Quarterly Journal of Experimental Psychology,**55**, 3, 705-731, 2002

Cognitive Euroscience: Scalar variability in price estimation and the cognitive consequences of switching to the Euro

Stanislas Dehaene

INSERM, CEA, Orsay, France

J. Frederico Marques

Universidade de Lisboa

Stanislas Dehaene (corresponding author)

INSERM, Unité 334, Service Hospitalier Frederic Joliot, CEA/DRM/DSV,
4 Place du general Leclerc, 91401 Orsay Cedex, France

J. Frederico Marques

Faculty of Psychology and Education, University of Lisbon,

Alameda da Universidade, 1649-013 LISBOA, PORTUGAL

ABSTRACT

The present paper examines the ability to evaluate prices in a familiar currency (French Francs, Portuguese Escudos or Irish Punt) and in an unfamiliar currency (Euro). Participants evaluated prices for different items either by selecting the most appropriate price from a set of alternatives (in a timed or not timed version) or by directly producing a price estimate for each item. The results followed Weber's law: the standard deviation of estimated prices was proportional to their mean. The Weber fraction was stable for the familiar currency in different countries, but was significantly higher for the unfamiliar currency. We suggest that price estimation relies on a learned mapping between items and their prices as represented by distributions of activation on an internal number line. The observed Weber fraction reflects the degree of expertise with a given currency as well as a minimal variability intrinsic to the number line itself.

Acknowledgments

We thank Anthony Barrett and Pedro Martins for their help in building the Irish materials and providing us participants' contacts. J. Frederico Marques was supported by the Centro de Psicometria e Psicologia da Educação da Universidade de Lisboa. S. Dehaene was supported by INSERM and the McDonnell foundation.

INTRODUCTION

The present paper examines how our current knowledge of the psychological principles of number processing applies to a practical domain: the ability to evaluate prices in different currencies. Interactions between cognitive psychology and applied research on price perception can be fruitful in both directions. As we shall see, price estimation is a valid psychological task that casts new light on the organization of mental representations of number. Conversely, knowledge of psychological mechanisms of number processing provides a theoretical foundation from which to evaluate, and hopefully to mitigate, the practical difficulties that can be anticipated when the majority of European countries will switch to a new currency, the Euro.

The mental representation of numbers has been the focus of considerable work in cognitive psychology (for reviews see Butterworth, 1999; Dehaene, 1992; Dehaene, 1997). Adult humans appear to possess a semantic representation of numerical quantities, which helps to evaluate proximity relations between numbers. Two characteristics constitute the signature of this representation (Dehaene, Dehaene-Lambertz, & Cohen, 1998). First, the distance effect reflects the fact that, when subjects have to discriminate two numbers or to decide which of two numbers is the larger, their reaction time and error rate is inversely related to numerical distance. For instance, subjects are faster in deciding that 9 is larger than 2 than in deciding that 6 is larger than 5. Second, the number size effect reflects the fact that, for an equal distance between the numbers, reaction time and error rate increase with number size. Thus, subjects are slower in deciding that 9 is larger than 8 than in deciding that 2 is larger than 1, although the distance is one unit in both cases (Moyer & Landauer, 1967).

The theoretical interpretation of those effects is that the mental representation of numbers is fuzzy and obeys Weber's law, so that larger numbers are increasingly less discriminable. According to one mathematical formulation of this property, numerical quantity is internally represented by a linear "accumulator", and the variability in this internal representation increases linearly with the size of the numbers involved, with a fixed proportionality constant (Whalen, Gallistel, & Gelman, 1999). Another formulation stipulates that the internal representation of numbers takes the form of a compressive logarithmic scale, so that increasingly larger numbers are subjectively closer together on the mental 'number line'. This representation would be subject to a fixed variance (Dehaene, 1993; Dehaene & Mehler, 1992). Whether and how the two linear and logarithmic formulations can be distinguished remains a matter of debate (for a recent discussion, see Brannon, Wusthoff, Gallistel, & Gibbon, 2000; Dehaene, 2001) Both imply that the standard deviation of numerical estimates increases in direct proportion to the number being estimated, a property called scalar variability (the proportionality constant is then called the Weber fraction). Assuming that numerical discrimination is hindered by overlapping internal representations, both models also imply that numbers that are closer in size are increasingly more confusing, explaining the distance effect, and also that this effect of proximity worsens as the numbers get larger, explaining the number size effect.

Scalar variability was initially demonstrated in animals' perception of numerosity (Mechner, 1958; Meck & Church, 1983). Recently, Whalen et al. (1999) tested directly the hypothesis that humans also exhibit scalar variability during numerosity perception and production tasks. In the perception task, subjects viewed a fast random sequence of flashes and were asked to estimate, without counting, their approximate number. In the converse production task, they were given a number and asked to tap with their finger, as fast as possible and without counting, an approximately equivalent number of taps. In both cases, the standard deviation of the subjects' responses increased linearly with number size, indicating scalar variability compatible with Weber's law.

Whalen et al. (1999) interpreted their result as indicating that a central abstract representation of numerical quantity, common to perception and production contexts, was subject to Weber's law. However, their findings were limited to a situation in which one of the two numbers was non symbolic, and hence perceptual or motor errors could have contributed to the observed response variability. We reasoned that, if the semantic representation of numerical quantities is Weberian, then scalar variability should be observed whenever this representation is accessed, even when both the input and output stimuli are symbolic rather than perceptual or motor. Price estimation appears as an ideal task to test this hypothesis. As consumers, we all carry in our minds a reasonable price for most products. Given a product name, we can quickly provide a reasonable number, expressed in a familiar currency. If the underlying representation of numerical quantity is Weberian, the variability in the estimated price should increase in direct proportion to the price of the item to be estimated. Because this task involves well-defined symbolic inputs and outputs (the name of the product and the numerical answer), variability in the input-output function cannot be attributed to perceptual or motor factors, but must reflect the internal structure of the mental representations involved.

The exact nature of this variability might even provide a new test of the linear versus logarithmic models of the number line. Suppose that the price estimation task requires two steps: first, the generation of an internal quantity sample on the internal number line, and second, the accurate naming of that approximate quantity with number words or digits. Then the distribution of subjects' responses should accurately reflect the internal variability in the representation of a given quantity. According to the linear model, this distribution should be Gaussian and therefore symmetrical. According to the logarithmic model, however, it should be skewed and should follow the "log-normal" law because the samples should be distributed normally on a logarithmic scale.

What is the evidence that price estimation indexes the internal quantity representation? Previous marketing and consumer research suggests that price knowledge is not stored in the form of exact declarative knowledge, but involved an inherently variable and approximate component. Dickson and Sawyer (1990) questioned shoppers just after they selected a product from the supermarket shelf and observed that less than half were able to state the correct exact price. Recalled prices varied greatly and, on average, fell lower than the actual price. Some of the shoppers even had in mind a price that was not only inaccurate, but was higher or lower than all of the actual prices in the brand choice set. Other experimenters have suggested that price knowledge is not random, but obeys Weber's law (Webb, 1961). For instance, Lambert (1978) presented subjects with partially modified advertisements for products and asked subjects to detect "anything that differed from the brand you usually buy". The frequency with which subjects noticed a change in price did not depend on the absolute size of the price change, but only on its ratio to the price of the item. Likewise, (Grewal & Marmorstein, 1994) demonstrated that subjects do not perceive the value of a proposed saving in absolute terms, but only relative to the mean price of a product. This explains the economical paradox that consumers often spend more time looking for a small saving on a cheap product than for an objectively much larger saving on a more expensive products (Grewal & Marmorstein, 1994). While those studies suggest scalar variability in price representation, they have not directly verified it by measuring the Weber fraction across a large range of prices. In the present paper, we demonstrate how this can be done accurately for single subjects in a few minutes of testing.

Another interesting feature of price estimation is that the numerical representations that subjects form for prices are only valid once a currency is fixed. What happens when this reference changes? Can subjects quickly rescale their previous map of prices? Or do they have to slowly relearn, for each product, on which region of the number line its price maps? Once subjects are informed of the exchange rate between two currencies, mental calculation can be used to convert prices. However, this calculation can be quite difficult, as most exchange rates are non-integers. Furthermore, if the above construal of the price estimation task is correct, subjects normally do not use calculation when thinking about prices, but rather map directly each product onto an approximate region of their mental representation of quantities. Thus, if the introduction of a new currency implies suddenly having to calculate, this should have a strong impact on reaction times and error rates in price estimation.

Millions of Europeans are currently involved in such an experiment. In January 2002, they will all be forced to use a new currency, the Euro, in all aspects of their daily life. To facilitate the transition, since January 1999 the Euro has begun to be introduced alongside the local currency in price tags, bills, and bank forms. In the present paper, we attempt to determine the psychological costs incurred during this transition period. Naturally, many political and sociological factors can also be expected to affect the ease with which people in different countries will switch to the Euro (Marques, 1999). An extensive international study on attitudes towards the euro, collected in the summer of 1997, showed important differences between countries in attitude towards the euro, related to several factors including perceived equity (i.e. ratio of benefits to contributions to the European Union), attitude towards Europe, and national identity (Müller-Peters, 1998; Müller-Peters et al., 1998; Pepermans & Verleye, 1998). In the present paper, however, we focus exclusively on cognitive psychological factors. We develop a simple mathematical method to analyze price estimation data and to measure the Weber fraction and the response distribution in both the native currency and the Euro. We also measure the objective response time cost when having to use an unfamiliar currency.

