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The cyclical behaviour of inventories: European cross-country evidence from the early 1990s recession

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This paper employs data for a panel of firms from France, Italy and the UK to study the effect of the recession of the early 1990s on inventory investment, controlling for cyclical fluctuations at the firm level. The results clearly show some common patterns across countries, pointing to the relevance of financial factors (namely, the level of leverage) in propagating initial recessionary shocks. Moreover, Italian firms, especially if 'small and young', seem more likely to suffer from a reduction in the value of collateralizable assets possibly originated by restrictive policy actions.

I. INTRODUCTION

Recent work in macroeconomics has emphasized the role of credit market imperfections in propagating original disturbances from various sources, amplifying their real effects. In the present paper this issue is addressed by providing novel microeconometric evidence from a largely unexploited dataset on balance sheets of European firms. The focus is on three countries, France, Italy and the UK, studying the cyclical behaviour of inventories, since they are likely to be the component of firms' assets most responsive to financial pressure and general adverse macroeconomic conditions. The recent launch of the European Monetary Union (EMU) and the current debate over the UK participation provide additional interest to this investigation, which might yield important insights on the differences and similarities in the channels of monetary transmission across European countries.

The paper is organized as follows. The next section provides a brief discussion of the relevant literature on financial factors and cyclical fluctuations and spells out the motivation of our work. Section III describes the data used (additional details on the construction of the final panels of firms are provided in the Appendix). Section IV presents the empirical specifications of the inventory equations and discusses methodological issues. Section V is the core of the paper, where empirical results are presented both for the full samples and for groups of firms partitioned according to proxies for the degree of access to capital markets. The main conclusions are summarized in the final Section VI.

II. MOTIVATION AND RELATED LITERATURE

Theoretical research has framed the issue of the effects of capital market imperfections on real economy fluctuations within the 'principal-agent' view of credit markets, studying how endogenous changes in the agency cost of lending over the business cycle are responsible for amplifying the real impact of initial shocks, giving rise to a 'financial accelerator' effect (see Bernanke *et al.* 1996, 1999, for overviews). Informational asymmetries between lenders and borrowers make external finance, if not fully collateralized, more expensive than internal sources of funds. This external finance premium compensates lenders for evaluation and monitoring activities and its size is negatively

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correlated with borrowers' net worth (the sum of net liquid assets and the collateral value of fixed assets). Negative shocks to net worth reduce the amount of internal finance available to borrowers, at the same time increasing the premium on external finance: these effects reinforce each other and determine sharp reductions in the borrowers' level of investment and production.

Empirical work on these issues has taken several directions. One strand of literature, initiated by Fazzari *et al.* (1988) and surveyed by Schiantarelli (1995) and Hubbard (1998), focused on the potential effects of financial constraints on firms' investment expenditure adding proxies for the availability of internal funds to investment equations. A related literature concentrated on the 'lending channel' of monetary policy transmission, giving a primary role to the response of bank loan supply in the face of monetary tightening (see, among others, Bernanke and Blinder, 1992; Kashyap *et al.* 1993; Oliner and Rudebusch, 1995) using time-series data on aggregate credit and bank loans during monetary policy restrictions. Such time-series studies have gradually acknowledged the fundamental identification problems encountered in the interpretation of aggregate, macroeconomic data. In fact, empirical work on the dynamic effects of monetary policy carried out using mainly Structural Vector Autoregressive (VAR) techniques suffers from a fundamental identification problem: a decline of bank loans following a monetary policy restriction does not necessarily reflect a squeeze in loan supply but might well be endogenously driven by a fall in loan demand. In the latter case there would not be a specific role for banks in the transmission of monetary policy shocks. Even results from the analysis of changes in the composition of firms' external finance between bank loans and commercial paper (Kashyap *et al.*, 1993, for the USA) are not conclusive if a monetary contraction causes a general shift of all forms of credit from small to large firms: a decline in the bank-loan share might be observed in aggregate data, with no substitution away from bank debt at the firm level, only because large firms typically rely less heavily on bank debt than small firms (Oliner and Rudebusch, 1995).

Two main insights emerge from this earlier literature. First, even if there is scant evidence in favour of a 'lending channel', a broader 'credit channel', perfectly consistent with the financial accelerator view, might be at work, whereby (maybe policy-induced) increases in market interest rates cause an increase in the premium for external debt of all sorts charged on at least certain classes of borrowers. Second, any convincing evidence of a specific role of financial market imperfections must come from empirical research focusing on the differential response of agents to recessionary shocks: identification must be achieved through agents' heterogeneity (Gilchrist and Zakrajšek, 1995).

Taking this lead, the recent empirical literature has focused on firms' behaviour, mainly in the USA, following two empirical approaches. Using time-series data on firms disaggregated by size class (as a proxy for capital market access), Gertler and Gilchrist (1994) provide substantial evidence of differential behaviour of small versus large firms over the business cycle and in response to monetary policy tightening. Overall, small firms experience a deeper contraction than large firms in a wide variety of variables, including sales and inventories. Broadly similar results are reported also by Bernanke *et al.* (1996) and Oliner and Rudebusch (1996). However, also non-financial reasons may explain the observed differential behaviour among firms, especially for sales, if size is correlated to other characteristics which are relevant to firms' cyclical behaviour. For example, if small firms are concentrated in sectors with more cyclical product demand or are marginal suppliers to other, maybe larger, firms or to the market, they may well be more severely hit by recessions that large firms.

To control for alternative shocks and different adjustment mechanisms among firms, some recent studies use microeconomic firm-level data.[1] Kashyap *et al.* (1994) provide a case-study of the behaviour of inventory accumulation focusing on the 1982 recession and using a cross-section of US firms. Their main finding is that inventory investment of firms without access to public capital markets is significantly liquidity-constrained during the recessionary episode. Gilchrist and Zakrajšek (1995) and Bernanke *et al.* (1996) analyse firm-level data for US manufacturing firms, confirming the findings of substantial cross-sectional differences between borrowers with different possibilities of financial market access.

The present work contributes to the latter strand of empirical literature, providing a case-study of inventory investment behaviour during the recession experienced in the early 1990s by several European countries including France, Italy and the UK. As Fig. 1 shows, in 1993 the rate of change of GDP was negative in France (−1.3%) and Italy (−0.9%), whereas in the UK recession reached its trough in 1991 (−1.5%). Industrial production followed the same pattern with even more pronounced fluctuations. In all countries industrial production decreased for three consecutive years: from 1991 to 1993 for France and Italy (more sharply in the former country), with the most serious decline in 1993 (−4.6% and −2.4%

[1]Alternatively, other studies use time-series data on firms disaggregated by geographical region (Carlino and De Fina (1998) for the USA) or by sector (Dedola and Lippi (2000) for the USA and four large European economies).

