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An empirical evaluation of the impact of three urban transportation policies on transit use

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ABSTRACT

The impact on transportation mode choice of policies implementing metro network expansion, fare subsidies and automobile use and ownership regulation was evaluated econometrically using data for 41 world cities. Controlling socioeconomic and demographic variables, it was found that an increase in metro network extension of 10% generates an average decrease in automobile use of 2%. The results also showed that regulation of automobile use or ownership leads to a significant rise in public transit use. By contrast, no evidence was discovered suggesting that transit fare subsidies produce significant increases in transit ridership.

1. Introduction

The main objective of this paper is to study empirically the impact of three urban transport policies: (i) expansion of a metro or urban rail transit network, (ii) subsidization of public transit fares, and (iii) regulation of automobile use and ownership. This study uses cross-sectional data collected for 41 cities in different countries to evaluate the impact on public transit use of these three urban transportation policies.

The specialized literature has addressed this issue on several occasions, however, the results are not conclusive. Typically, the methodologies are based on local analysis focused on specific public transport systems. In this paper we answer this question using aggregate data from different cities of the world applying the methods of multiple regression analysis and controlling for socioeconomic and demographic variables (similar to the approach employed in Pucher et al., 1983).

To test the robustness of our conclusions we defined four versions of a base econometric model, each one using a different dependent variable to represent public transit use. Also, to ensure that the statistical non-significance of a given estimate was not due to collinearity we ran four separate specifications of each version, three of which regressed on a single independent variable representing one or other of the three transit policies while the fourth did so on all three variables. Thus, a total of 16 model formulations were utilized, all of which yielded consistent results.

Although we control relevant characteristics of the cities, there may be the concern with the relevance of omitted controls. To show that our models seem to be well specified, we report the p-values of the Ramsey Regression Equation Specification Error Test. We cannot reject the null hypothesis that the models do not have omitted variables in any case.

The database used in our analyses was constructed with information taken from various sources and references for different metro and rail transit systems around the world. Included in our sample were systems in cities for which comparable data was available on all of the variables in our model, and particularly on the modal split between public and private transportation and the public transit operating subsidy.

We generated a set of estimates that led us to the following conclusions: first, a 10% metro or rail network expansion produces a reduction in automobile use averaging more than 2%; second, there appears to be no evidence that fare subsidies stimulate transit ridership or reduce car use; and third, regulation of car use or ownership brings about a significant increase in transit demand. These findings are consistent with those previously reported in the specialized literature for the three policies.

In the remainder of this article, Section 2 reviews the literature on the three policies under analysis; Section 3 introduces our econometric model and its various formulations and summarizes our results; and Section 4 presents our main conclusions together with some recommendations suggested by the empirical analyses we carried out. Finally, the data used for deriving the estimates are set out in graphical format in the Appendix A.

2. Review of the literature

In what follows we review the evidence contained in previously published works on the impacts of metro network

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expansion, fare subsidies and automobile use regulation on certain aspects of public transit.

Existing studies indicate that although expanding metro lines involves high levels of investment and capital costs, it is usually successful in reducing the progressive rise in private vehicle use and reversing trends toward lower participation in public transit.

Comparing cities within the US, Litman (2009) found that transit ridership per-capita is up to five times larger in cities with metro or light rail transit (LRT) than in cities operating only with buses. Litman (2009) also reported lower motorization rates in cities with metro or LRT, despite their usually higher per-capita income. Schumann (2005) compares the evolution of Sacramento that introduced a LRT in 1985, with Columbus, a city of similar size that maintained a bus-only system. Over the following 17 years, service levels and passenger flows rose significantly in Sacramento whereas in Columbus they declined. Henry and Litman (2006) studied US cities between 1996 and 2003, finding that cities with metro or LRT displayed greater growth in transit use than bus-only cities. In a similar way, Litman (2005) found that in households located near metro stations, automobile ownership was lower than in urban zones without metro service. Baum-Snow and Kahn (2005) demonstrated that although commuting on public transit has waned over recent decades, the decline in cities with a metro is 20% to 23% whereas in cities without a metro the figure is more than 60%.

Regarding the effects of expanding the metro or LRT system, Bento et al. (2003) estimated that a 10% expansion of a metro network reduced automobile use by 4.2%, a greater decline than has been generated by any road pricing scheme implemented to date anywhere in the world. While, Winston and Langer (2004) showed that road congestion decreases in cities that expand rail transit, but increases where bus service is improved. This appears to occur because buses attract fewer motorists, compete for existing road space with other vehicles (thus making their own contribution to congestion) and have less positive impact on land use accessibility.

Empirical evidence indicates that a metro or urban rail transit system generally tends to reduce trips in private vehicles for two reasons: first, due to increased transfers of car users to the metro, and second, the reduced acquisition of cars in households located in urban areas close to metro stations. Research by Neff (1996) and Newman and Kenworthy (1999) reveals that each passenger-mile that is transferred from automobiles to rail transit represents an average reduction of between 1.4 and 9 vehicle-miles.

Results are inconclusive regarding the promotion of public transit use through the subsidization of fares. A series of studies dating back to the late 1970s and early 1980s reported that the direct benefits of such policies to riders were minimal and that environmental and secondary benefits were either negligible or non-existent (Altshuler et al., 1981; Meyer and Gómez-Ibáñez, 1981; Hilton, 1974; Hamer, 1976; Webber, 1976).