EXPERIMENT 1

Multiple choice amongst prices in the usual currency

The goal of experiment 1 was to provide background data on the ability of normal adults to evaluate prices in their usual currency. For each of 40 items, participants were asked to select the most appropriate price amongst several choices. This experiment provided the opportunity to evaluate the scalar variability hypothesis in price evaluation: is the standard deviation of participants' price estimates proportional to the meanprice? We also evaluated whether sex has an effect on accuracy and bias in price estimation.

Method

Participants

46 participants (20 males) ranging in age from 16 to 60 (average 35.5 years) took part in this experiment on a voluntary basis. The participants belonged to the administrative or scientific communities in the Paris area. Their education ranged from 13 to 26 years (average 19.0 years). They were contacted through the broadcast of an e-mail questionnaire, which they completed and returned on a voluntary basis. All were native of France.

Stimuli

A list of 40 familiar products or services was selected. For each, seven proposed prices were selected as follows. First, the experimenter generated a plausible price for each item (subsequent data processing, described below, ensured that this partially arbitrary choice had no influence on the results). Second, a random increase or decrease of up to 10% was applied to avoid round numbers. Third, a random rank between 2 and 6 was assigned to this price in the series of seven proposed prices. Fourth, the remaining elements of the series were generated as a geometric series with a power of 1.5. This ensured that each of the proposed prices was 50 % higher than the preceding one, and that the seven proposed prices spanned slightly over one order of magnitude. Finally, the prices were rounded as follows. If the highest price of the series was greater than 100 FF, then all prices were rounded to the nearest integer; otherwise they were rounded to one significant digit after the decimal point, and printed with two digits after the decimal point (to appear in the familiar presentation with francs and centimes). Here are some sample stimuli:

A pack of cigarettes:

4,90 FF   7,40 FF   11,10 FF 16,60 FF 24,90 FF 37,40 FF 56,00 FF

A bottle of Champagne:

56 FF   84 FF   127 FF   190 FF   285 FF   427 FF   640 FF

Procedure

Participants were told that they would be presented with various familiar products or services, followed by a list of prices in ascending order, expressed in French francs. They were asked to select the price that fell closer to the normal or typical price of each item. They were asked to answer quickly, intuitively, and without calculating. They were asked to select only one price by marking it with an «x ».

Results

Application of an automated message parsing program resulted in the extraction, for each item i and each participant s, of a price estimate $P_{i,s}$. The item-based analysis was based on the mean m_i and standard deviation sd_i of the estimates across participants, as well as on the Weber fraction defined as $w_i = \frac{sd_i}{m_i}$. As shown in Figure 1a, there was a linear relation between standard deviation sd_i and mean price m_i. Simple regression showed that 95.1% of the variability in $\underline{sd_i}$ across the forty items could be accounted for by the following equation: $\underline{sd_i}$ = .337 $\underline{m_i}$ ($\underline{t}$=44.1, 39 d.f., $\underline{p}<10^{-32}$). The constant term, when introduced in this regression, was found non-significant, suggesting that a simple proportionality relation between $\underline{sd_i}$ and $\underline{m_i}$ accounted for the data.

The above analysis is not fully appropriate because linear regression assumes linearly additive noise, which is clearly contradicted by the data. Furthermore, a significant regression indicates that $\underline{sd_i}$ increases $\underline{m_i}$, but does not test whether the shape of the relation is strictly linear. To further test this, we used a log-log regression. In this representation (Figure 1b), scalar variability predicts that variability estimates should follow a straight line with a slope of 1 (log $\underline{sd_i}$ = log $\underline{m_i}$ + log $\underline{w}$, where $\underline{w}$ is the Weber fraction). A slope different from 1 would indicate that the relation between $\underline{sd_i}$ and $\underline{m_i}$ is not linear, but a more complex power function (e.g. a slope of 2 would indicate a quadratic relation). Strict linearity was observed, however. The log-log regression was highly significant (t=33.2, 39 d.f., $p<10^{-28}$), yielding a very precise estimate of the slope, which did not differ significantly from 1 (mean ± 2 standard errors = 1.03 ± .06). The intercept of the regression gave a Weber fraction value of .295, close to the above estimate of .337. However, this estimate is sensitive to the estimation of the slope of the regression, hence has a wide 95% confidence interval (.229-.381). The Weber characteristic can be measured more simply by plotting the Weber ratio $\underline{w_i} = \frac{sd_i}{m_i}$ for each item against its mean price (Figure 1c). Although $\underline{w_i}$ varied somewhat across items, no systematic linear or non-linear relation with price was found, suggesting that the Weberian relation provided a good description of price estimation. Across items, the mean Weber fraction was .346± .038.

(Figure 1 about here)

Figure 2 shows graphically the distribution of participants' responses when working with their familiar currency and with the Euro. To visualize the distribution of responses with very different orders of magnitude, each individual price estimate $\underline{P_{i,s}}$ was first divided by the mean estimated price of the corresponding item, $\underline{m_i}$, to yield an index of deviation from the central price estimate. The distribution of the resulting indices (Figure 2) was highly asymmetrical, with a long tail (skewness = 1.74). A better fit of this distribution was achieved with a log-normal distribution (χ^2 (18 d.f.) = 146) than with a simple normal distribution (χ^2 (18 d.f.) = 339), though the deviations from the theoretical distribution were significant in both cases ($p<10^{-4}$). The skewness of the distribution indicates that subject make errors that are numerically larger when over-estimating than when under-estimating. This is rather intuitive, as it merely means that, if a subject underestimates the cost of a 100 FF item as costing only 50 FF, the "equivalent error" in the direction of over-estimation is 200 FF (same ratio to 100 FF) rather than 150 FF (same difference to 100 FF). It does, however, imply that subjects sample at random from a logarithmically organized internal number line rather than from a linear continuum. This was further confirmed by plotting the distribution of the logarithms of the subjects' responses (log $\underline{P_{i,s}}$), after normalization by subtracting the logarithm of the mean response for that item (log $\underline{m_i}$). This distribution was essentially symmetrical (skewness = -.23) and a relatively good fit was obtained with a simple normal distribution χ^2 (16 d.f.) = 129).

Note that, if the shape of the price distributions is log-normal, estimating the Weber relation through the standard deviation of responses is not optimal. Ideally, the Weber relation should be evaluated by calculating the standard deviation of the *log* responses (sdl_i = sd(log $\underline{P_{i,s}}$)). As predicted by Weber's law, we verified that this value remained constant (mean = .33) over the range of items, as shown by a non-significant linear regression with the mean of the log response (ml_i = mean(log m_i)) (t=.48, n.s.). However, because this demonstration, although theoretically more appropriate, is based on a null effect, in subsequent analyses we continue to report the more standard analysis based on the positive finding of a linear relation between standard deviation and mean price.

(Figure 2 about here)

Finally, because of earlier reports of sex differences in mathematical cognition (Benbow, 1988; Hyde, Fennema, & Lamon, 1990), the results were examined for differences between participants. The participant-based analysis was based on the computation of the location $l_{i,s}$ of each participant's estimate relative to the overall group: $l_{i,s} = \frac{P_{i,s} - m_i}{sd_i}$. Note that $l_{i,s}$ is expressed in standard deviation units and is therefore comparable even for items that differ widely in mean price. Each participant's bias and dispersion measures b_s and d_s were calculated, respectively, as the mean and standard deviation of the $l_{i,s}$ across items for a given participant. Bias b_s measures whether a given participant tends to systematically under- or overestimate prices relative to the group mean. Bias is negative for underestimation, and positive for overestimation. Dispersion d_s evaluates the extent to which the responses of this participant varies across items. Dispersion is low if a participant's responses are tightly grouped (even if they systematically fall below or above the mean); it is high if the participant's responses deviate widely sometimes above and sometimes below the group mean. Though no bias effect was found, dispersion was found to be systematically more elevated by 14.1% in females than in males (t=2.35, 44 d.f., p<.023).

Discussion

The results validate the hypothesis that price estimation follows Weber's law. Participants approximate prices with a variability that is proportional to the estimated price. The measured Weber fraction of .34 is somewhat more elevated than the value of .15 observed by Whalen et al. (1999).This may indicate that participants have an inherently variable and approximate internal representation of prices, with a high variance. Furthermore, this variance tended to be somewhat higher in females than in males. Finally, distribution analyses suggest that this variability is best represented as originating from Gaussian-shaped fluctuations on a logarithmic internal scale.

Before accepting these conclusions, however, three alternative interpretations must be considered. First, the Weberian relation could be an artifact of the design of the multiple-choice task. For each item, the proposed prices formed a geometric series. As a result, as the items increased in price, the range of proposed prices also increased in direct proportion. One may wonder whether this alone could have imposed a Weberian relation in the results. Second, it is possible that the Weber fraction reflects a judgment based only on familiarity and frequency of purchase. The higher the price, the less frequently the product is bought and so participants' choices are more uncertain and variable. Finally, a third hypothesis is that the Weber fraction is not a property of the participants' mental representations of numbers, but a genuine relation in our environment. Actual prices do vary in the external world in direct proportion to the mean price (Grewal & Marmorstein, 1994). Perhaps the participants' variable estimates merely represent an accurate depiction of the range of prices for a given item in their environment. Experiment 2 evaluated these possibilities.

EXPERIMENT 2

Production of prices in the usual currency

Experiment 2 was identical to experiment 1 except that participants were asked to directly produce a price estimate for each item, rather than select amongst proposed prices. This permitted us to evaluate whether the Weberian relation was induced by the multiple-choice procedure. Furthermore, participants also evaluated the familiarity and intrinsic variability of each item's price. This permitted to evaluate to what extent the variability in price estimates was related to an actual variability of prices in the environment.

