t-2}	a_2	.187	.122	13.1
C_t	a_3	1.40	1.11	21.1
C_{t-1}	a_4	.690	1.12	53.9
r_t	a_5	8.42	3.07	.80
r_{t-1}	a_6	-6.19	3.00	4.37
E_t	a_7	1.63	1.24	19.3
E_{t-1}	a_8	.102	1.21	93.3
Y_t	a_9	-1.07	.590	7.37
Y_{t-1}	a_{10}	1.28	.569	2.82

[a.] OLS results are for the period January 1987 through December 1992; T=70 observations on the dependent variable. Asymptotic p-values are the null hypothesis of a zero coefficient.

Table 4a
OLS estimation of the ECM equation (6) and model diagnostics

Variable	Coefficient	Std. Error	Asymptotic p-value (%)
Intercept	-14.4	4.27	.14
$(P_{t-1} - P_{t-2})$	-.187	.111	9.70
$(C_t - C_{t-1})$	1.40	.868	11.3
$(r_t - r_{t-1})$	8.42	2.80	.39
$(E_t - E_{t-1})$	1.63	1.09	14.2
$(Y_t - Y_{t-1})$	-1.07	.477	2.78
(v_{t-1})	-.317	.0947	.14

Functional Form[1] (Ramsey)   1.83 (18%)

Constant Variance[2] (White)   16.1 (14%)

ARCH test [3] (Engle)   1.20 (27%)

Normality[4] (Jarque-Bera)   .324 (85%)

Serial Correlation[5] (Breusch-Godfrey)   15.0 (24%)

Residual diagnostic p-values in parentheses
[1.] RESET test with fitted squared terms only.
[2.] White test with linear and squared terms only (12 constraints)
[3.] ARCH test with l lagged squared residual.
[4.] Skewness=0, Kurtosis=3 (two constraints)
[5.] LM test with 12 lags

Table 4b
IV estimation of the ECM equation (6) and model diagnostics

Variable	Coefficient	Std. Error	Asymptotic p-value (%)
Intercept	-17.1	18.2	35.1
$(P_{t-1} - P_{t-2})$	.254	.299	39.9
$(C_t - C_{t-1})$	1.24	1.99	53.7
$(r_t - r_{t-1})$	7.63	4.91	12.5
$(E_t - E_{t-1})$	1.26	1.46	38.9
$(Y_t - Y_{t-1})$	-1.03	.515	4.96
(v_{t-1})	-.380	.404	35.0

Functional Form[1] (Ramsey)	.392 (53%)
Constant Variance[2] (White)	15.7 (15%)
ARCH test[3] (Engle)	.0838 (77%)
Normality[4] (Jarque-Bera)	.452 (79.8%)
Serial Correlation[5] (Breusch-Godfrey)	21.1 (5.0%)

Residual diagnostic p-values in parentheses

[1.] RESET test with fitted squared terms only.

[2.] White test with linear and squared terms only (12 constraints)

[3.] ARCH test with l lagged squared residual.

[4.] Skewness=0, Kurtosis=3 (two constraints)

[5.] LM test with 12 lags

Table 5
Maximum Likelihood Estimates of the state space model (7) and (8)
January 1987 - December 1992[a.]

Variable	Coefficient	Standard Error
Square Meters of Floor Space	1.10	.170
Floor Location	-.00955	.0692
Ilotype 1 - Middle to Upper Income Residential	.166	.125
Ilotype 2 - Mixed Use: Residential and commercial	-.0136	.132
Arrondissement Dummy 1 - Arr. 1, 4-8 (City Center)	.443	.0862
Arrondissement Dummy 2 - Arr. 15-17 (Beaux Quartier)	.276	.085
Arrondissement Dummy 3 - Arr. 9-11, 13 (SW Periphery)	.0364	.111
Intercept	-.0933	.00215
$(a_{t-1} - P^*_{t-1})$	-.392	.0514
$(a_{t-1} - a_{t-2})$	-.517	.170

[a.] All statistics are based on estimates 87,242 total transactions. For the hedonic characteristics, average and standard error statistics are calculated using 72 sets of parametric hedonic coefficient estimates for equation (7).