Some critics of fare subsidies suggest that they simply inflate costs without generating better or cheaper services (Altshuler et al., 1981; Bonnell, 1981; Gwilliam et al., 2004). Pucher et al. (1983) argue that this approach may even increase transit companies' operating costs while reducing their productivity. Karlaftis and McCarthy (1998) concluded that subsidies had no general effect on the performance of the transit system of Indiana. Phillipson and Willis (1990) found that in the Australian city of Adelaide, a free transit service (i.e., a 100% reduction in fares) would probably boost patronage by 30%. However, only a very small share of this would be from private car users and it was estimated that automobile trips would decline by less than 2%. Similarly, Robert and Jonsson (2006) estimated for Stockholm a free public transit would not bring a substantial reduction of private vehicle use or its associated CO_2 emissions. Rather, they discovered that the primary result of the measure would be the migration to the transit system of foot and bicycle trips. Hensher (2008) and Noland et al. (2006) also found that fare subsidies would do little to reduce CO_2 emissions associated to private vehicles.

In contrast, Macharis et al. (2006) studied the case of a free travel policy for students adopted by the Brussels public bus system. Based on survey data, they find a significant increase in the demand for bus trips by students. Findings of Parry and Small (2009) indicate that fare reductions increase social benefits at the margin. However, it does not address the problem of the subsidy's ineffectiveness in reducing private vehicle trips or even in reducing final fares. In a recent paper, Sharaby and Shiftan (2012) found a 7.7% increase in passenger trips and an 18.6% in boardings after the implementation of an integrated bus fare system in Haifa, Israel. Although the findings are consistent with those published by NEA (2003), in the latter analysis the improvement was also partly due to the operational integration of the network and the information. The NEA study also states that the demand effects in the long term tend to be smaller.

As regards policies regulating automobile use, Meland and Polak (1993) found that the implementation of a toll ring in Trondheim, Norway, changed the behavior of travelers, whether in mode choice, time of day, route, destination or frequency. Another Norwegian study discovered that trips within toll rings during toll periods declined from 6% to 7% in Bergen, 8% in Oslo and 10% in Trondheim (Larsen, 1995). Jaensirisak (2003) argued that improvements in public transit do not by itself appear to be enough to capture automobile users. The policy would have to be supplemented by direct restrictions on road traffic and ensure a better balance between demand and supply of road space.

3. The model

Our proposed model in its various formulations are all estimated using cross-sectional data for 41 important cities around the world following the methodology in Pucher et al. (1983). The data used to calibrate the model, presented here in the Appendix A, were drawn from the following sources: Millennium Cities Database (Kenworthy and Laube, 2001), EMTA (2007, 2009), MVA (2005), Patisson (2000), Bristol City Council, the Transport Statistics Bulletin (UK Dept. for Transport website), Jane's Urban Transport Systems 2000–2001 and other official Internet websites of relevant organizations.

It is worth pointing out that we are not interested in finding estimates of the demand and the supply functions for transit separately, or to describe the detailed response of each after changes in market conditions describing endogenous relationships as in Taylor et al. (2009). All three policies that we consider are exogenous to contemporary shocks in markets conditions, and thus our estimates are consistent and reflect long run total effects of the policies on the use of transit. For instance, subway extensions do not occur endogenously during a positive shock in usage. Despite the decision to extend the subway depends on demand conditions, it does not depend on contemporaneous demand shock. Indeed, the construction of subway extensions may well take several years.

3.1. Design of the model

The model attempts to determine empirically the impact on public and private transportation use of the implementation of three urban transportation policies:

i. expansion of the metro or urban rail system
ii. subsidization of transit fares
iii. regulation of private vehicle (car) use and ownership.

The use of public or private transportation may be represented in various ways, such as aggregate modal split or number of daily or annual trips. With the data gathered for this study we were able to determine the relative modal split for each city included in our sample. Thus, we denote π as the proportion of trips taken on public transit and $1-\pi$ as the proportion on private transportation.

To control city characteristics we incorporated data on population, population density, per-capita income, and the number of automobiles per 1000 inhabitants (i.e., the motorization rate).

As regards the variables used as indicators of the three policies, metro or rail line extension is represented by network lengths in kilometers, transit fare subsidies are given by the percentage of operating costs (i.e., excluding investment and capital costs) that are not covered by operating revenues, and the regulation of private vehicle use and ownership is incorporated as a dichotomous variable that takes the value of 0 if a city has no significant regulation policy and 1 if it does. In our sample the cities we consider to have implemented effective regulation of automobile use are: London, Singapore, Oslo and Stockholm, all of which apply cordon tolling, and Hong Kong, Tokyo and Seoul, which use car taxes or value pricing. Despite that other cities in the sample also have some kind of regulation to the use of automobile (as road space rationing or specific taxes to automobiles and gasoline), we consider that in these seven cities the regulation on the tenure and use of automobiles are qualitatively larger than in the remaining cities in the sample.