Method

The same list of items as in experiment 1 was presented by electronic mail to 10 native French participants drawn from the same population as above (6 males; mean age 31.3 years; mean education 20.8 years). For each item, participants were asked, first, to estimate the frequency with which they buy it, on a scale from 1 (very rarely) to 7 (very frequently); second, its normal or typical price; and third, the variability of this price in different shops or stores, on a scale from 1 (very constant) to 7 (highly variable).

Results

The relation between standard deviation and mean price was again found to be linear (figure 3), with 94.9% of variance being accounted for by the equation $sd_i = .433\ m_i$ (t=41.5, 39 d.f., $p<10^{-30}$). The log-log regression (see Figure 3) was also significant (t=20.7, 39 d.f., $p<10^{-21}$) and its slope did not differ from 1 (1.06 ± .10). Across items, the mean Weber fraction was .304, a value that did not differ significantly from the one observed in experiment 1 (t=1.72, 39 d.f., p=.09).

The results of experiments 1 and 2 were further compared by regression analyses. Prices obtained by both methods were extremely well correlated (r^2 = 96.6%). In this regression, no constant term was needed, indicating direct proportionality of the prices in the two sets. However, the slope of 1.13 differed significantly from 1 (t=9.19, 39 d.f., $p<10^{-10}$), indicating that the prices in experiment 1 tended to be more elevated than in experiment 2, perhaps due to a different population of participants or to the different response mode.

(Figure 3 about here)

Correlation analyses indicated that, as expected, there were modest but significant correlation amongst product prices, buying frequency, and price variability. The higher the price, the lower the estimated frequency of buying (r=-.377, p=.017), and the higher the estimated price variability (r=+.502, p=.001). Frequency and variability were also negatively correlated (r=-.528, p=.0005). In spite of these correlations, however, neither frequency nor variability alone sufficed to explain the linear increase of the standard deviation with the mean price. In a partial regression analysis, even after removing the effects of frequency and variability on sd_i, the residuals remained linearly correlated with mean price, according to the equation $\text{residual } sd_i \propto .327\ m_i$ (t=9.61, 38 d.f., $p<10^{-11}$). The same was true for a similar retrospective analysis of experiment 1, using the frequency and variability estimates from experiment 2 ($\text{residual } sd_i \propto .279\ m_i$ (t=9.65, 38 d.f., $p<10^{-11}$).

The Weber fractions in both experiments were correlated across items (r^2 = 18.8% ; t=2.96, 38 d.f., p=.005), indicating that the variability in the Weber fractions for different items was partially systematic and replicable in both experiments. To explore the hypothesis that Weber fractions are not a property of the human numerical cognition system, but rather reflect objective properties of the products, and particular the actual variability of their prices, correlations were computed between the Weber fractions observed in both experiments and the estimated buying frequency and price variability. Weber fractions increased significantly with estimated price variability (two-tailed tests; experiment 1 : r=.404 ; p=.010 ; experiment 2 : r=.396, p=.011), and also tended to decrease with buying frequency (two-tailed tests; experiment 1 : r=-.302, p=.058 ; experiment 2 : r=-.307, p=.054). Importantly, however, even after correction for those factors, there was no sign of a tendency for the residual of the Weber fraction to vary with the mean price, either on a linear or on a log-log plot, in either experiment 1 or 2 (r^2 <5.2%; t<1.44, 38 d.f., n.s.). The mean residual Weber fraction, as interpolated for the most familiar and highest frequency items, was a non-zero value of .130 ± .046 in experiment 1, and .225 ± .034 in experiment 2.

Because of the small number of participants in experiment 2, no distribution and between-participants analyses could be performed.

Discussion

The results of experiment 2, using self-generated prices, largely confirmed those of experiment 1 using a multiple-choice procedure. The average prices and Weber fractions obtained by both procedures were highly correlated. This indicated that our first hypothetical artifactual explanation of Weber's law for prices can be rejected: Weber's law in experiment 1 was not simply due to the presentation of a geometrical range of proposed results, since it was also obtained in experiment 2 where participants could spontaneously offer any numerical response.

A second possible artifact proposes that the higher the price, the less frequently the product is bought and hence the greater the uncertainty about its prices. This explanation too could be rejected. Participants' estimates of buying frequency and price variability did vary modestly, but significantly with price in the predicted direction (higher price, lower frequency, higher variability). However, the almost perfectly linear relation between standard deviation and mean price could not be reduced to a frequency of buying effect, since it remained significant even after the effects on the standard deviation of frequency and estimated price variability were removed. In other words, our list contained sufficiently many products that were high-priced but familiar (e.g. a pocket-size paperback book), or conversely were cheap but infrequently bought (e.g. a deck of cards), so that we could break the correlation between the two variables and show that Weber's law was due to price itself.

A third possible artifact is that scalar variability is not a deep property of the mental representation of prices, but is merely an internalization of an objective property of the external world, namely the fact that the higher the price, the larger its range of variability. In support of this explanation, experiment 2 revealed that the Weber fraction for different items varies with estimated buying frequency and variability. First, there is a frequency effect on the accuracy of price representations, as indicated by a significant negative correlation across items, between the Weber fraction and the estimated buying frequency. Not surprisingly, participants have a more variable representation of the prices of items that they buy less frequently. Second, and more crucially, the Weber fraction also correlated positively with estimated price variability.

This indicates that participants partially agree that some items have more variable prices than others in their environment, and that this « objective » variability is reflected in the « subjective » variability of their estimated prices.

Does this mean, however, that all of the variability in estimated prices reflects the variability in actual prices? Is the Weberian property entirely due to participants having internalized an objective property of pricing in their environment? The results suggest otherwise, because even items that were rated as highly invariable in price were estimated with some minimal Weber fraction. For instance, the price of postage is fixed in France, and participants did rate it very close to the minimum of the rating scale for actual variability (average 1.4, where 1 means « very constant »). Yet it was estimated with a Weber fraction of .135. Findings were similar for other largely invariable items such as a baguette (average variability rating 1.7, Weber fraction .145), a daily newspaper (respectively 1.7 and .116), and a pack of cigarettes (2 and .174). It is noteworthy that these Weber fractions are clearly non-zero and much closer to the value of about .15 observed by Whalen et al. (1999) for numerosity perception and production. This suggests that the Weber fraction for prices reflects both a representation of the variability in actual prices and a minimum residual variability due to properties of the internal representation of magnitudes on the mental number line.

EXPERIMENT 3

Multiple choice amongst prices in a novel currency

Experiments 1 and 2 validated the multiple-choice methodology and the Weber fraction measure of inter-item and inter-participant variability in price estimation. Experiment 3 asks how these observations are affected when participants are asked to estimate prices in an unfamiliar currency, the Euro (€). As noted in the introduction, the main motivation for this experiment was to apply the above mathematical tools to quantify the amount of imprecision associated with responding in an unfamiliar currency. This also provided an opportunity to examine whether scalar variability would still hold when evaluating quantities with an unfamiliar measure. The experiment was run in April 1999. At this time, the exchange rate had been announced four months before, and prices were labeled already in both currencies in many shops. However, the currency was only available for payments by check or bank-to-bank. No coins or banknotes were in circulation yet. Hence, participants had little concrete experience with the Euro. Two predictions were made. First, because of their lack of familiarity with the Euro, participants should make gross errors in price estimation, and should thus show a greater Weber fraction than when estimating prices in their familiar currency. Second, assuming that participants apply the numerical intuition they developed with their familiar currency to the new prices, they should tend to overestimate prices in Euros. This is because the actual prices, when converted in Euros, are divided by about 6 (exactly 6.55957) and hence seem too small.

Those predictions depend critically on participants' ability to follow the instructions and to respond intuitively, without calculating. If the participant applied the strategy of generating the equivalent price in French Francs, followed by mental calculation to compute the corresponding price in Euro, then one would expect little difference between performances in the two currencies. To evaluate participants' ability to comply with the instructions, a short assessment of intuition- versus calculation-based responding was added at the end of the questionnaire.

Method

16 native French participants from the same population as above (12 males; mean age 32.8 years ; mean education 19.9 years) participated via electronic mail. The method was identical to experiment 1 except for the following changes. First, participants were asked to select amongst prices proposed in Euro, which were generated by applying the exchange rate of 6.55957. Second, the following two questions were asked after completion of the test : What percentage of trials do you think that you answered based only on your intuition, without calculating ? Can you estimate you degree of intuition for prices in Euros, on a scale from 1 to 7 (1=no intuition ; 7=excellent intuition) ?

Results

When participants estimated prices in Euros, the relation between standard deviation and mean price was again linear (figure 3), with 95.0% of variance being accounted for by the equation sd_i = .331 m_i (t=39.9, 39 d.f., $p<10^{-30}$). As in experiments 1 and 2, this remained true even after the effects of buying frequency and price variability on sd_i were partialled out (residual sd_i α .236 m_i (t=10.1, 38 d.f., $p<10^{-11}$). The log-log regression of sd_i on m_i was also significant (t=35.0, 39 d.f., $p<10^{-29}$), with a slope not significantly different from 1 (1.01$\pm$ .06) indicating a strictly linear Weber relation. The distribution of normalized responses was again better fitted by a log-normal distribution (χ^2 (17 d.f.) = 81.5) than by a normal distribution (χ^2 (17 d.f.) = 93.1).