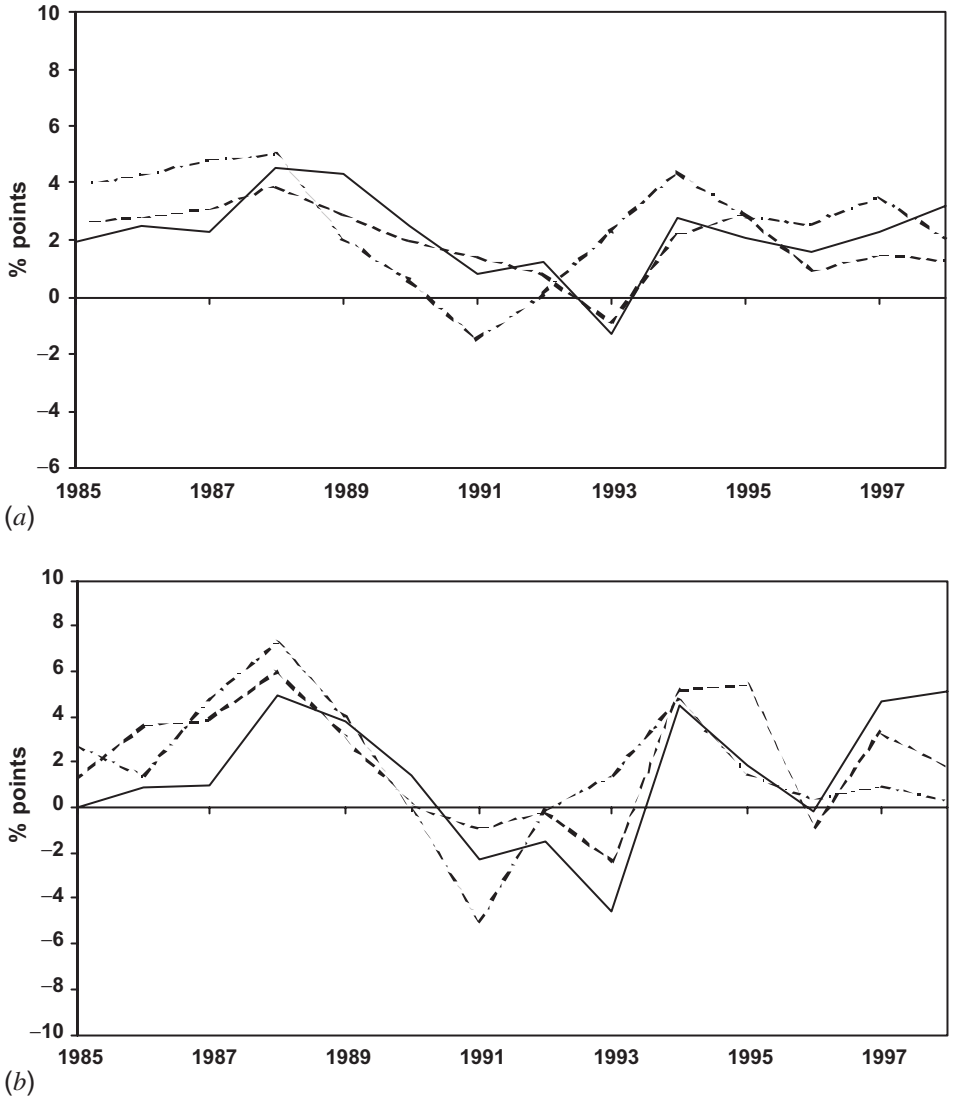

Fig. 1. *Annual growth rates of GDP and industrial production (1985–1998) (a) GDP annual growth rates, (b) industrial production annual growth rates: (——) France, (- - - - -) Italy, (-·-·-·-·-) UK*

respectively), and from 1990 to 1992 in the UK, with the sharpest fall in 1991 (−5%). In Italy, industrial production decreased also in 1996 by 1%. The main recessionary episode in the early 1990s follows in all countries a tightening of monetary policy, occurred in France and Italy in the second half of 1992 and the first quarter of 1993 and in the UK between the end of 1990 and the beginning of the following year. Other policy actions may have accompanied, if not induced, the recession, especially in Italy, where also a severe fiscal contraction occurred in those years (and occurred again in 1996, when a fiscal policy tightening was required to fulfil EMU entry conditions).

An additional motivation for the cross-country analysis is provided by the current debate on the effects and transmission channels of the common monetary policy in the EMU area (see Guiso *et al*. (1999) for a comprehensive survey of empirical results). A detailed analysis of firms' inventory response to adverse macroeconomic conditions may be useful in order to understand the specific channels of monetary policy transmission in each EMU country, providing valuable information on potential asymmetries of ECB's policy actions. The evidence so far available, mainly based on large-scale national macroeconomic models or smaller-scale structural VAR systems, does not yield a consistent set of results on the differences and similarities in the monetary transmission channels across countries. Disaggregated, firm-level data seem more likely to deliver harder evidence on this issue, since they exploit the advantage of much more variability in firms' behaviour at the cross-sectional level than at the cross-country level.

To this aim, a panel of French, Italian and UK firms is used to look first at the behaviour of inventories during

Table 1. *Descriptive statistics on real sales*

	France		Italy		UK	
Number of firms	2093		2254		1560	
Number of observations	18 837		15 778		14 040	
Sample period	1989–1997		1991–1997		1989–1997	
Statistics on real sales:						
Period	1989–1997	1989	1991–1997	1991	1989–1997	1989
Mean	60.00	58.27	52.09	47.51	224.36	226.06
First quartile	13.50	12.55	14.24	12.52	16.81	16.01
Median	21.98	20.78	20.61	17.86	30.97	29.52
Third quartile	48.44	46.74	36.96	31.45	78.51	74.97

Note: Sales are measured in millions of ECUs at constant exchange rates and constant prices (base year: 1990).

the recession of the early 1990s both in aggregate and for different classes of firms.[2] Inventories are the real component of firms' assets that is likely to be more responsive to financial constraints; in fact, the financial pressure caused by a negative shock affects all firms' assets, but the relative lower liquidation and adjustment costs of inventories determine their potentially larger response to recessionary shocks.

III. DATA DESCRIPTION

Annual balance sheet data at the firm level for three large European economies (France, Italy and the UK) is used, obtained from Amadeus, a commercial data-bank containing (unconsolidated and/or consolidated) balance sheet information on European firms.[3] Using data from a common source should guarantee a reasonable degree of comparability across countries with partly different accounting practices and classifications of elementary balance sheet items. Available data start in 1989 for France and the UK, and in 1991 for Italy; the end of the sample is 1997 for all countries.

A balanced panel has been constructed by considering only those firms in the manufacturing sector with continuous observations on the relevant variables throughout the sample period. To avoid outliers, firms with anomalous observations on the crucial variables (sales, inventories and various measures of financial pressure) have been discarded.[4] Overall, the sample contains data from around 6000 firms (2093 from Italy, 2254 from France and 1560 from the UK), for a total of almost 49 000 observations.[5] For comparison, the scant empirical evidence on inventory behaviour in European countries is either based on smaller panels of publicly traded UK companies (Guariglia, 1999, 2000) or on time series aggregates of small and large Italian firms (Rondi *et al*., 1998). Cross-country comparisons of inventory investment are not available, whereas Bond *et al*. (1997) provide a cross-country analysis of firms' fixed investment behaviour in the 1980s using data from 361 Belgian firms (quoted and non-quoted), 1365 French firms (quoted and non-quoted), 228 quoted German firms, and 571 UK firms (consolidated and quoted), with remarkable size heterogeneity, due to the larger average size of German and UK firms.