The base econometric model can now be formulated as:

$$y_i = X_i'\beta + \alpha_1 \ \ln(Kms_i) + \alpha_2 \ \ Regulation_i + \alpha_3 \ \ Subsidy + \varepsilon_i \tag{1}$$

where X_i is the control variables vector (population, density, per-capita income and motorization rate), $\ln(Kms_i)$is the natural logarithm of the metro network length, $Regulation_i$ is the dichotomous variable for automobile use and ownership regulation, $Subsidy_i$ is the percentage of operating costs not covered by operating revenues. The dependent variable $y_i = \ln(Trips_i)$ indicating that the number of transit trips is replaced by a proxy which is the natural logarithm of the proportion π of all trips accounted for public transit multiplied by the city population.

$$\ln(Trips_i) = \ln(\pi_i \ \ Population_i) \tag{2}$$

The advantage of using the natural log of the number of trips as the dependent variable is that the coefficients of Eq. (1) can then be interpreted as elasticities, semi-elasticities or percentage changes.

To test the robustness of our results we also estimate three other versions of the model in which the respective dependent variables are: (i) π; the proportion of all trips on public transit; (ii) $\ln(\pi/(1-\pi))$, the natural log of the ratio of public to private transportation use; and (iii) a proxy of the number of private transportation trips, similar to the proxy for public transit trips:

$$\ln(Priv. \ \ Trips_i) = \ln((1-\pi_i)Population_i) \tag{3}$$

In the model versions employing as the dependent variable the proportion of trips and the log of the ratio of public to private transportation use, we do not include the city population as a control given that the dependent variables are already normalized for this factor.

For each of the model versions we formulate four specifications to estimate the values of the policy variable parameters α_1, α_2 and α_3 in Eq. (1). Three of the specifications each contain one or other of the policy variables and the fourth includes all three of them. The reason for adopting this multiple approach is that with a single specification including the variables, the parameters may lose statistical significance due to possible collinearity even though the corresponding policies are in fact effective. On the other hand, excluding variables from the model could cause bias in the estimates. Thus, with four specifications we can evaluate the robustness of the estimators to the extent that both the estimated parameters and their statistical significance do not change greatly from one specification to the next. The parameter estimations themselves were performed using ordinary least squares (OLS) while the variances were estimated using an estimator robust to heteroscedasticity. In regressions with cross-sectional data it is common for the homoscedasticity of errors assumption to be violated, with the result that the traditional OLS estimator of the variance matrix is very likely inconsistent. To sidestep this problem we used the Huber–White variance estimator, which is robust to the presence of heteroscedasticity (White, 1980).

As for the exogenous variables, it is true that by using cross-sectional rather than panel data we cannot control non-observable individual effects that could be a source of endogeneity to the extent they are correlated with the model's explanatory variables. Nevertheless, for a number of reasons we believe our results are still reliable.

To begin with, by controlling sociodemographic variables (population, density, GDP per capita and motorization rate) we reduced the non-observable individual effects that could potentially be correlated with the explanatory variables and lead to bias and inconsistency in the parameter estimations.

Second, the variables of interest are the three urban transit policies, which depend in large measure on political or other factors that are exogenous to the model and therefore are unlikely to be correlated with non-observable individual effects.

Third, the main source of information with cross-sectional data is the variability of the variables between different cities (between variation), not the variability of them over time within a single city (within variation). Since all of the variables in the model change little over time, estimating using deviations from the variable means (Least Squares Dummy Variable, or LSDV) is not efficient. Indeed, variables such as metro network length or the existence of automobile use regulation may not vary at all in the sample over time, making it impossible to estimate their effects with the LSDV method.

Fourth, even with panel data OLS performs relatively well in practice (Beck and Katz, 1995). Plümper and Troeger (2007) have shown that when explanatory variables in a panel analysis vary little over time (within variation), as is the case here, an OLS estimator performs much better than an LSDV estimator even when the non-observable individual effects are correlated with the explanatory variables. In their simulations the authors found that the root mean squared error (RMSE) of an LSDV estimator is close to double (1.74 times) that of a pooled OLS estimator.

Finally, it would not be possible to construct panel data for the 41 cities and correct for any individual effects simply because in most cases the necessary information does not exist.

On the other hand, there is a positive correlation between GDP/Pop and Cars, which is well documented in the specialized literature (Button et al. (1992); Badoe (2007); Guevara and Thomas (2007); Oyedepo and Makinde (2009)). But this does not change our results. In fact, it is essential that the two variables be specified in the econometric model in order to better control aggregate demand for trips and its separate relationships with them. Excluding either of the variables would introduce endogeneity into the regression.

Besides, to check for the right specification of the model, in all estimations we report the p-value of the Ramsey Regression Equation Specification Error Test (Ramsey, 1969). The Ramsey test is a classic test that checks for the presence of serious specification errors, in terms either of omitted variables or the chosen functional form. It verifies whether non-linear combinations of the explanatory variables have predictive power in regard to the explained variable. In all cases, we do not reject the null hypothesis that the models do not have omitted variables, which supports our econometric approach.

3.2. Results

The results of the estimations of our model using the natural log of the number of transit trips as the dependent variable are shown in Table 1. In all four specifications, the control variables all appear with the expected signs but are not always significant, possibility due to collinearity.

The estimators in the fourth specification, less exposed to bias given that all independent variables are present, indicate that a 10% increase in per-capita income would induce a 2.1% reduction in transit use. They also show that as a city's density grow, the use of transit rises. This is to be expected given that higher density brings with it greater congestion and therefore higher automobile use costs which act as a disincentive. The estimates also generated an elasticity of transit trips with respect to the motorization rate of between −0.229 and −0.314, so that an increase of 10% in the rate reduces transit trips by 2.29% to 3.14%.