Across items, the mean Weber fraction was .362. This value did not differ significantly from the one observed in experiment 1 (.346 ; t=.94, 39 d.f., two-tailed p=.35), but it did significantly exceed the value of .305 observed in experiment 2 (t=2.68, 39 d.f., two-tailed p=.011). As shown in Figure 4, the distribution of the log responses was clearly broader in Euro than in Francs. Fits by a normal distribution revealed a 40% higher variability in Euros than in Francs (standard deviations = .174 versus .124).

(Figure 4 about here)

The Weber fractions when estimating in Euros were highly correlated with the ones obtained when estimating the prices of the same items in French Francs (exp 1, r^2 = 34.7%, p<.00004 ; exp. 2, r^2 = 29.2%, p<.0002). As in previous experiments, Weber fractions increased significantly with estimated price variability (r=.323 ; p=.041) and decreased with buying frequency (r=-.345, p=.029). Regression analyses across items also indicated that the prices estimated in Euro correlated very well with those estimated in French Francs in experiment 1 (r^2 = 97.3%) and experiment 2 (96.5%). No constant term was needed, indicating direct proportionality of the prices estimated on those different occasions. However, the regression slopes in both cases were significantly smaller than the expected value of 6.55957, which is the exchange rate that prevails between Euros and Francs. For experiment 3 against experiment 1, the slope $\pm$ one standard error was 6.363 $\pm$ .039, and for experiment 3 against experiment 2, it was 5.600 $\pm$ .077 (both $p<10^{-6}$). This suggests that participants tend to systematically overestimate prices in Euros. Indeed, when converted to the same currency, the prices in experiment 3 exceeded those estimated in experiments 1 and 2 by an average of 7%, a value that was significant across items (exp 1: z=3.87, 39 d.f., one-tailed p<.0002 ; exp 2 : z=1.73, one-tailed p=.046).

The questionnaire revealed that participants experienced difficulties in answering intuitively without calculating, as required by the instructions. The estimated percentage of trials in which the participants did not calculate was only 40% on average, and a majority of participants (9/16) thought that they had to calculate on more than half of trials. Participants also rated themselves as having a relatively poor intuition of prices in Euro (average rating = 3.06 on a scale from 1 to 7, range 1-5).

Finally, the effect of sex on dispersion, which was observed in experiment 1, was also replicated. Dispersion was more elevated by 24.6% in females relative to males (t=2.10, 14 d.f., one-tailed p<.028). No effect was observed on bias.

Discussion

When participants selected prices in an unfamiliar currency, they still followed the basic scalar variability rule. However, two differences between judgements with familiar and unfamiliar currencies were observed. First, the Weber fraction tended to be higher, suggesting that estimation was less precise in the unfamiliar currency (though this reached significance only relative to experiment 2). Second, there was a systematic overestimation of prices in the unfamiliar currency. This effect was in the direction we predicted. Given the exchange rate between Euro and Franc (1 € = 6.55957 FF), accurate prices in Euro appear as smaller numbers than the corresponding in French Francs (for instance, a stamp is 3 FF, but only .45 €). Participants may have compensated, voluntarily or not, for those counter-intuitively small numbers by proposing exaggerated prices when performed the task in Euros than when performing it in Francs.

One weakness of experiment 3 is that although we were interested in studying intuitive price judgements in the absence of explicit calculations, participants avowedly found it quite difficult to comply with this instruction and to answer without calculating. Lemaire and collaborators (Lemaire & Lecacheur, 2001; Lemaire, Lecacheur, & Ferréol-Barbey, 2001) have extensively studied the various explicit strategies that French subjects use to convert prices from Francs to Euros and vice versa. Use of such calculation strategies in our task may explain why only relatively subtle differences were found in participants' performance with familiar and unfamiliar currencies (exp. 1 versus 3). Given that participants had had little practice using the Euro in real-life situations at the time of testing, we expected to find a much larger difference in Weber fractions than was actually observed. By using a calculation strategy, participants might have achieved a greater accuracy than if they had been using estimation alone. Naturally, if this interpretation is correct, it must be at the expense of response time, which must be considerably slower when participants have to calculate than when they do not. To evaluate this hypothesis, experiment 4 replicated experiments 1 and 3 while measuring response time to each item.

EXPERIMENT 4

Timed multiple choice tests in native and novel currencies

Method

16 native French students (9 males; mean age 24.8 years; mean education 18.6 years) participated in an individual 15-minutes computerized test. All participants participated in two blocks, one in which prices were presented in Francs, and the other in which they were

presented in Euros, in random order. The same items and price choices as in experiments 1 (Francs) and 3 (Euros) were used. In each block, the Expe6 software (Pallier, Dupoux, & Jeannin, 1997) was used to present the items in random order, on one line centered on screen, followed one second later by the seven price choices ordered from left to right. Participants selected the price that seemed most appropriate by using the arrow keys to move a rectangular frame which initially surrounded the middle (fourth) proposed price, then pressing the space bar to validate their choice. Reaction times were measured from the onset of the price choices to the bar press. The same two questions as before (percentage of trials answered without calculating, and estimation of intuition for prices) were asked after completion of both Franc and Euro blocks.

Results

Response times were entered into an analysis of variance with participants as the random factor, order (Euro first or Franc first) as a between-participants factor, and currency as a within-participant factor. The results revealed only a main effect of currency ($F(1,14)=6.34$, $p<.025$): participants were 387 ms slower when judging prices in Euro (mean RT= 3131 ms) than in their familiar currency (mean RT=2745 ms). This effect was confirmed by a similar ANOVA with items as the random factor: the vast majority of items (33/40) were judged significantly more slowly when their prices were presented in Euro than when they were presented in the familiar currency ($F(1,39)=27.1$, $p<.0001$).

The relation between standard deviation and mean price was again linear (figure 3). In Francs, 98.5% of variance was accounted for by the equation $sd_i = .457\ m_i$ ($t=53.5$, 39 d.f., $p<10^{-37}$), while in Euros 93.8% of variance was accounted for by the equation $sd_i = .427\ m_i$ ($t=32.1$, 39 d.f., $p<10^{-28}$). The log-log regression of sd_i on m_i was also significant (Francs: t=35.4, 39 d.f., $p<10^{-29}$; Euros: t=44.6, 39 d.f., $p<10^{-33}$), with a slope very close to 1 ($1.06 \pm .06$ in Francs, $.96 \pm .04$ in Euros), indicating a linear Weber relation in both currencies. The distribution of normalized responses was again better fitted by a log-normal distribution (Francs: χ^2 (14 d.f.) = 47.1; Euros: χ^2 (20 d.f.) = 45.5) than by a normal distribution (Francs: χ^2 (15 d.f.) = 126.0; Euros: χ^2 (19 d.f.) = 146.2).

Across items, the mean Weber fraction was .333 in Francs and .479 in Euros. The difference was highly significant ($t=6.00$, 39 d.f., $p<3.10^{-7}$), indicating that price estimation was much more accurate in the familiar currency than in Euro. A similar result was obtained when comparing the distributions of the log responses (figure 4): Gaussian fits indicated a much larger standard deviation of the log responses in Euros than in Francs (respectively .251 and .117). Nevertheless, the Weber fractions when estimating in Euros were slightly correlated with the ones obtained when estimating the prices of the same items in Francs (r^2 = 11.2%, $p<.035$). In agreement with previous experiments, Weber fractions in Francs decreased significantly with increasing buying frequency ($r=-.32$, $p=.046$) and increased significantly with increasing familiarity with the product ($r=+.36$, $p=.023$). The corresponding correlations were not significant for prices in Euros, however (respectively $r=.12$, $p=.45$; and $r=-.18$, $p=.27$), presumably because participants had a much less reliable estimation of prices in this unfamiliar currency.

Regression analyses across items indicated that the average prices estimated in Euro correlated very well with those estimated in Francs (r^2 = 99.0%). No constant term was needed, indicating direct proportionality of the estimated prices. However, the regression slope

(6.082 ± standard error .101) was significantly smaller than the actual exchange rate of 6.55957 ($p<10^{-5}$). This replicated experiment 3, indicating that French participants overestimate prices in Euros. Indeed, when converted to the same currency, the prices in Euro were above the prices in Francs by an average of 5.8%, a value that fell short of significance (t=1.63, 39 d.f., p=.056).

Participants judged that they had complied with the instructions and responded without calculating on 95.3% of trials in their familiar currency, but only on 60.0% of trials in Euro, a highly significant difference (t=3.97, 15 d.f., p=.0006). They also rated their intuition of prices on a 1-7 scale as 5.25 in their familiar currency, but only 3.25 in Euro, again a significant difference (t=6.08, 15 d.f., $p=10^{-5}$).

Finally, the effect of sex on dispersion was replicated, female participants again showing dispersion scores 18.9% higher than male participants (t=3.67, 14 d.f., p=.0025). In this experiment, sex also influenced bias: male participants select lower prices than female participants (t=-4.09, 14 d.f., p=.001). No significant effect was found in the Euro block, presumably because of greater response variability.

Discussion

As predicted, participants were much slower in judging prices in Euro than in their familiar currency. Importantly, however, knowing that their responses were timed improved participants' compliance with the instruction to avoid any calculation: the estimated percentage of trials solved without calculating jumped from 40% in experiment 3 to 60% in experiment 4. This was reflected in a much larger discrepancy in the Weber fractions. Whereas in experiment 3 only a small difference was found, in experiment 4 the Weber fraction in Euro was almost 50% higher than its value in Francs. This was not due to a speed-accuracy trade-off, since participants spent greater time on the Euro trials. Rather, participants spontaneously rated their intuition of prices in Euro much lower than in Francs. Presumably, this loss of intuition resulted in a greater imprecision, as measured by the Weber fraction and therefore a greater reliance on calculation.