Appendix A

The data sources for this analysis were obtained from a variety of sources. The house price and characteristic data were obtained from the Chambre de Notaires of Paris. The fundamental series were obtained from three sources: the monthly series were obtained from the *Bulletin Mensuelle de la Statistique* (BMS) published by the Institut National de la Statistique et Etudes Economiques (INSEE), the quarterly series were obtained from OECD and from DATASTREAM which compiles data from INSEE, and the annual series were obtained from the *Annuaire de la statistique* published by INSEE. The series obtained for household formation, household revenue, employment, and the interest rate series are French national series. The rental indices and CPI are for the city of Paris.

We used a monthly frequency in our estimation and applied the Chow-Lin (1971; 1976) generalized least squares (GLS) procedure to interpolate the needed series for the market fundamentals. This procedure is the best linear unbiased estimator and is, therefore, preferable to Kalman filter techniques for interpolation. We interpolate from two different series: quarterly series using monthly, related series and annual series using monthly related series. Initial diagnostics suggest that the quarterly residuals for these series follow a first order autoregressive process. The covariance matrix needed for the GLS estimator is obtained by assuming that the residual structure is AR(1). The covariance across the observations is estimated using the two-stage iterative maximum likelihood estimator in TSP. It is not clear whether the residuals from the annual series exhibit autocorrelation, because there are only six data points for these series. For this reason, we use ordinary least squares to perform the interpolations of the annual series using monthly, related series.

The quarterly index of residential construction costs was obtained from DATASTREAM. The series was interpolated to a monthly series using a monthly index of French overall building costs. The results from the Cho-Lin procedure are

Table A.1

Dependent Variable: Residential Building Costs

Variable	Coefficient Estimate	Standard Error
Intercept	337.41	98.10
French Construction Cost Index	1.36	.22
$\hat{e}_{-1}$	.75	.13
R^2	.95	

The coefficient on $\hat{e}_{-1}$ is the quarterly autocorrelation coefficient. All of the estimated coefficients are statistically significant at the .05 level or better.

The quarterly rental price index for Paris is interpolated using a monthly index for total housing services for France and the overall consumer price index for Paris. The results from the Cho-Lin interpolation procedure are

Table A.2

Dependent Variable: Paris Rental Price Index

Variable	Coefficient Estimate	Standard Error
Intercept	144.55	96.01
Consumer Price Index for Paris	-1.52	1.24
French Housing Services Index	1.65	.65
$\hat{e}_{-1}$	.63	.16
R^2	.97	

The coefficient estimate for the total housing service index and that for the autocorrelation coefficient are statistically significant at the .05 level, however, the Paris CPI is statistically significant at only the .10 level.

The quarterly level of total employment is interpolated using monthly levels of the demand for and offers for employment in France and monthly values for the level of total population in France. The results of the Cho-Lin estimating procedure is

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Table A.3

Dependent Variable: Total Employment

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Variable	Coefficient Estimate	Standard Error
Intercept	.689E07	.522E07
Demand for Employment in France	-49.38	79.37
Offers of Employment in France	513.93	1254.26
Total French Population	271.61	94.182
R^2	.997	

The results indicate that total French population is the only statistically significant determinant of total employment in France. The offer and demand rates are not statistically significant determinants of employment.

The annual level of household revenue is interpolated using monthly series for the Paris, CPI, labor wage index for manufacturing production and an index for industrial production.

Table A.4

Dependent Variable: Household Total Revenue

Variable	Coefficient Estimate	Standard Error
Intercept	144.55	96.01
Paris CPI	1.968	.446
Wage Index for manufacturing	-.038	.035
Index for Industrial Production	.123	.358
R^2	.97	

The only statistically significant determinant of revenue per household is the Paris CPI. The index for industrial production and the wage index for manufacturing have no statistically significant effect on revenues per household.

The annual level of household size was interpolated using monthly series on French population, French marriage rates, and French birth rates. The results of the Cho-Lin interpolation are

Table A.5

Dependent Variable: Total Number of Households

Variable	Coefficient Estimate	Standard Error
Intercept	15,041	1750.84
French Population		
	.644	.030
French marriage rates		
	1.65	.65
French birth rates		
	1.02	.75
R^2	.97	

The results show that the most important determinant of the number of households is the French population and the marriage rate. The number of births does not have a statistically significant effect on the observed level of household formation.

We use ratios of employment to number of households, because the French population variable is used to interpolate both series. In an effort to control for the effect of population on these two variables we use the ratio of total employment to total number of households. Our employment measure is thus the average number of employees per household.

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