The estimator of the marginal effect of the metro network on transit use, α_1 in Eq. (1), is positive and highly significant in all specifications. Its value indicates that on average, an increase of 10% in metro network length induces a rise of approximately 3% in transit trips.

This positive effect on transit ridership is consistent with results reported previously in the literature. Vuk (2005) estimated that an average of 70% of users of a new metro came from buses, 15% were former automobile users and the remaining 15% constituted new demand, that is, trips that previously were not taken in any mode. In a study by Knowles (1996) it was estimated that demand induced by a light rail system in Manchester was greater than 20% while Monzón (2000) found a new Madrid metro line induced a demand increase of 25%. (Golias, 2002) detected that 16% of users of the recently inaugurated Athens metro were former car users, and Monzón (2000) calculated a figure of 26% for the same phenomenon in the case of a suburban train line. In short, more metro leads to more total trips and fewer trips in private vehicles.

Returning to our results, we found that a policy of effective automobile use regulation significantly increased transit patronage while reducing car trips. This is consistent with the findings of Jaensirisak (2003), Wilson (1988), McCarthy and Tay (1993), Luk (1999) and Menon (2000) for the case of Singapore; Larsen (1995) for Oslo; GOL (2000) for London; and Ahlstrand (2001) for Stockholm.

Finally, for the transit fare subsidy policy we found that the parameter is close to zero and not statistically significant. This is consistent with results reported in studies by Vickrey (1980), Hay (1986), Robert and Jonsson (2006), Noland et al. (2006) and Hensher (2008), all of whom concluded that public transit demand is not affected by subsidized fares. The implication is that this demand is very inelastic. A study by Steg (2005) revealed that automobile use is not motivated solely by functional considerations but also by important symbolic and affective factors and therefore has a value far above any subsidy that could be given for taking public transit. Similarly, Dargay (2007) concluded that car users are not sensitive to changes in relative transportation cost changes, a finding that is consistent with empirical evidence that transit fare subsidies are not sufficient to induce them to change transportation modes.

The estimates for the model versions using the other three alternatives for the dependent variable are shown in Tables 2–4.

Table 1
Parameter Estimation, Goodness of Fit and Ramsey test

Dep. variable	Ln(*Trips*) Model 1	Ln(*Trips*) Model 2	Ln(*Trips*) Model 3	Ln(*Trips*) Model 4
Ln(*Population*)	1.072[a]	1.211[a]	1.305[a]	1.079[a]
	(0.104)	(0.072)	(0.091)	(0.094)
Ln(GDP/Pop.)	−0.132	−0.066	0.003	−0.208[a]
	(0.083)	(0.081)	(0.099)	(0.075)
Ln(Density)	0.181[a]	0.103	0.112	0.199[a]
	(0.097)	(0.087)	(0.106)	(0.092)
Ln(Cars)	−0.265[a]	−0.229[a]	−0.314[a]	−0.243[a]
	(0.078)	(0.074)	(0.087)	(0.075)
Ln(*Kms*)	0.291[a]			0.251[a]
	(0.112)			(0.115)
Regulation		0.31[a]		0.305[b]
		(0.141)		(0.18)
Subsidy			0.002	0.004
			(0.004)	(0.003)
Constant	0.277	−1.077	−2.099	0.305[b]
	(1.89)	(1.431)	(1.714)	(0.18)
Observations	41	41	41	41
R^2	0.91	0.90	0.89	0.92
Ramsey test	0.36	0.78	0.72	0.73

[a] Standard errors in parentheses: at the 5% level.
[b] Standard errors in parentheses: indicates significance at the 10% level.

Table 2

Dep. variable	π Model 1	π Model 2	π Model 3	π Model 4
Ln(GDP/Pop.)	−8.938[a]	−9.633[a]	−6.09[b]	−12.315[a]
	(3.033)	(2.596)	(3.4)	(3.246)
Ln(Density)	6.298	3.733	3.834	6.748[b]
	(3.86)	(3.632)	(4.704)	(3.818)
Ln(Cars)	−7.988[a]	−1.036	−1.937	−6.438[a]
	(3.293)	(3.043)	(3.71)	(3.223)
Ln(*Kms*)	14.348[a]			12.085[a]
	(4.273)			(4.211)
Regulation		20.041[a]		15.517[b]
		(7.759)		(8.995)
Subsidy			−0.071	0.133
			(0.172)	(0.127)
Constant	131.872[b]	126.71[a]	108.471[b]	15.517[b]
	(75.076)	(50.032)	(58.393)	(8.995)
Observations	41	41	41	41
R^2	0.35	0.26	0.12	0.41
Ramsey test	0.78	0.70	0.73	0.67

[a] Standard errors in parentheses: at the 5% level.
[b] Standard errors in parentheses: indicates significance at the 10% level.