Two other important findings were replicated in experiment 4. First, participants showed a systematic bias of overestimation in Euro relative to Francs. Second, males again showed lower dispersion scores than females. Both findings are further studied in experiments 5 and 6.

EXPERIMENT 5

Generalization to another country: Portugal

In experiment 5, we replicated the above experiments 1-4 with subjects from another country, Portugal. The first goal was to study which of the above effects could be replicated. The second, more specific goal, was to examine the effect of the exchange rate on price over or underestimation. In Portugal, the familiar currency, the Escudo, is much lower than the Euro (1 Euro = 200.482 Escudos). Thus, one may expect that subjects will have an even greater tendency to overestimate prices in Euro, because the proposed numbers in Euro will look essentially 200 times too small to them. However, because the exchange rate is very close to a round number, it is also conceivable that subjects will think of alternative strategies. For instance, the price in Euro can be obtained by multiplying the price in Escudos by 5 and correcting by a factor of 1000. If subjects use this strategy, they might be expected to underestimate, rather than overestimate, the Euro prices.

Method

Items were translated to Portuguese, and seven proposed prices were generated in both Escudos and Euros using the procedure described in the method section of experiment 1. In the e-mail based multiple choice experiment, participants were contacted through the second author contacts and belonged to the administrative or scientific communities in the Lisbon area. 24 native Portuguese participants participated in the familiar currency version , similar to experiment 1 (13 males; mean age 28.7 years; mean education 16.7 years); and 23 other in Euro version, similar to experiment 3 (13 males, age 31.2 years, education 16.4 years). All participants answered on a voluntary basis and were sent a small report providing the general results and a brief account of the research in which they participated.

In the timed multiple choice experiment, similar to experiment 4, 23 university students participated (6 males; age 19.4 years; education 14.1). Finally, 19 other university students estimated price variability and frequency of buying as defined in experiment 2 (4 males; age 18.2 years; education 13.1). The two university students groups participated for partial fulfillment of an introductory psychology course requirement and were tested individually at the Faculty of Psychology and Education, University of Lisbon.

Results

For participants in the timed test, response times were entered into an analysis of variance with participants as the random factor, order (Euro first or Franc first) as a between-participants factor, and currency as a within-participant factor. The results revealed a main effect of currency ($F(1,25)=26.0$, $p<.0001$) : participants were considerably slower when judging prices in Euro (mean RT= 5152 ms) than in their familiar currency (mean RT=3465 ms). There was no main effect of order ($p>.29$), though order interacted with currency ($F(1,25)=5.89$, $p<.023$), revealing a small priming effect: the slowing down in Euro was smaller, though still significant ($p=.01$, effect size 854 ms), when participants had first judged the prices of the same items in their familiar currency than when they started with the judgement in Euro (effect size 2461 ms). The effect of currency was confirmed by a similar ANOVA with items as the random factor. In fact, all items were judged significantly more slowly when their prices were presented in Euro than when they were presented in the familiar currency ($F(1,39)=181.7$, $p<.0001$).

The relation between standard deviation and mean price was again linear in all groups (Figure 3). For the e-mail test, in Escudos 87.8% of variance was accounted for by the equation $sd_i = .315\ m_i$ ($t=18.7$, 39 d.f., $p<10^{-20}$), while in Euros 97.3% of variance was accounted for by the equation $sd_i = .431\ m_i$ ($t=40.8$, 39 d.f., $p<10^{-32}$).For the timed test, in Escudos, the equation was $sd_i = .413\ m_i$ (variance 96.5%; $t=67.9$, 39 d.f., $p<10^{-41}$), while in Euros it was $sd_i = .547\ m_i$ (variance 93.8%; $t=35.5$, 39 d.f., $p<10^{-30}$). The log-log regressions of sd_i on m_i were all significant (all $ps<10^{-29}$), and all slopes fell very close to 1 ($1.01 \pm .06$ in Escudos, email test; $1.03 \pm .04$ in Euros, email test; $1.06 \pm .06$ in Escudos, timed test; $1.02 \pm .04$ in Euros, timed test), indicating a linear Weber relation in all cases. The distribution of normalized responses was always better fitted by a log-normal distribution (Escudos, email: χ^2 (18 d.f.) =195.6; Euros, email: χ^2 (21 d.f.) =160.1; Escudos, timed: χ^2 (18 d.f.) =79.1; Euros, timed: χ^2 (23 d.f.) = 65.2) than by a normal distribution (Escudos, email: χ^2 (18 d.f.) =307.8;

Euros, email: χ^2 (19 d.f.) =177.7; Escudos, timed: χ^2 (18 d.f.) =183.0; Euros, timed: χ^2 (21 d.f.) = 240.4).

The mean Weber fraction was higher in Escudos than in Euros, indicating that price estimation was much more accurate in the familiar currency (e-mail test, respectively .380 versus .448, t=3.36, 39 d.f., p=.0009; Timed test, .375 versus .531, t=7.35, 39 d.f., $p<10^{-8}$). Indeed, Gaussian fits to the distribution of the log responses indicated a much larger standard deviation in Euros than in Francs (email version: .245 versus .172; timed version: versus .160). Nevertheless, the Weber fractions when estimating in Euros were slightly correlated with the ones obtained when estimating the prices of the same items in Escudos, suggesting that there were intrinsic determinants of price variability in the items themselves (e-mail test, r^2 = 19.1%, p=.005; timed test, r^2 = 13.3%, p=.012). In agreement with previous experiments, for the timed test in Escudos, Weber fractions decreased significantly with increasing buying frequency (r=-.55, p=.0002) and increased significantly with the rated price variability of the product (r=.44, p=.0048). Although these correlations did not reach significance for the e-mail test in Escudos, nor for both tests in Euros, they were always in the appropriate direction (negative with frequency, positive with variability). Even for the items judged as less variable, however, the Weber fraction remained non negligible. For instance, for a stamp, it was .204 in the e-mail test and .333 in the timed test; for a pack of cigarettes, respectively .172 and .252; and for a cup of coffee, .213 and .205.

Regression analyses across items indicated that the average prices estimated in Euro correlated very well with those estimated in Escudos (e-mail test, r^2 = 99.5%; timed test, r^2 = 96.4%). No constant term was needed, indicating direct proportionality of the estimated prices. Results concerning estimation biases were somewhat different in the timed and e-mail versions of the experiment. In the timed test, the regression slope (256.5 $\pm$ standard error 3.9) was significantly larger than the expected value of 200.482, which is the exchange rate that prevails between Euros and Escudos. Thus, participants tended to systematically underestimate prices in Euros. Indeed, when converted to the same currency, the Euro prices were below the Escudo prices by an average of 8.4%, a value that was significant across items (t=2.88, 39 d.f., p=.006). In the e-mail test, however, the regression slope of 206.7 $\pm$ 2.4 was only slightly, though still significantly larger than 200 (p=.014), and when converted the Euro prices exceeded the Escudo prices by a non-significant .8%.

Because the vast majority of participants were female, sex differences could not be evaluated. The questionnaire showed that participants judged that they had complied with the instructions and responded without calculating more often when working in their familiar currency than when working in Euro (e-mail test: 74.5% versus 57.3% of trials, t=2.02, 45 d.f. one-tailed p=.025; timed test, 66.8% versus 57.5%, t=1.89, 26 d.f., one–tailed p=.035). They also rated their intuition of prices as higher in their familiar currency than in Euro (e-mail test: 4.80 vs. 3.60, t=3.22, 45 d.f., one-tailed p=.0012; timed test: 4.80 vs. 3.00, t=7.05, 26 d.f., one-tailed $p<10^{-7}$).

Discussion

Portuguese participants behaved very similar to the previous groups of French participants. First, they estimated prices with a variability directly proportional to the estimated price. Second, their Weber fraction was again around .38 when the familiar currency was used. Third, the Weber fraction was considerably larger, by a factor of almost 50%, when participants had to estimate prices in the novel currency. Fourth, it also took them much longer to select a price in the novel currency than in the familiar currency.

While French subjects overestimated prices in Euro, Portuguese subjects underestimated them. This suggests that biases in the perception of the Euro are more complicated than we initially thought. In both countries, the exchange rate is such that prices in Euro appear as much smaller numbers than prices in the original currency (in other words, the Euro is a higher-valued currency). Clearly, that factor alone does not suffice to account for subjects' perceptual biases. It seems likely that the exact strategy used by the subjects must be considered here. One possibility is that Portuguese subjects compare the Euro with the value of 1000 Escudos, a strategy encouraged by the small value of the Escudo, the nearly round exchange rate of 1 € = 200.482 Escudos, and the possible use of a strategy of multiplying the number in Escudos by 5. If so, then the number representing the price in Euro would actually feel too large. We postpone a full discussion of this issue, however, until after the last experiment where we examine yet a third country.

EXPERIMENT 6

Generalization to another country: Ireland

In experiment 6, we generalized the above findings to a third country, Ireland. Beyond the mere replication of the results, the choice of Ireland was made because of all European countries switching to the Euro, it is the only one in which the present currency, the Irish Punt (IEP), is of a higher value than the Euro (1 € = .787564 IEP). The exchange rate is a non-integer slightly below one. This provides a unique opportunity to test the idea that biases in the estimation of Euro prices are related to the exchange rate. If this idea is correct, Irish subjects should judge that prices in Euro feel too large, and hence they should underestimate them.