Basic descriptive statistics on the distribution of firms in the sample by size (measured by real sales) are reported in Table 1 for the whole time period and for the initial year in each sample. The data shows that the size distribution of French and Italian firms is quite similar, and only in the sample for the UK there is a greater proportion of relatively large firms.[6] Besides having broadly comparable numbers of firms and size distributions across countries, the sample has the advantage of including both quoted and non-quoted firms, therefore covering also the lower tail of the overall firms' size distribution in each country where most firms are unlisted.

[2]Also Bond *et al*. (1997) provide a cross-country study of a panel of European firms, analysing the role of financial factors in affecting firms' investment behaviour, but with no specific focus on recessionary episodes.
[3]Higher frequency data (e.g. quarterly as in the empirical analysis of US firms by Carpenter *et al*, 1994, 1998), though more desirable to capture the interaction between cyclical inventory fluctuations and financial factors, are unavailable for most European countries.
[4]The Appendix provides more details on the treatment of the data and some discussion of our sample building strategy.
[5]In preliminary work, also a sample of German firms had been constructed from the same data source. Data for only 325 firms are available and the size distribution is not comparable with the other three countries in the sample (e.g. the median German firm has real sales three times as large as the median UK firm). It was therefore decided to exclude German firms from the investigation.
[6]This may be partly due to the presence for about 50% of UK firms only of their consolidated balance sheets (see the Appendix for details).

Table 2. *Descriptive statistics on inventories and leverage measures (whole sample)*

	France	Italy	UK
Inventories			
Mean	0.134	0.174	0.133
First quartile	0.072	0.092	0.073
Median	0.118	0.149	0.123
Third quartile	0.177	0.227	0.179
Leverage			
Mean	0.634	0.734	0.588
First quartile	0.519	0.650	0.454
Median	0.647	0.760	0.591
Third quartile	0.762	0.844	0.725
Short-term leverage			
Mean	0.498	0.558	0.452
First quartile	0.371	0.436	0.322
Median	0.492	0.566	0.432
Third quartile	0.618	0.683	0.565
Debt maturity			
Mean	0.786	0.753	0.779
First quartile	0.700	0.663	0.667
Median	0.818	0.774	0.831
Third quartile	0.907	0.864	0.930

Note: 'Inventories' are expressed as a ratio to sales; 'leverage' is computed as the ratio of total debt (short- and long-term) to total liabilities (debt and shareholders' funds); 'short-term leverage' is computed as the ratio of short-term debt to total liabilities; 'debt maturity' is computed as the ratio of short-term debt to total debt.

Additional descriptive statistics, reported in Table 2, concern the variables used below in the empirical analysis: inventories (as a ratio to sales), and three measures of financial pressure, such as leverage (the ratio of short- and long-term debt to total liabilities, i.e. debt and shareholders' funds), short-term leverage (the ratio of short-term debt to total liabilities), and debt maturity (the ratio of short-term debt to total debt). On average in the sample period Italian firms display higher leverage measures (both total and short-term), whereas the debt maturity indicator is quite similar across countries. Finally, also the distribution of the inventories–sales ratio is similar across countries, with Italy displaying slightly higher values.

For each country, the analysis of firms' inventories behaviour is carried out first using all firms in the sample. Then, the focus is on the differential behaviour of firms with different capital market access. This should be useful to capture additional responses to recessionary shocks, since firms with more difficulties in raising funds on the market should be more severely affected in periods of downturn in economic activity. In order to capture differential financial market access, the sample in each country is partitioned according to two criteria. The first is based on the previously described *size* distribution: firms in each country are partitioned into two dimensional classes ('large' and 'small' firms) using a common size threshold of 20 millions ECU at 1990 prices (which is close to the lowest median firm size among the three countries) applied to the distribution of real sales in the first year of the sample (1989 for France and the UK, 1991 for Italy). Therefore, firms are not allowed to move from one dimensional class to the other during the period. Though somewhat restrictive, this choice may be justified both by the relatively short time span analysed (9 years at most) and by a focus on a specific episode occurring towards the beginning of the sample period.[7] As discussed by Gertler and Gilchrist (1994), firm size is a reasonable proxy for capital market access since it is strongly correlated with factors, such as the degree of idiosyncractic risk, the availability of collateral and the existence of a bond rating, more directly relevant in determining the existence and magnitude of the premium on external finance.[8]

Firms in each country are also partitioned according to their *age*, as revealed by the year of foundation: firms with more than ten years of age in 1989 are defined as 'old' and those with less than ten years as 'young'. In so doing, again it is not allowed for firms to transit from young to old during the sample period.[9] Age may capture better than size the firms' track record, a relevant information in determining the availability and cost of external finance. This may be particularly true for Italy, where many small firms are relatively old and have developed over time long-run relationships with financial intermediaries (typically, commercial banks), overcoming informational problems.[10] It is noted here that working with a balanced panel may bias the results against finding significant differences between firms of different size if smaller or younger firms are more likely to default due to adverse financial conditions during recessions.

[7]The dimensional split among firms may be obtained using other criteria: e.g. by using the median of the real sales distribution in each country as the threshold or by defining those firms in the upper third of the size distribution as 'large' and those in the lower third as 'small' (in the following empirical analysis both such alternative split criteria were applied with qualitatively very similar results).
[8]Moreover, splitting the sample by firm size as a way of identifying crucial effects is a widely used technique in the literature on investments and financial constraints (Hubbard, 1998; Schiantarelli 1995).
[9]Devereux and Schiantarelli (1989) define as 'old' the UK *quoted* firms with at least 12 years of age; Carpenter and Rondi (2000) classify as 'old' those Italian firms with more than 15 years of age.
[10]Furthermore, as suggested by Carpenter and Rondi (2000), the peculiar ownership structure of a large fraction of Italian firms, based on long-lasting family control, represents a constraint on firms' growth, reducing the correlation between size and age.

Applying the size and age criteria to the sample produces four groups of firms whose main features are described in Table 3. Each cell in panel A of the table reports the number of firms, the average size (as measured by real sales) over the whole sample period, and the average age in years measured in 1989. In all countries there is a sizeable dispersion of firms across different groups. In particular, the number of firms in the off-diagonal cells (including the 'small and old' and the 'large and young' firms) is around 46% in France, 53% in Italy and even 38% in the UK, where the 'large and old' firms are a big share of the overall sample (55%). It is concluded that the partition into 'size' and 'age' classes provides significantly different information on firms and may capture their ability to access capital markets in a different (and complementary) fashion. Then, using both split criteria in the empirical analysis can yield valuable additional information. As for firms' dimension, the common cut-off across countries produces a size split broadly consistent with some of the main features of national industrial structures: 'small' firms, which amount to only 35% in the UK, are 47% in France and 56% in Italy, reflecting a well-known peculiarity of the Italian manufacturing sector. Moreover, more than one third of Italian firms are 'young', against less than 20% in both other countries. The fraction of 'young' firms is higher in the 'small' category than in the 'large' group in France and in the UK (around 20% compared with around 15%) but not in Italy (34% against 35%). The average age (in 1989) of 'young' firms is uniform across countries (4–5 years), whereas the average age of the 'old' firms ranges from 26 in Italy to 45 in the UK, with French firms in between (36).