Table 3

Dep. variable	$\text{Ln}(\pi/(1-\pi))$ Model 1	$\text{Ln}(\pi/(1-\pi))$ Model 2	$\text{Ln}(\pi/(1-\pi))$ Model 3	$\text{Ln}(\pi/(1-\pi))$ Model 4
Ln(GDP/Pop.)	−0.397[a]	−0.433[a]	−0.258[b]	−0.552[a]
	(0.137)	(0.118)	(0.15)	(0.146)
Ln(Density)	0.3[b]	0.18	0.18	0.315[b]
	(0.178)	(0.167)	(0.218)	(0.174)
Ln(Cars)	−0.368[a]	−0.045	−0.088	−0.29[a]
	(0.15)	(0.137)	(0.168)	(0.146)
Ln(*Kms*)	0.663[a]			0.549[a]
	(0.198)			(0.19)
Regulation		0.962[a]		0.75[b]
		(0.37)		(0.419)
Subsidy			−0.004	0.006
			(0.008)	(0.006)
Constant	3.532	3.324	2.441	0.75[b]
	(3.359)	(2.173)	(2.558)	(0.419)
Observations	41	41	41	41
R^2	0.36	0.27	0.12	0.43
Ramsey test	0.88	0.19	0.16	0.93

[a] Standard errors in parentheses: at the 5% level.
[b] Standard errors in parentheses: indicates significance at the 10% level.

The values are consistent with those in Table 1, demonstrating that the results are quite robust to the model specification.

More specifically, the values in Table 2 indicate that with a 10% extension of the metro network, the proportion of all trips taken on public transit increases between 1.1% and 1.4%. Adopting car use regulation boosts the proportion by 20% while a transit fare subsidy, as with the number of trips estimates (Table 1), has no significant effect.

The estimates presented in Table 3 demonstrate that with the ratio of transit to private trip proportions as the dependent variable, metro expansion and car use regulation once again have significant effects while a fare subsidy has little. The estimators, however, are more difficult to interpret in this case. Since we are working with just two alternatives (public or private transport) with π representing the proportion of trips on public transport, the variable π/(1-π) in Table 3, known in the specialized literature as the odds ratio, is in our case the ratio between the probabilities of traveling on public and on private (i.e., not public) transport. Thus, this ratio expresses how much more likely a public transport trip will occur than a private one. If we have two alternative scenarios (0 and 1), for example, $\ln[\pi_0/(1-\pi_0)]-\ln[\pi_1/(1-\pi_1)]=\beta[X_0-X_1]$ where the ratio $\ln[\pi_0/(1-\pi_0)]-\ln[\pi_1/(1-\pi_1)]$ is referred to as the log of the odds ratio, and therefore $[\pi_0/(1-\pi_0)]/[\pi_1/(1-\pi_1)]=\exp(\beta[X_0-X_1])$. β can be interpreted as the log of the odds ratio for the likelihood that two scenarios (0 and 1) will differ by one unit with respect to the independent variable X (for example, the effect of an additional kilometer of metro between the two scenarios: $X_0-X_1=1$). The term $\exp(\beta)$ is just the odds ratio itself for the likelihood that two situations differ by a single unit of X.

Table 4

Dep. variable	Ln(*Priv. Trips*) Model 1	Ln(*Priv. Trips*) Model 2	Ln(*Priv. Trips*) Model 3	Ln(*Priv. Trips*) Model 4
Ln(*Population*)	0.877[a]	0.777[a]	0.671[a]	0.894[a]
	(0.146)	(0.091)	(0.129)	(0.129)
Ln(GDP/Pop.)	0.157	0.128	0.02	0.252[a]
	(0.109)	(0.089)	(0.119)	(0.096)
Ln(Density)	−0.066	0.007	0.01	−0.077
	(0.114)	(0.09)	(0.121)	(0.102)
Ln(Cars)	0.204[b]	0.139	0.237[a]	0.151
	(0.116)	(0.093)	(0.121)	(0.103)
Ln(*Kms*)	−0.279[a]			−0.221[b]
	(0.13)			(0.131)
Regulation		−0.436[a]		−0.43[b]
		(0.212)		(0.241)
Subsidy			−0.0002	−0.003
			(0.003)	(0.003)
Constant	−1.476	−0.573	0.679	−0.43[b]
	(2.017)	(1.431)	(1.799)	(0.241)
Observations	41	41	41	41
R^2	0.84	0.84	0.82	0.86
Ramsey test	0.90	0.24	0.21	0.91

[a] Standard errors in parentheses: at the 5% level.
[b] Standard errors in parentheses: indicates significance at the 10% level.

When the number of private vehicle trips is the dependent variable, as shown in Table 4, a 10% metro expansion produces a reduction of approximately 2% in automobile use. This is conceptually similar to the result given in Bento et al. (2003) and points up the marked effectiveness of this policy in reducing private traffic. As for car use regulation, in cities where it has been implemented such use is significantly lower. Finally, transit fare subsidies have no significant effects on this factor.

Thus, in none of the model specifications summarized in the four tables did we find any statistically significant relationship between fare subsidies and public or private transportation use. This result is plausible if we recall that a regular automobile user incurs costs over the long run of more than US$20 a day. In the face of such an amount, a fare subsidy is not an attractive incentive for switching one's choice of transportation mode. The average transit fare, as Cox (2002) observed, is not a relevant factor for car users. Consistent with this finding, Robert and Jonsson (2006) found that transit subsidies primarily capture persons who would otherwise walk or cycle rather than private vehicle drivers.