Method

Items were translated in English, and seven proposed prices were generated in both Irish punts and Euros using the procedure described in the method section of experiment 1. 20 native Irish participants participated in an e-mail based multiple choice experiment in their familiar currency, similar to experiment 1 (10 males; mean age 39.5 years; mean education 17.1 years). 20 other participants participated in a similar experiment in Euro, similar to experiment 3 (9 males, age 31.2 years, education 18.4 years). Participants were contacted through two residents in Ireland and through several Irish Universities e-mail databases for faculty. All participants answered on a voluntary basis and were sent a smallreport providing the general results and a brief account of the research in which they participated.

Results

The relation between standard deviation and mean price was again linear (Figure 3). In Irish punts, 83.2% of variance was accounted for by the equation $\underline{sd}_i = .335\ \underline{m}_i$ ($\underline{t}$=15.7, 39 d.f., $\underline{p}<10^{-17}$), while in Euros 85.8% of variance was accounted for by the equation $\underline{sd}_i = .310\ \underline{m}_i$ ($\underline{t}$=18.2, 39 d.f., $\underline{p}<10^{-19}$). In both cases, the log-log regression of $\underline{sd}_i$ on $\underline{m}_i$ was significant (both $\underline{ps}<10^{-26}$), and the slopes fell very close to 1 (1.03 ± .06 in Punt; 1.04 ± .08 in Euros), indicating a linear Weber relation. The distribution of normalized responses was again better fitted by a log-normal distribution (Punt: χ^2 (16 d.f.) =96.0; Euros: χ^2 (17 d.f.) =126.6) than by a normal distribution (Punt: χ^2 (17 d.f.) =178.9; Euros: χ^2 (19 d.f.) =333.4).

The mean Weber fraction was higher in Punts than in Euros, indicating that price estimation was more accurate in the familiar currency (respectively .349 versus .444, t=3.03, 39 d.f., p=.004). This was confirmed by examining the distribution of the log responses (fig. 4), which was Gaussian with a greater standard deviation in Euros than in Punts (respectively .171 versus .137). Weber fractions were slightly correlated across the two currencies (r^2 = 11.7%, p=.031). Although we did not independent estimates of buying frequency and price variability as rated by Irish participants, there was a good correlation between those measures as provided by the French and the Portuguese participants (frequency: r=.51, t=3.69, 39 d.f., p=.0007; variability: r=.77, t=7.47, 39 d.f., $p<10^{-8}$), leading us to apply the average of those measures to the Irish population. In agreement with previous experiments, Weber fractions for prices in Punts decreased significantly with increasing buying frequency (r=-.36, p=.021) and increased significantly with increasing familiarity with the product (r=.34, p=.030). As in most of the above experiments, these correlations were not significant for the Weber fraction in Euros, presumably due to greater response variability, though they were all in the same direction than with the native currency.

Regression analyses across items indicated that the average prices estimated in Euro correlated very well with those estimated in Punts (r^2 = 97.8%). No constant term was needed, indicating direct proportionality of the estimated prices. The regression slope (.855± standard error .018) was significantly larger than the expected value of .787564, which is the exchange rate that prevails between Euros and Punts. This suggested that participants tended to underestimate prices in Euros. However, this was contradicted by the analysis of converted prices, which indicated that after conversion the Euro prices were above the Punt prices by an average of 5.4%, a value that was significant across items (z=2.12, p=.040). Examination of individual items suggests that the discrepancy may be due to a complex interaction with price. For most items, participants tended to overestimate Euro prices, but this tendency was reversed to an underestimation for the few highest-priced items, which are given a much higher weigh in the least-square based regression analysis.

The effect of sex on dispersion was replicated for participants responding in their familiar currency, females participants having dispersion scores 17.3% higher than males (r=.42, t=1.96, 18 d.f., one-tailed p=.033). No other effect approached significance. Participants judged that they had responded without calculating more often when working in their familiar currency than when working in Euro (76.5% versus 57.0% of trials, t=1.88, 38 d.f. one-tailed p=.034). They also rated their intuition of prices as marginally higher in their familiar currency than in Euro (5.1 vs. 4.35, t=1.59, 38 d.f., one-tailed p=.060).

Discussion

Nearly all the findings with French subjects were replicated with Irish subjects. Estimated prices followed Weber's law, and the Weber fraction jumped from .35 in the native currency to .44 in the Euro. Response speed was not measured, but subjects again claimed to have to calculate more when responding in Euros. Finally, the sex effect on dispersion was replicated, males being more accurate than females. The only prediction that failed concerned subjects' biases. A complex pattern of results was found, with overestimation in Euros for most items (contrary to our initial predictions) and underestimation for a few high-priced items. We have no ready explanation for this pattern, especially considering that it is not found in other countries. Many factors may potentially bias price estimation in Euros, including the roundness of specific prices (a factor that probably influences price estimation in the native currency as well, as when judging that 599 is a smaller price than 600); the particular strategy used for price conversion (estimation or calculation; and the exact nature of the calculation used) ; poor knowledge and/or rounding of the exchange rate; and even, perhaps, feelings of superiority or inferiority relative to the Euro. More research will be needed to sort out the contributions of those factors.

GENERAL DISCUSSION

Our first main finding is that subjects from different countries are highly consistent when performing price estimation tasks. Regardless of the details of the task (timed or not timed; spontaneous price production or forced choice), a linear relation is systematically found, across items, between the estimated price and the standard deviation of this price. This is the signature of Weber's law. Our results extend Whalen et al.'s (1999) work by showing that this law holds over several orders of magnitude and, more importantly, holds even in a situation where the inputs and the outputs are symbolic. The systematic variability observed in our estimation task cannot arise from failures at the comprehension or response levels, but must come from one or more of the internal representations involved.

Information about the nature of those internal representations comes from two sources. First, examination of the shape of the response distributions revealed that the responses are distributed log-normally, i.e. they followed a classical Gaussian curve only when plotted on a logarithmic scale (Figure 4). Interpretation should be cautious at this stage because although the effect could be due to the fact that most experiments involved a multiple choice situation in which the proposed alternatives themselves followed a geometric progression and might have biased the results towards a logarithmic encoding. The only experiment in which the responses were open-ended (experiment 2) involved too few trials to enable evaluation of the shape of the distributions. Nevertheless, this finding, if validated by further work, provides a new way to evaluate the internal representation of numerical quantities and seems to suggest that the internal 'number line' is best modeled by a logarithmic(Dehaene, 1989; Dehaene, 2001) than by a linear function(Brannon et al., 2000).

A second source of information comes from a close examination of the value of the Weber fraction. Across different countries, the average value of the Weber fraction in the native currency is always close to .35. However, across items, this value varies with buying frequency (more familiar items have a smaller Weber fraction) and with estimated price variability (items whose price is judged as more variable have a larger Weber fraction). This implies that subjects mentally represent both parameters. It also means that a large part of the observed variability does not come from an intrinsic randomness in the representation of numbers, but from properties of the mapping from items to prices. Nevertheless, even frequently encountered items whose price is highly invariable have a non-zero Weber fraction, with a value close to .15. This value, which is close to that observed by Whalen et al. (1999), may represent the intrinsic variability associated with the number system itself.

These findings suggest a possible scenario for the acquisition of price expertise. According to this hypothesis, each encounter with a product and its price leads to the creation of an association between the mental representation of the product and the corresponding number, represented as a distribution of activation at the appropriate location on the number line. The width of this distribution is determined by the Weber fraction of about .15, which characterizes the intrinsic variability of the mental quantity representation. After exposure to multiple product-price pairs, the strength of the mapping builds up in direct proportion to the frequency with which a given product is bought. However, its dispersion can also become wider than the initial Weber fraction if there has been significant variation in the range of prices encountered for a given product. In the end, the mental mapping from any product to the number line representation of its price is therefore characterized by a strength factor, which varies with buying frequency, and a dispersion, which reflects both the actual variability of prices in the environment and a minimal residual variability due to the Weberian property of the number line itself.

Note that the postulated mapping links each product directly to a given number, expressed in a specific currency. According to the proposed model, there is no abstract semantic representation of prices independently of the currency used to express them. This can readily explain why subjects fail to generalize their price knowledge when the currency is changed. All they have is a sense of which range of numbers is approximately appropriate; but those numbers are simply inadequate in the new currency. Indeed, our results reveal that the Weber fraction is significantly higher, by about 50%, when subjects rate prices in an unfamiliar currency. Even this poor performance is achieved at the expense of speed, as subjects avowedly have to calculate to compensate for their lack of price intuition. If the proposed model is correct, such intuition can only develop by a slow process of association that requires exposure to many product-price pairs. Hence, our experiments support the intuitive idea that a change in currency, although seemingly computationally trivial, carries a significant psychological toll in the form of a temporary loss of price intuition.

Another drawback for large-scale currency changes is posed by the presence of systematic biases in subjects' perception of the new prices. We found that French subjects overestimated prices in Euro by as much as 7%, suggesting that they would not notice if a shop owner took advantage of the switch to the Euro to increases prices by this amount. In Portugal, subjects underestimated prices in Euro, suggesting that they would find the new prices too expensive. Finally, in Ireland, a complex interaction with price with found, with overestimation for cheap items and underestimation for expensive items. We have not found a systematic explanation for these variations across different countries, though we suspect that the direction and roundness of the exchange rate as well as the conversion strategies adopted play a significant role.