Panel B of Table 3 reports the (whole period) average of inventories and leverage measures for the four firms' groups in each country. In all samples firms display remarkably similar structural features across groups. This similarity ensures that potential different behaviour of firms belonging to distinct groups cannot simply be attributed to their different asset/liability structure. Only the 'small and young' (S/Y) group shows some systematic differences in the leverage distribution, consistent across all three leverage measures.

Finally, Fig. 2 focuses on the dynamics of the main variables of interest showing for each country the rate of change of the median of the distribution of sales and inventories (measured in real terms) over the sample period. Both variables display a strongly procyclical dynamics, with slightly more pronounced fluctuations for inventories in France and, to a lesser extent, in the UK. This firm-level evidence is broadly consistent with the stylized facts on

Table 3. *B. Descriptive statistics for firms' groups (mean and standard deviation in brackets)*

		L/O	L/Y	S/O	S/Y
France	Inventories	0.134	0.146	0.135	0.121
		(0.090)	(0.095)	(0.093)	(0.093)
	Leverage	0.622	0.647	0.634	0.669
		(0.172)	(0.170)	(0.172)	(0.157)
	Short-term leverage	0.486	0.492	0.504	0.530
		(0.171)	(0.177)	(0.167)	(0.170)
	Debt maturity	0.781	0.761	0.796	0.793
		(0.157)	(0.175)	(0.145)	(0.159)
Italy	Inventories	0.175	0.180	0.174	0.169
		(0.126)	(0.130)	(0.123)	(0.122)
	Leverage	0.720	0.732	0.732	0.763
		(0.151)	(0.142)	(0.144)	(0.148)
	Short-term leverage	0.540	0.543	0.556	0.601
		(0.166)	(0.173)	(0.166)	(0.182)
	Debt maturity	0.742	0.736	0.754	0.781
		(0.139)	(0.159)	(0.142)	(0.156)
UK	Inventories	0.135	0.128	0.133	0.125
		(0.079)	(0.080)	(0.083)	(0.083)
	Leverage	0.580	0.616	0.576	0.653
		(0.185)	(0.172)	(0.202)	(0.187)
	Short-term leverage	0.441	0.475	0.455	0.484
		(0.175)	(0.171)	(0.184)	(0.173)
	Debt maturity	0.771	0.779	0.802	0.756
		(0.189)	(0.184)	(0.175)	(0.195)

Table 3. *A. Sample composition by size and age*

		France			Italy			UK		
		Large	Small	Total	Large	Small	Total	Large	Small	Total
Old	Number of firms	903	790	1693	638	844	1482	854	436	1290
	Size (mill. ECU)	99.5	14.4	59.8	107.4	16.2	55.5	366.4	16.0	247.9
	Age (years)	39	33	36	29	24	26	48	39	45
Young	Number of firms	180	220	400	348	424	772	159	111	270
	Size (mill. ECU)	115.7	15.8	60.8	80.0	17.2	45.5	177.1	18.0	111.7
	Age (years)	4	4	4	5	5	5	4	5	5
Total	Number of firms	1083	1010	2093	986	1268	2254	1013	547	1560
	Size (mill. ECU)	102.2	14.7	60.0	97.8	16.6	52.1	336.7	16.4	224.4
	Age (years)	33	27	30	20	17	19	41	32	38

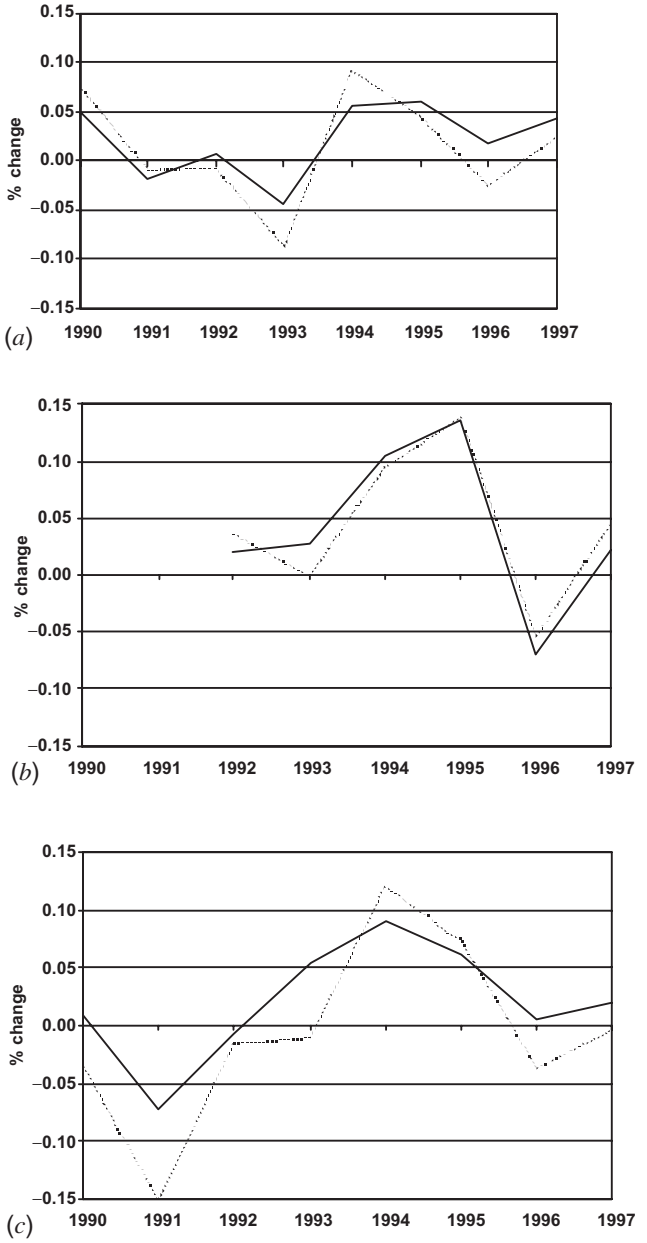

Fig. 2. *Real growth rates of sales and inventories (medians) (a) France, (b) Italy, (c) UK: (——) sales, (·····) inventories*

aggregate inventory behaviour reported by Ramey and West (1999).

IV. EMPIRICAL SPECIFICATION

In this section the specification of the estimated inventory equation is motivated and relevant estimation issues briefly discussed.

The following general autoregressive distributed lag model for inventories and sales (as discussed e.g. by Blinder and Maccini, 1991), augmented by financial variables as in Carpenter *et al.* (1994, 1998) and Guariglia (1999) is estimated:

$$inv_{it} = \beta_1 inv_{it-1} + \beta_2 sales_{it} + \beta_3 sales_{it-1} - \beta_4 lev_{it-1} + \alpha_i + \alpha_t + \varepsilon_{it} \quad (1)$$

where the dependent variable is the (log of) the end-of-period stock of real inventories (inv_{it}), $sales_{it}$ is (the log of) real sales and lev_{it-1} denotes (the log of) firm i's leverage as measured at the beginning of period t. The error term in the equation includes a firm-specific fixed effect (α_i) which captures any time-invariant influence of unobservable variables on individual firms' behaviour due, e.g., to different storage costs and rate of obsolescence across firms, a time effect (α_t), and an idiosyncratic component (ε_{it}). Guariglia (1999) estimates an error-correction version of Equation 1, with additional dynamics, on UK data. For the USA, Carpenter *et al.* (1994) derive a similar equation from a structural model with stock adjustment and a buffer-stock role for inventories, and interpret the estimated coefficients accordingly. For the purposes of this paper, lagged inventories and current and lagged sales may be considered simply as control variables allowing to test for the relevance of financial factors; therefore a structural interpretation is not attributed to the first three estimated coefficients.