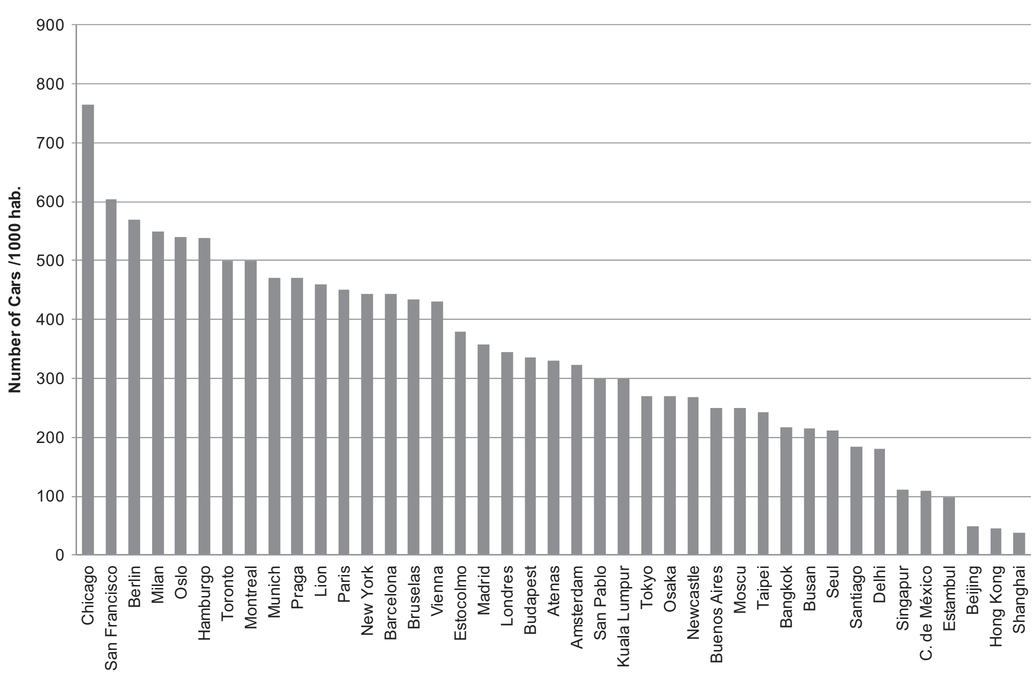

Fig. A1. Automobile ownership, by city.
Sources: Millennium Cities Database (Kenworthy and Laube, 2001), EMTA (2007, 2009), MVA (2005), Patisson (2000), Bristol City Council, Transport Statistics Bulletin (UK Dept. for Transport website).

4. Conclusions

A regression analysis was performed on cross-sectional data for 41 cities around the world to evaluate the impact of metro network expansion, fare subsidies and automobile use regulation on transportation mode use. The econometric model employed controlled socioeconomic and demographic variables on each city such as population, density, per-capita income and motorization rate. Various model formulations were defined and estimated to test the robustness of the results. Ramsey specifications tests cannot reject the null hypothesis of no omitted variables in any of the models.

The estimates produced by the regression analysis led us to conclude that a policy of expanding metro or train networks stimulates the use of public transit. On an average, a 10% extension of a city rail network generates an increase in transit

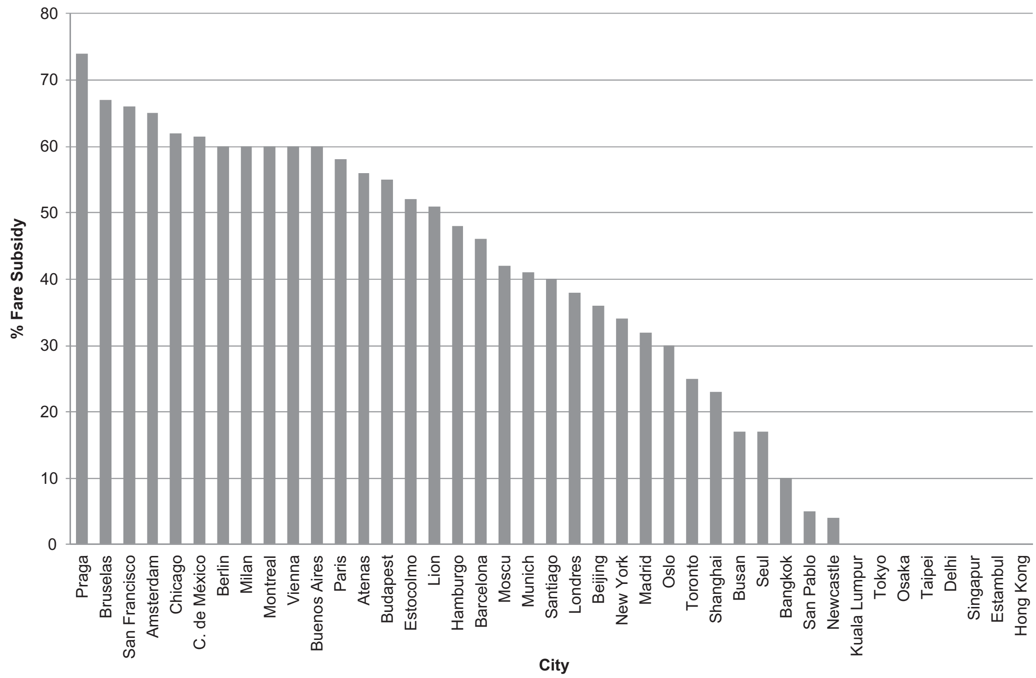

Fig. A2. Fare subsidies (share of operating costs not covered by revenues), by city.
Source: Jane's Urban Transport Systems 2000–2001, FTA (2007), EMTA (2007, 2009), MVA (2005) and other official websites of relevant organizations.