A significant effect of sex on dispersion was observed repeatedly in French subjects and was replicated in Ireland, though not in Portugal. Females' estimations were more variable than males by about 20%. It seems possible that, for educational or biological reasons, the mental representation of numbers is more accurate in males, as suggested by several large-scale studies of sex effects in mathematics(Benbow, 1988; Hyde et al., 1990). Another possible interpretation is that females actually possess amore accurate sense of the range of possible prices for a given item and hence propose a greater variety of responses. Thus, this intriguing sex effect is currently ambiguous and is reported here as an exploratory finding that should be replicated in more formal studies, controlling in particular for the neutrality of the selected items with respect to sex.

Finally, a clear limitation of our work is that we have only worked with subjects that have had very little experience with the new currency, during a time period in which the introduction of the Euro was only partial. Many questions therefore remain open to future investigations (see also Marques, 1999). How will Weber fractions evolve following the official introduction of the Euro in January 2002? How long will it take for price intuition in Euros to reach present levels? Will the speed of adaptation be the same for all age groups, or will younger subjects adapt faster? The tasks and statistical tools introduced here provide a methodology with which to evaluate these questions.

REFERENCES

Benbow, C. P. (1988). Sex differences in mathematical reasoning ability in intellectually talented preadolescents: Their nature effects and possible causes. Behav. Brain Sci., 11, 169-232.

Brannon, E. M., Wusthoff, C. J., Gallistel, C. R., & Gibbon, J. (2000). Numerical subtraction in the pigeon: Evidence for a linear subjective number scale. Psychol. Sci., in press.

Butterworth, B. (1999). The Mathematical Brain. London: Macmillan.

Dehaene, S. (1989). The psychophysics of numerical comparison: A re-examination of apparently incompatible data. Percept. Psychophys., 45, 557-566.

Dehaene, S. (1992). Varieties of numerical abilities. Cognition, 44, 1-42.

Dehaene, S. (1993). Numerical Cognition. Oxford: Blackwell.

Dehaene, S. (1997). The number sense. New York: Oxford University Press.

Dehaene, S. (2001). Subtracting pigeons: logarithmic or linear? Psychol. Sci., 12, 244-246.

Dehaene, S., Dehaene-Lambertz, G., & Cohen, L. (1998). Abstract representations of numbers in the animal and human brain. Trends Neurosci., 21, 355-361.

Dehaene, S., & Mehler, J. (1992). Cross-linguistic regularities in the frequency of number words. Cognition, 43, 1-29.

Grewal, D., & Marmorstein. (1994). Market price variation, perceived price variation, and consumer's price search decisions for durable goods. Journal of Consumer Research, 453-460.

Hyde, J. S., Fennema, E., & Lamon, S. J. (1990). Gender differences in mathematics performance: A meta-analysis. Psych. Bull., 107, 139-155.

Lambert, Z. V. (1978). Differential thresholds in consumer perception of retail prices. The Journal of Psychology, 1978(100), 139-150.

Lemaire, P., & Lecacheur, M. (2001). Older and Younger Adults' Strategy Use and Execution in Currency Conversion Tasks: Insights from French Franc to Euro and Euro to French Franc Conversions. Journal of Experimental Psychology: Applied, in press

Lemaire, P., Lecacheur, M., & Ferréol-Barbey, M. (2001). Strategy Use in Currency Conversion Tasks: The examples of Euros to French Franc and French Franc to Euro Conversions. Current Psychology Letters, 4, 39-50.

Marques, J. F. (1999). Changing “EUROpe”: The euro as a new ground for psychological research in numerical cognition. European Psychologist, 4, 152-156.

Mechner, F. (1958). Probability relations within response sequences under ratio reinforcement. J. Exp. Anal. Behav., 1, 109-121.

Meck, W. H., & Church, R. M. (1983). A mode control model of counting and timing processes. J. Exp. Psychol. Animal. Behav. Processes, 9, 320-334.

Moyer, R. S., & Landauer, T. K. (1967). Time required for judgements of numerical inequality. Nature, 215, 1519-1520.

Müller-Peters, A. (1998). The significance of national pride and national identity to the attitude toward the single European currency: A Europe-wide comparison. Journal of Economic Psychology, 19, 701-719.

Müller-Peters, A., Pepermans, R., Kiell, G., Battaglia, N., Beckmann, S., Burgoyne, C., Farhangmehr, M., Guzman, G., Kirchler, E., Koenen, C., Kokkinaki, F., Lambkin, M., Lassarre, D., Lenoir, F. R., Luna-Arocas, R., Marell, A., Meier, K., Moisander, J., Ortona, G., Quintanilla, I., Routh, D., Scaciati, F., Uusitalo, L., van-Everdingen, Y. M., van Raaij, W. F., & Wahlund, R. (1998). Explaining attitudes towards the euro: Design of a cross-national study. Journal of Economic Psychology, 19, 663-680.

Pallier, C., Dupoux, E., & Jeannin, X. (1997). Expe5: an expandable programming language for on-line psychological experiments. Behavior Research, Methods, Instruments and Computers, 29, 322-327.

Pepermans, R., & Verleye, G. (1998). A unified Europe? How euro-attitudes relate to psychological differences between countries. Journal of Economic Psychology, 19, 681-699.

Webb, E. J. (1961). Weber's law and consumer prices. Amer. Psychol., 16, 450.

Whalen, J., Gallistel, C. R., & Gelman, R. (1999). Non-verbal counting in humans: The psychophysics of number representation. Psychol. Sci., 10, 130-137.

<u>Figure 1. Scalar variability and Weber's law in price estimation</u>A, linear regression between the mean and the standard deviation of the estimated prices across 40 different products in experiment 1 (multiple choice, Francs). The regression curve was linear and constrained to pass through the origin. B, same data and predicted values on a logarithmic scale; C, approximate stability of the Weber fraction (ratio of standard deviation and mean price) over several orders of magnitude.