The leverage term is included in the equation to test whether inventory investment is sensitive to balance sheet variables proxying for the degree of financial pressure (in the empirical analysis below all the previously defined measurers are used as indicators of financial pressure). Other variables capturing the same effect are commonly employed in the literature, namely the beginning-of-period stock of liquid assets (cash and marketable securities) as in Kashyap *et al.* (1994), the coverage ratio (the ratio of pre-tax and pre-interest earnings to total interest payments, interpretable as the flow counterpart of leverage, see Nickell and Nicolitzas, 1999) as in Guariglia (1999) and Carpenter *et al.* (1998), and particularly cash flow as in Carpenter *et al.* (1994, 1998). Since the data set does not allow for construction of a reliable cash flow measure for all countries, (total and short-term) leverage is used as a proxy for financial pressure. It is believed that leverage may nevertheless be a suitable variable to consider for two main reasons. First, cash flow may also contain information on firms' future profitability not fully accounted for by the sales variable (even though inventories should react more to short-term profit expectations likely to be more closely correlated to sales); in turn this would determine an upward bias in the cash flow estimated coefficient. Second, cash flow and sales may be highly collinear, making it difficult to estimate relevant coefficients

with precision.[11] Both problems are less likely to apply to a leverage measure.

Controlling for sales allows the estimated α_t to be interpreted as evidence of fluctuations in inventories in addition to the firm-level business cycle. These additional effects may well be an important part of a financial propagation mechanism amplifying the impact of recessionary shocks and restrictive policy actions. Indeed, in the time period under study all countries suffered from a pronounced demand-driven cyclical downturn at least partly attributable to restrictive monetary and fiscal policies; however, from the perspective of this paper, the identification of the structural source of shocks originating the recessionary episode is not a crucial issue. The time dummies then capture the common cyclical response of inventory investment during recession which is independent from the firm-specific sales' fluctuations.[12]

Estimation issues

In order to remove the firm-specific effect all equations are estimated in first differences. A constant term is kept in estimation to allow for a possible time trend in the levels of the dependent variables. The estimation period is therefore 1991–1997 for France and the UK and 1993–1997 for Italy, since two observations are lost by lagging the variables and by constructing first differences. Estimation is carried out by the Generalized Method of Moments (GMM) using twice or more lagged variables in levels as appropriate instruments for the transformed lagged dependent variables. The assumption of no serial correlation in ε_{it} is essential for the consistency of the GMM estimator; if the disturbances are not serially correlated there should be evidence of first-order serial correlation in differenced residuals but no evidence of second-order correlation (see Arellano and Bond, 1991, 1998). For this reason in Table 4 and 5 the results of first-order and second-order residual serial correlation tests (m_1 and m_2 denote the p-values of the relevant test statistics) and the Sargan tests of over-identifying restrictions are reported. In all equations its also allowed for the sales variables to be predetermined, using twice or more lagged values as instruments.[13] In the following section results for the aggregate samples and for four groups of firms, resulting from the joint application of our two splitting criteria (dimensional and age), are reported.

V. RESULTS

Results from the estimation of inventory equations 1 are shown in Tables 4 and 5. Table 4 shows results for the aggregate samples and Table 5 reports for each country results for the four subsamples of firms: 'large and old', 'large and young', 'small and old' and 'small and young'. For each equation estimated coefficients are reported together with their standard errors; moreover, the p-values of test statistics for first- and second-order residual serial correlation (m_1 and m_2) are displayed. In addition, w_t denotes the p-value of the statistic (with a $\chi^2(6)$ distribution for France and the UK and a $\chi^2(4)$ distribution for Italy) testing the joint significance of the time dummies α_t. The estimated values for the time dummies with 95% confidence intervals are plotted against time in Fig. 3 for the aggregate sample, for the years 1991–1996 for France and the UK and 1993–1996 for Italy. When estimation is carried out on firms' subsamples, the four sets of estimated α_t's are portrayed in Fig. 4.

Time dummies coefficients measure the percentage deviation of the dependent variable in each year from its value in 1997, after controlling for the behaviour of sales and for the effect of the leverage measure. Moreover, since a constant term is included in (first-difference) estimation, a linear time trend effect (if present) is removed from the magnitude of the estimated coefficients. Such trend could capture gradual improvements in inventory management (e.g. the diffusion of a 'just-in-time' technology). Year 1997 has been chosen as a benchmark for the evaluation of time effects since it is the last observation in the sample and because in that year the growth rates of GDP and industrial production were fairly close in the countries as shown in Fig. 1. For all countries the baseline specification in Equation 1 has been expanded by introducing an additional variable constructed by interacting each financial variable (either lev_{it-1}, $stlev_{it-1}$ or mat_{it-1}, denoting total leverage, short-term leverage and debt maturity

[11]Moreover, the lack of data on 'cash' held by firms in the data set precludes the computation of a reliable measure of liquid assets. For the same reason, it was not possible to construct a measure of 'net leverage' by subtracting cash and other liquid assets to both total debt and total liabilities.

[12]If forward-looking firms are subject to common shocks to sales, so that current sales matter also because they predict the future, the time effect α_t must also pick up a common component in future expected sales not captured by the coefficients on current and lagged sales.

[13]It is well known that in dynamic panel data models where the autoregressive parameter is large and the number of time series (N) is small the first difference GMM estimator suffers from finite sample bias and poor precision, as shown in simulation studies, e.g. Blundell and Bond (1998). However, this is unlikely to be a problem in the present case since N is very large, ranging from 1560 to 2254 in the aggregate estimates, and inventories are known not to be a very persistent phenomenon (Ramey and West, 1999). In fact, Blundell and Bond (1998) find that the sample bias becomes negligible for $N = 500$ and a true autoregressive parameter around 0.8.

Table 4. *Inventory equations: Aggregate results (dependent variable:* inv_{it}*; standard errors in parentheses)*

	France	Italy	UK	France	Italy	UK	France	Italy	UK
inv_{it-1}	0.337	0.512	0.365	0.346	0.520	0.369	0.346	0.525	0.382
	(0.059)	(0.057)	(0.114)	(0.060)	(0.057)	(0.114)	(0.061)	(0.057)	(0.114)
$sales_{it}$	0.614	0.480	0.691	0.625	0.435	0.728	0.643	0.371	0.646
	(0.146)	(0.197)	(0.262)	(0.148)	(0.196)	(0.269)	(0.147)	(0.197)	(0.263)
$sales_{it-1}$	0.244	0.242	0.165	0.238	0.215	0.164	0.228	0.224	0.220
	(0.119)	(0.127)	(0.160)	(0.121)	(0.127)	(0.172)	(0.120)	(0.127)	(0.167)
lev_{it-1}	−0.315	−0.564	−0.283	–	–	–	–	–	–
	(0.035)	(0.063)	(0.037)						
$lev_{it-1} \cdot R_t$	0.014	−0.044	0.007	–	–	–	–	–	–
	(0.022)	(0.022)	(0.025)						
$stlev_{it-1}$	–	–	–	−0.135	−0.247	−0.203	–	–	–
				(0.021)	(0.030)	(0.028)			
$stlev_{it-1} \cdot R_t$	–	–	–	0.007	−0.027	−0.002	–	–	–
				(0.017)	(0.016)	(0.020)			
mat_{it-1}	–	–	–	–	–	–	−0.043	− 0.160	− 0.058
							(0.021)	(0.031)	(0.030)
$mat_{it-1} \cdot R_t$	–	–	–	–	–	–	−0.016	− 0.019	− 0.004
							(0.022)	(0.027)	(0.046)
m_1	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00
m_2	0.47	0.09	0.03	0.47	0.10	0.03	0.45	0.10	0.03
Sargan	0.01	0.18	0.01	0.01	0.24	0.01	0.01	0.06	0.01
w_t	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00	0.00