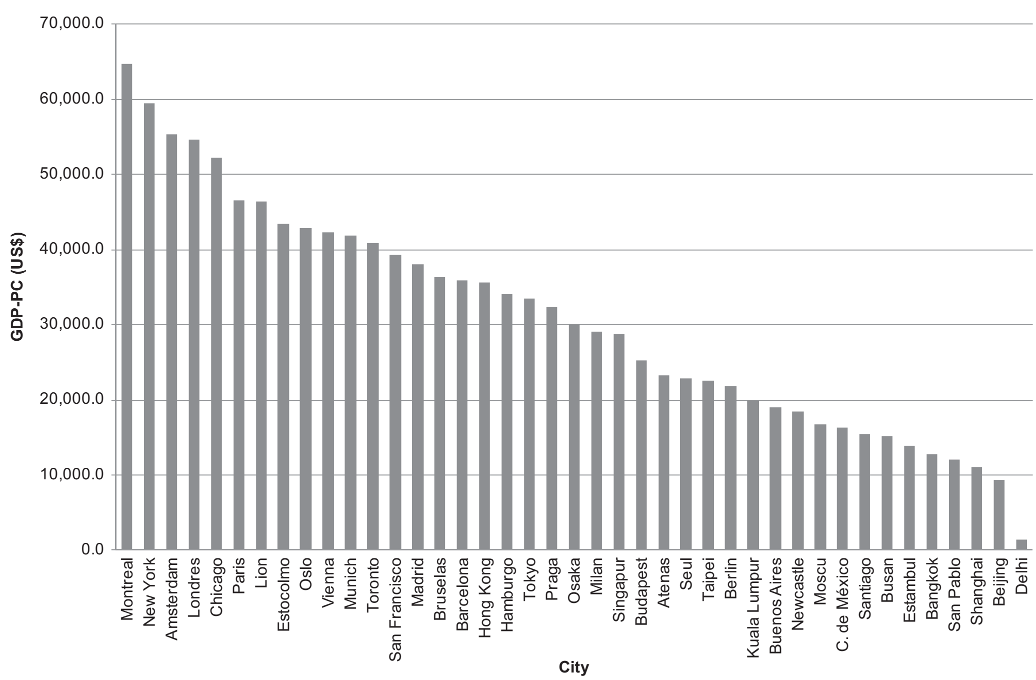

Fig. A3. GDP per capita, by city.
Source: derived from data obtained at various official websites of relevant organizations.

use of almost 3% and a decrease in automobile use of more than 2%.

We further concluded that regulation of automobile use and possession via policies such as road pricing and taxes on car acquisition also have a positive impact on transit patronage. In cities that have implemented effective regulation of this type, the use of cars has dropped by an average of 20–30% while transit use has risen in similar proportions.

Lastly, we found no evidence that fare subsidies encourage the use of transit as an alternative to private cars, corroborating previously published research on the issue. This does not imply that such subsidies should be eliminated, but it does indicate that they will not reduce automobile use. The funding involved could be reallocated to the financing of metro or rail network extensions, which as we have just seen, do promise to generate a significant impact on the use of public and private transportation.

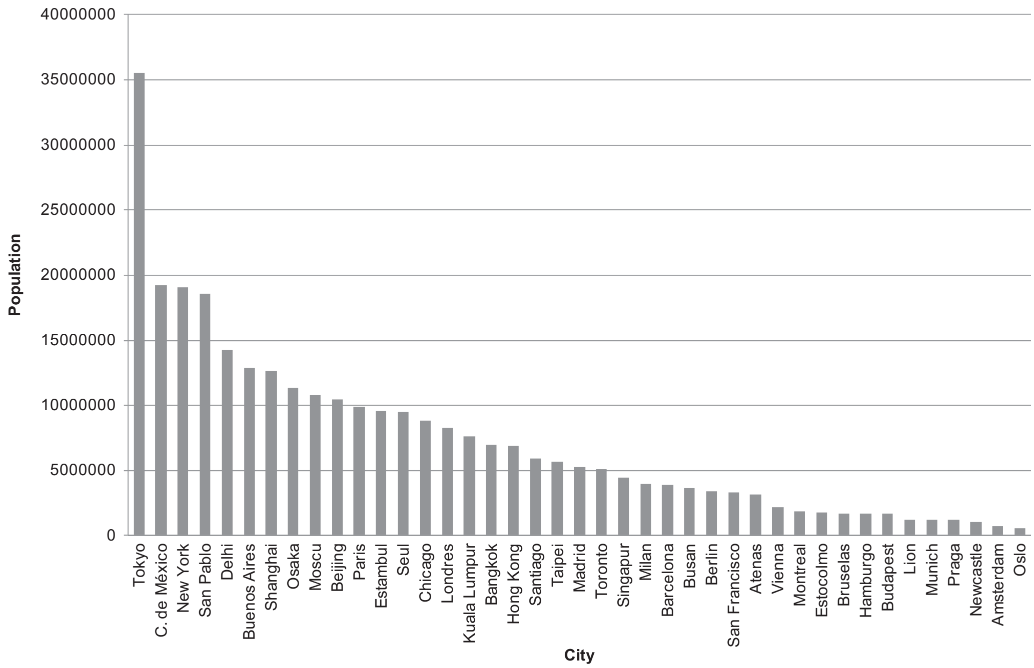

Fig. A4. Population, by city.
Source: derived from data obtained at various official websites of relevant organizations.