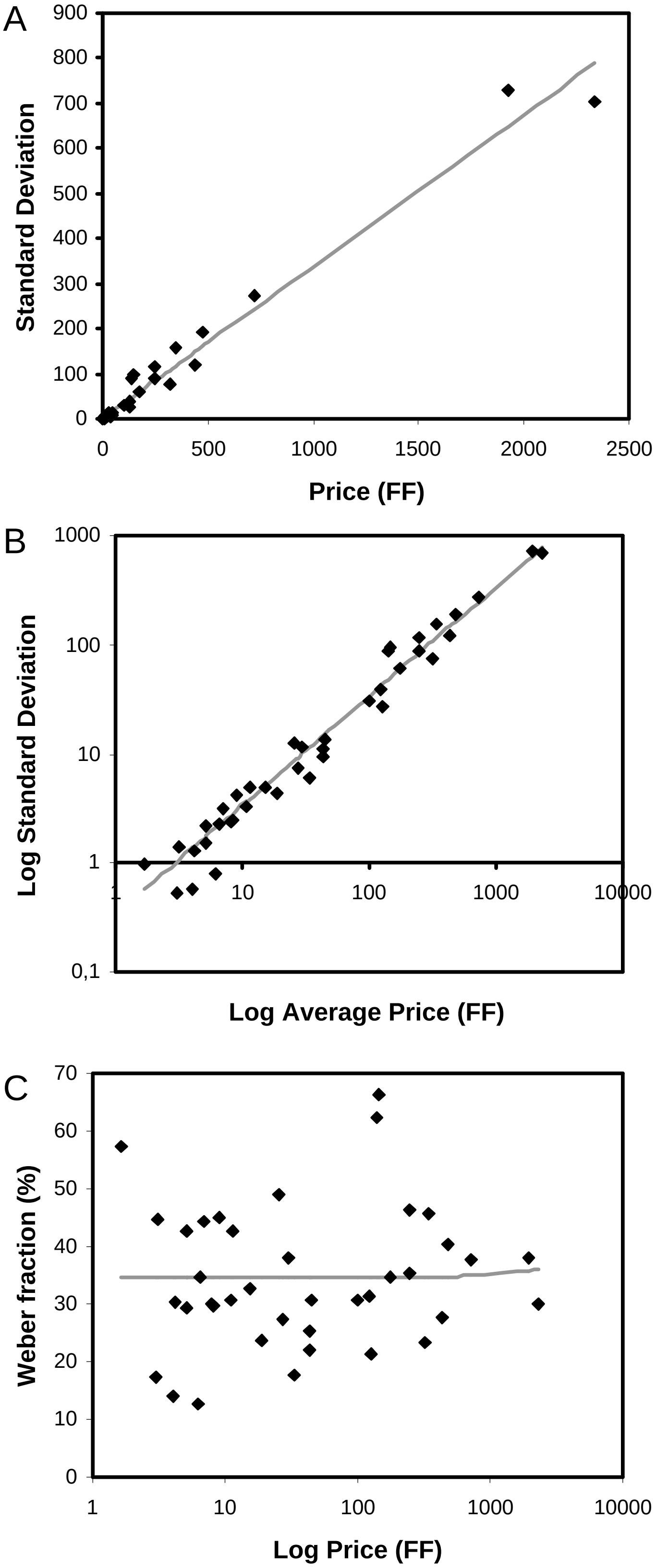
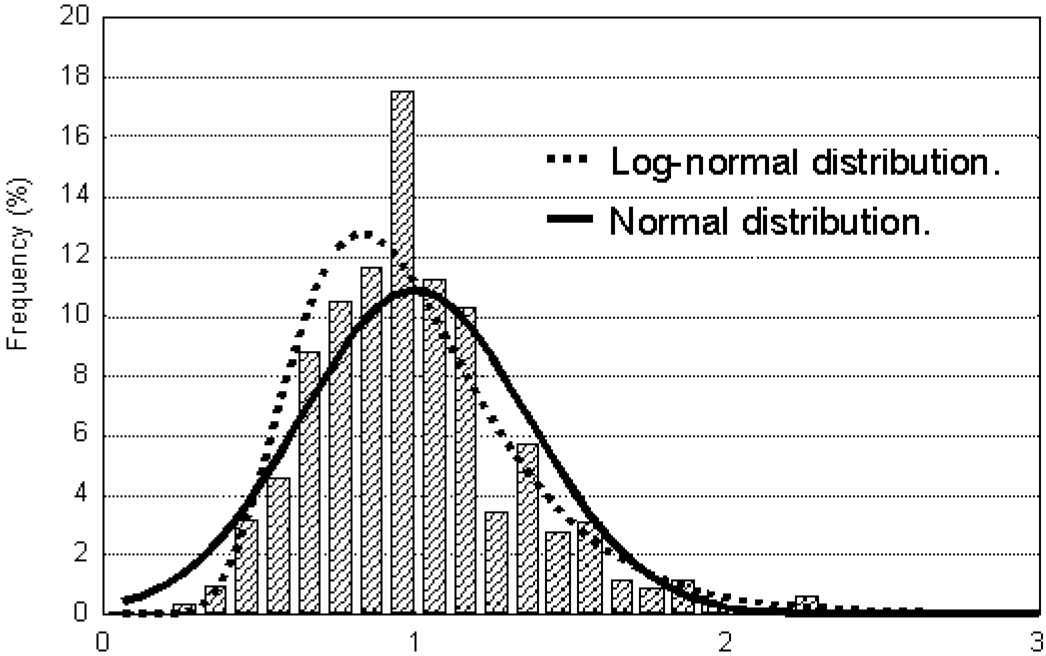
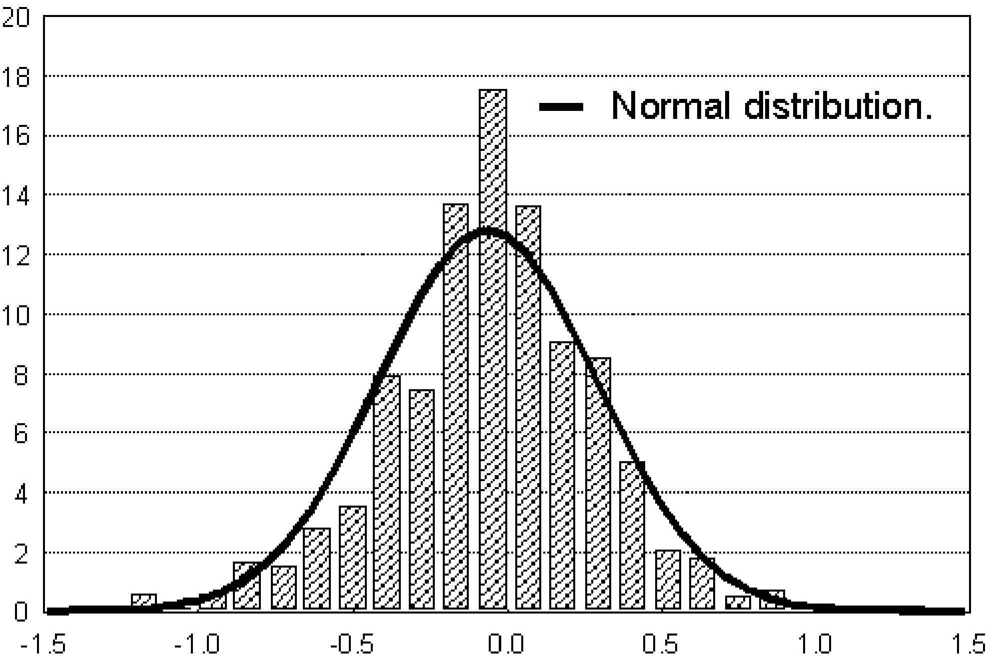


Figure 2. Shape of the distribution of responses. Responses were normalized by dividing by the mean price for each item. On a linear scale (top panel), the distribution is consistently skewed and is better fitted by a log-normal than by a normal curve. On a log scale (bottom panel), the distribution appears symmetrical and essentially normal.

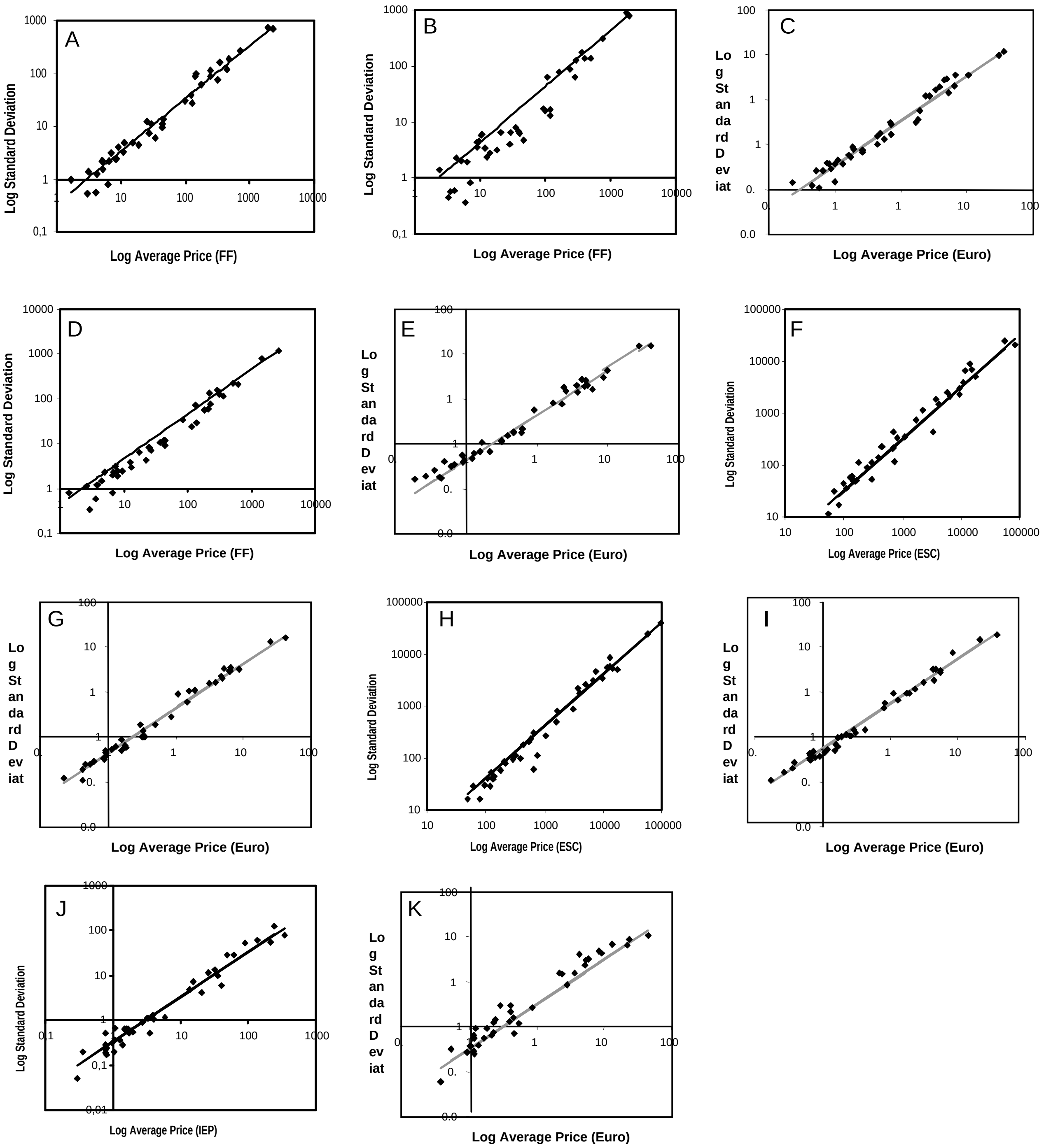


Linear scale for normalized price.



Logarithmic scale for normalized price.

Figure 3. Reproducibility of scalar variability with various currencies, countries, and experimental methods. Format similar to figure 1 B. Note that although the axes are on a logarithmic scale, the regression curve shown is based on a relation of direct proportionality between estimated price and standard deviation (as indicated by a slope of 1 on a log-log scale), not a more general power law. Panel A, experiment 1 (multiple choice, Francs). Panel B, exp. 2 (spontaneous production, Francs). Panel C, exp. 3 (multiple choice, Euros). Panels D and E, exp. 4 (timed choice, Francs and Euros). Panels F, G, H and I, exp. 5 (multiple choice in Escudos and Euros, and timed choice in the same currencies). Panels J and K, exp. 6 (multiple choice, Irish punts and Euros).



<u>Figure 4. Increased variability of price estimates in an unfamiliar currency</u>. The distribution of numerical responses around their mean is plotted on a logarithmic scale (same representation as in figure 2, bottom panel; the fitted curve is a normal distribution). In all experiments in which the familiar and unfamiliar currencies were compared, the distribution is broader in the latter.