Note: Sample period: 1991–1997 for France and the UK; 1993–1997 for Italy. All models are estimated in first differences. Time dummies are included in all equations both as regressors and as instruments. One-step standard errors (robust to heteroscedasticity) are reported in parentheses. Instruments: inv_{it-2}, ..., $sales_{it-2}, \ldots, \Delta lev_{it-1}$ or $\Delta stlev_{it-1}$ or Δmat_{it-1}, $\Delta(lev_{it-1} \cdot R_t)$ or $\Delta(stlev_{it-1} \cdot R_t)$ or $\Delta(mat_{it-1} \cdot R_t)$. m_1 and m_2 are tests for first- and second order residual correlation, asymptotically distributed as $N(0, 1)$ under the null of no serial correlation. w_t is a Wald test for the joint significance of the time dummies (for all tests p-values are reported).

respectively) with a dummy variable R_{it} taking the value of 1 in 1991–1992–1993 for France, in 1993 and 1996 for Italy and in 1991 and 1992 for the UK. This new variable is designed to test whether the disciplining role of debt is stronger in recession.

The aggregate results in Table 4 show a similar coefficients' pattern on the sales and lagged inventories variables across countries. When the error-correction restriction (as in Guariglia, 1999) is imposed in estimation of Equation 1, it is always strongly rejected on the aggregate data (and in the majority of disaggregated estimates). The tests on the residuals show the expected first-order serial correlation and detect some evidence of second-order correlation only for the UK sample. However, the Sargan tests reveal some problem with the instruments used in the equations for France and the UK. The set of time dummies is always statistically significant at the 1% level and again the coefficients' time pattern (Fig. 3) shows a marked cyclical behaviour in all countries, after controlling for firm-specific fluctuations in sales and beginning-of-period financial pressure. The magnitude of this effect is −8% in France in 1993 and around −2–3% for the UK in 1991 and for Italy in 1993. In this latter country, the inventory decumulation is much stronger in 1996 (−9%). The effect of the three financial pressure measures is consistently negative and strongly statistically significant in all countries, with a larger magnitude in Italy (more evidently so for the total leverage and debt maturity measures). In the aggregate samples the only additional recessionary effect is found for Italy when the total leverage variable is used. To gauge the quantitative importance of this effect we performed a very simple experiment. The impact on inventories of a leverage increase from 0.55 (approximately the first quartile of the overall distribution of firms in the sample) to 0.75 (approximately the third quartile) was computed for each country. According to the estimates, this increasing leverage implies a reduction in inventories of 11.5% for the UK, 10.3% for France and 20.5% (21.1% in recessionary years) for Italy.

Table 5 shows the results of the inventory equations estimated on the four available size/age firms' groups, using the leverage as a measure of financial pressure. Rather comfortingly, there is no evidence of second-order serial correlation in all subsamples at the 5% significance level. Moreover, the Sargan tests reject the choice of instruments only in three out of 12 subsamples. The strong negative effect of leverage is confirmed also for all subsamples. No appreciable difference across firms' groups is found

Table 5. *Inventory equations: subsample results (dependent variables:* inv_{it}*; standard errors in parentheses)*

	Large/Old	Large/Young	Small/Old	Small/Young
Panel A: France				
inv_{it-1}	0.308	0.472	0.334	0.401
	(0.072)	(0.076)	(0.070)	(0.137)
$sales_{it}$	0.585	0.669	0.480	0.855
	(0.309)	(0.202)	(0.167)	(0.318)
$sales_{it-1}$	0.121	0.170	0.189	−0.185
	(0.190)	(0.218)	(0.133)	(0.178)
lev_{it-1}	−0.297	−0.261	−0.288	−0.314
	(0.060)	(0.097)	(0.062)	(0.099)
$lev_{it-1} \cdot R_t$	0.051	0.060	−0.027	−0.023
	(0.042	(0.071)	(0.025)	(0.071)
m_1	0.00	0.00	0.00	0.00
m_2	0.69	0.11	0.40	0.29
Sargan	0.01	0.28	0.11	0.24
w_t	0.16	0.48	0.00	0.57
Panel B: Italy				
inv_{it-1}	0.519	0.397	0.531	0.399
	(0.101)	(0.120)	(0.095)	(0.100)
$sales_{it}$	0.498	0.804	0.044	0.757
	(0.353)	(0.329)	(0.205)	(0.425)
$sales_{it-1}$	0.150	0.579	0.231	−0.128
	(0.232)	(0.245)	(0.164)	(0.233)
lev_{it-1}	−0.349	−0.598	−0.638	−0.503
	(0.109)	(0.152)	(0.109)	(0.127)
$lev_{it-1} \cdot R_t$	0.053	0.001	−0.030	−0.107
	(0.038)	(0.050)	(0.037)	(0.052)
m_1	0.00	0.00	0.00	0.00
m_2	0.30	0.06	0.56	0.25
Sargan	0.21	0.47	0.71	0.63
w_t	0.00	0.02	0.00	0.23
Panel C: UK				
inv_{it-1}	0.138	0.563	0.162	0.149
	(0.108)	(0.080)	(0.152)	(0.082)
$sales_{it}$	0.287	1.279	0.828	0.853
	(0.237)	(1.163)	(0.413)	(0.404)
$sales_{it-1}$	0.234	−0.250	−0.083	0.008
	(0.151)	(0.240)	(0.157)	(0.283)
lev_{it-1}	−0.159	−0.608	−0.229	−0.198
	(0.051)	(0.201)	(0.060)	(0.102)
$lev_{it-1} \cdot R_t$	−0.031	−0.004	−0.008	−0.008
	(0.022)	(0.099)	(0.055)	(0.069)
m_1	0.00	0.02	0.00	0.00
m_2	0.70	0.33	0.54	0.37
Sargan	0.00	0.66	0.02	0.66
w_t	0.00	0.00	0.00	0.19

Note: See Table 4, Instruments: $inv_{it-2}, \ldots$, $sales_{it-2}, \ldots$, Δlev_{it-1}, $\Delta\ (lev_{it-1} \cdot R_t)$.