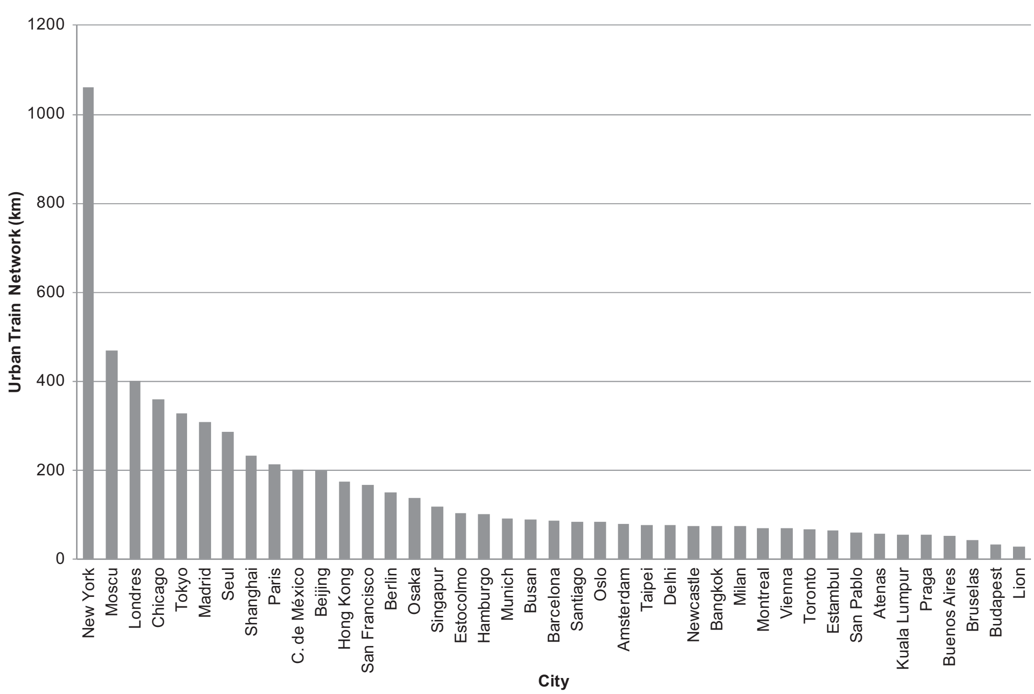

Fig. A5. Urban metro or rail network extension (kms), by city.
Source: derived from data obtained at various official websites of relevant organizations.

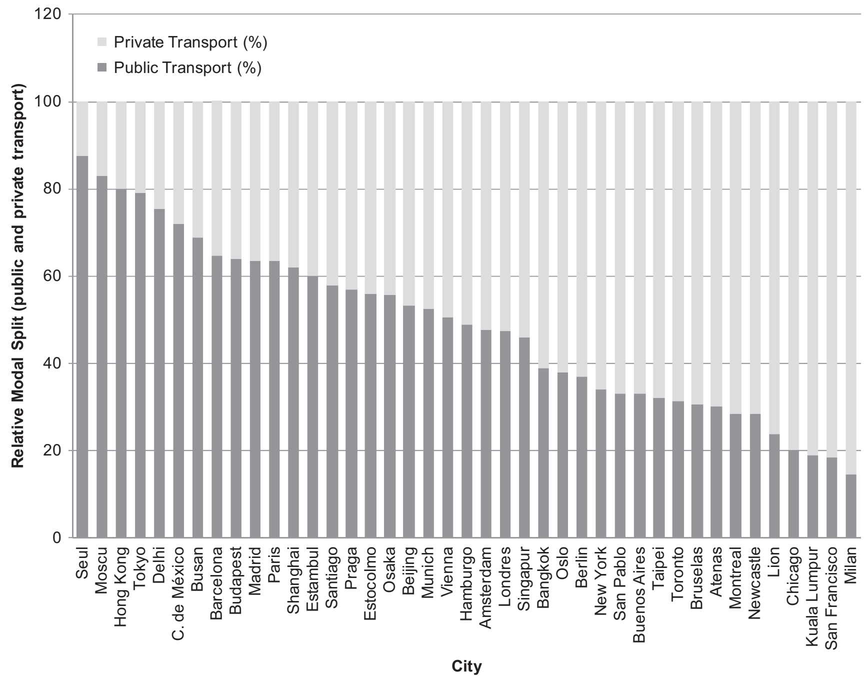

Fig. A6. Private and public transport relative modal split, by city.
Sources: Millennium Cities Database (Kenworthy and Laube, 2001), EMTA (2007, 2009), MVA (2005), Patisson (2000), Bristol City Council, Transport Statistics Bulletin (UK Dept. for Transport website) and other official websites of relevant organizations.

Appendix A. Data and Information Sources

(See Appendix Figs. A1–A6).

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