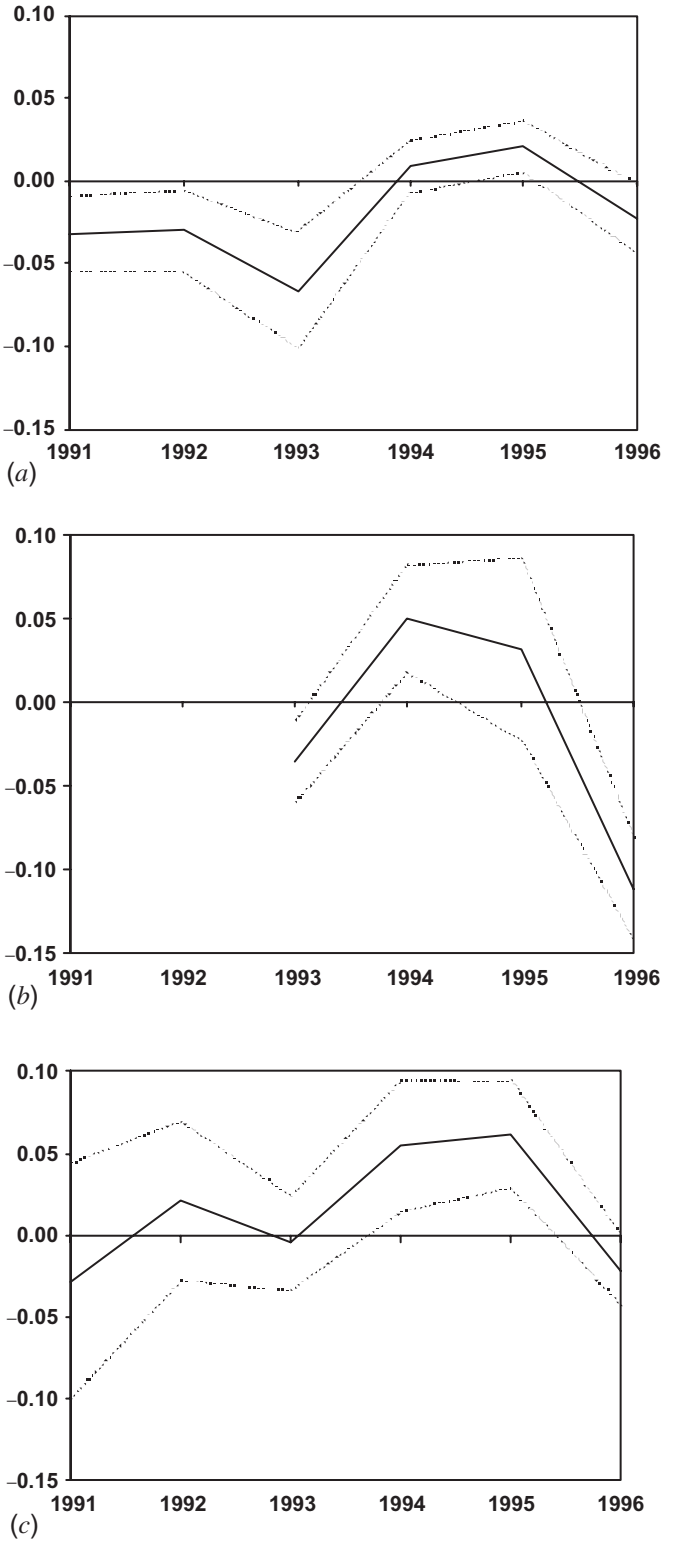

Fig. 3. *Coefficients on time dummies in equations for inventories (with 95% confidence intervals) (a) France, (b) Italy, (c) UK*

in France whereas 'small' UK firms (both 'old' and 'young') show a larger leverage effect than the 'large and old' firms.[14] In Italy the 'large and old' firms display a smaller leverage effect with respect to the other

[14]Furthermore, the subsample of 'large and young' UK firms shows a larger coefficient on leverage compared to the other subsamples. However, this result has to be taken with caution given both the limited sample size (159 firms) and the suspicious deformity in the magnitude of other coefficients.

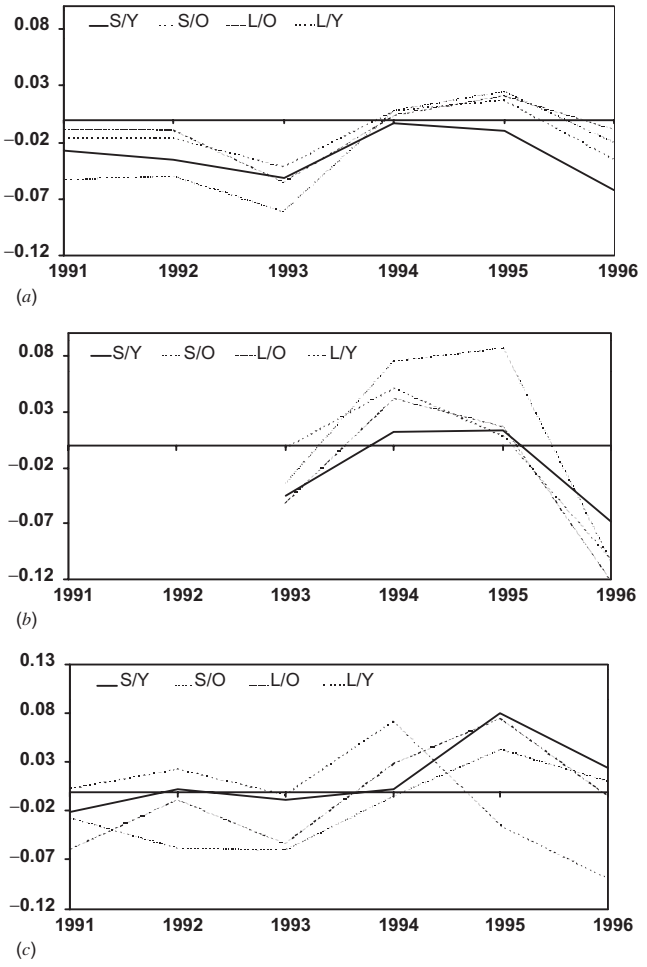

Fig. 4. *Coefficients on time dummies in equations for inventories for subsamples of firms (a) France, (b) Italy, (c) UK*

three groups. In addition, only for the 'small and young' Italian firms the sensitivity to financial pressure is magnified in recession years, with a negative and statistically significant estimated coefficient on the interaction variable $lev_{it-1} \cdot R_t$. Overall, when some within-country differences are detected in inventories behaviour, they are broadly consistent with a larger leverage effect for firms more likely to suffer from financial pressures.

Finally, as can be seen from Fig. 4, the pattern of the time dummies α_t is quite similar across different firms' groups in all countries, with the partial exception of the 'large and young' UK firms. As suggested by Gertler and Gilchrist (1994), if small or young firms use more flexible production technologies, a more pronounced reduction in inventories may not be due to financial factors but simply to a greater possibility of quickly adjusting inventories when production needs change. Allowing for different coefficients also on the variables capturing technological features (sale and lagged inventories) as well as different time trends, this possibility is explicitly taken into account when splitting the sample into four groups of firms.

To assess the robustness of the above results the two additional measures of firms' financial pressure are employed, namely short-term leverage and debt maturity, obtaining results not substantially different from those reported above. Moreover, given the annual frequency of the data, a different specification for the recession dummy was used, focusing on the 'worst' year for each country. Therefore R_t was defined as taking the value of one only in 1993 for France and Italy and only in 1991 for the UK. Moreover, a second-order autoregressive specification has been estimated, to evaluate whether the pattern of time dummies displayed in Figs. 3 and 4 is attributable to functional-from misspecification. Both extensions of the basic model yielded results not appreciably different from the one reported in Tables 4 and 5. Finally, the model was also re-estimated without imposing the logarithmic transformation on the financial variables, yielding qualitatively very similar results.

In conclusion, the results show that a leverage measure is able to capture financial pressure effects in all countries, which are stronger in Italy where smaller and younger firms seems to be hit more sharply in recessions. Moreover, in the aggregate, inventories display a procyclical pattern in excess of what is explained by firm-specific fluctuations in sales and this sensitivity is of a sizeable magnitude especially in France and Italy. On the whole, the empirical results suggest that although a common pattern clearly emerges, some interesting differences can also be found. In particular, cross-country differences in inventory behaviour seem to be more pronounced than differences between firms grouped by size and age within each country, with the notable exception of smaller and younger Italian firms.

VI. CONCLUSIONS

The main conclusion from the empirical analysis on firm-level data for three large European economies in the early 1990s is that in all countries, during the recessionary episode, firms reduced inventories significantly beyond the level justified by the cyclical behaviour of sales. To the extent that this recession was triggered by monetary policy tightening, these findings may support the broad 'financial accelerator' view of the monetary transmission mechanism. Moreover, a significant negative effect of the level of leverage is found on inventories in all countries and in all subsamples within each country. This in turn suggests that all types of firms respond to financial pressures by reducing the level of inventories. However, besides these common patterns, also some differences both across countries and within each country between various classes of firms (with a different access to financial markets) seem to emerge. In particular strong evidence is found that the leverage effect on inventories is significantly larger for Italian firms;

in addition, this effect is enhanced in recessions for the sizeable group of 'smaller and younger' Italian firms. From the European Monetary Union perspective, the evidence points towards broadly similar firms' responses to financial pressures, but with sizeably different magnitude across countries.

Clearly additional work on these issues is needed. For example, a more complete picture of the effects of recessions on firms' behaviour could be obtained by the investigation of the cyclical reaction of other items on both sides of firms' balance sheet (in particular trade credit, net trade debt, fixed investment and short- and long-term debt) and by considering potential cross-country differences in firms' cyclical behaviour at the sectoral level. Even though empirical work in this area is at present somewhat hampered by the limited availability of more exhaustive European firm-level data sets, it ranks high in the research agenda.

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APPENDIX

The data used in this paper is obtained from the cd-rom version of the Amadeus data-bank produced by Bureau Van Dijk (BVD, hereafter). Each cd-rom contains (unconsolidated and/or consolidated) balance sheet information on European manufacturing firms for the five most recent years for which data were available at the time of publication. To broaden the sample period use was made of two different cd-roms covering respectively the 1989–1983 (but only 1991–1993 for Italy) and 1993–1997 periods.

To build the panels the following initial procedures have been followed. First, firms located in France, Italy and the UK were identified separately in each cd-rom and (unconsolidated and/or consolidated) balance-sheet data were extracted. Second, balance sheet data coming from the two cd-roms were merged by using the BVD identification number as key variable and by exploiting the 1993 data (available in both cd-roms) as a further consistency check. Third, since in a limited number of cases both the unconsolidated and the consolidated versions of the balance-sheets were available, the latter one was discarded. Fourth, only firms for which balance-sheets (possibly with missing values) were available for the whole sample period, that is 1989–1997 for France and the UK and 1991–1997 for Italy, were kept into the panel. The authors are more than aware that the choice of imposing a balanced structure to the panel might introduce a sample selection bias. However, given the characteristics of the data source, the only available alternative choice would have been to include also those firms present in the second (1993–1997) but not in the first (1989–1993) cd-rom (being small the number of firms present in the first but not in the second cd-rom). However, this sample building strategy would have introduced a different, and possibly more serious, sample selection bias. In addition, it would have not been very helpful in the estimation of the leverage effect in recession, since data on recessionary years are concentrated in the first cd-rom. As it can be seen in Table A1, after these initial procedures 2751 firms were left (and therefore 24 759 firm-year observations) for France, 2581 firms (and 18 067 observations) for Italy and 2869 firms (and 25 821 observations) for the UK.

The second step consisted in removing from the initial samples firms with missing values for the variables used in the econometric estimates. In particular only firms for which the following balance sheet variables were available in each year were kept in these restricted panels: stocks recorded at book value (var6), shareholders' funds (var11), non-current liabilities (var14), current liabilities (var17), total shareholders' funds and liabilities (var21), and turnover (var25). The data were also asked to satisfy the obvious equality: var21 = var11 + var14 + var17. After dealing with the missing value problem, the initial samples dropped by 17.2% (from 2751 to 2277) for France, by 5.1% (from 2581 to 2450) for Italy and by 39.1% (from 2869 to 1748) for the UK. The higher fall in the UK sample can be almost entirely attributed to the lack of the turnover variable for a larger number of firms.

Final samples were obtained by applying two different trimming procedures to the intermediate samples. First from each intermediate sample firms were removed with extreme observations of the stock to turnover ratio (var6/var25). Operationally, firms with at least one observation above the 0.99 or below the 0.01 quantile were excluded. The purpose of this procedure is to exclude firms with anomalously high/low levels (in proportion to total sales) of the dependent variables in the estimated equations. Second, a very similar procedure was adopted for the logarithmic first differences of all variables used in estimation: stocks (var6), turnover (var25), leverage ((var14 + var17)/var21), short-term leverage (var17/var21), and maturity (var17/(var17 + var14)). Operationally, firms with at least one observation above the 0.999 or below the 0.001 quantiles were excluded. In this case, the purpose is to exclude from the samples firms with very high growth rates in absolute value. After the two trimming procedures the samples further reduced to 2093 firms (and 18 837 observations) for France, 2254 (and 15 778 observations) for Italy and 1560 (and 14 040 observations) for the UK.

Finally, in Table A2 firms are split according to a number of criteria. In particular both the French and the Italian panels are made exclusively by unconsolidated balance sheet data whereas for the UK 47.3% of the data represent consolidated figures. This might explain, at least partially,

Table A1. *Sample building procedures*

	Initial sample	Intermediate sample	Final sample
France	2751	2277	2093
Italy	2581	2450	2254
UK	2869	1748	1560

Table A2. *Firms splits (%)*

	Consolidation	Size (large)	Age (old)	Quotation
France	0.00	51.74	80.89	1.48
Italy	0.00	43.74	65.75	0.71
UK	47.31	64.94	82.69	16.28

why the proportion of large firms – defined as firms with real sales larger than 20 million Ecus in the first sample – is higher (64.9%) for the UK compared with France (51.7%) or Italy (43.7%). On the basis of the 'year of foundation' variable, Italian firms turn out to be younger than their British and French couterparts. In fact the proportion of firms older than ten years in 1989 is 65.8% for Italy which has to be compared with much higher figures for the UK (82.7%) and France (80.9%). Finally, according to the information provided by BVD, the proportion of firms listed on the Stock Exchange is very limited for Italy (0.78%), and France (1.48%). Only the UK sample shows a sizeable proportion of listed firms (16.0%), hampering the possibility of using the listed/non-listed status as an alternative candidate to proxy for the degree of capital market access.