Border Zone Mass Transit Demand in Brownsville and Laredo

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This study examines whether economic conditions in Mexico influence public transportation ridership levels in the border cities of Brownsville and Laredo, Texas. Besides the standard variables generally utilized to model bus ridership, additional indicators included in the empirical analysis are northbound pedestrian traffic and the real exchange rate index. Seemingly unrelated regression parameter estimates suggest that the volume of pedestrian border crossings in both cities is positively related to changes in ridership. The real exchange rate index in Laredo is negatively related to fluctuations in ridership, implying that peso appreciation increases transit utilization in this border city.

INTRODUCTION

Advocates of public transportation argue that it can play an important role in reducing air pollution and traffic congestion. Such outcomes can occur if it can reduce personal automobile usage. There is speculation that higher gasoline prices might convince people to ride public transit rather than drive. For instance, a local newspaper in south Texas recently attributed an observed increase in McAllen bus ridership to higher fuel prices (Janes 2011). Similarly, a higher level of transit service or lower fares may persuade more people to use public transportation. Econometric analysis may help provide accurate information about the extent to which these and other factors influence ridership.

Aside from the potential benefits from reduced automobile traffic, public transit is also promoted as a means of increasing mobility when using private automobiles is not a viable alternative. Much of the attention in this regard has centered on the role of transit in linking unemployed people with employment opportunities in urban areas (Blackley 1990, Hughes 1991). Less understood is the role of public transportation in linking pedestrian border crossers with destinations such as shopping centers in cities located at international boundaries. In some border zones, public transportation has been proposed as a way of facilitating cross-border shopping (Dascher and Haupt 2011). Such projects could produce economic benefits in areas such as the southern border of Texas, where Mexican shoppers constitute a significant portion of retail trade (Coronado and Phillips 2007). This study examines a variety of factors that may affect transit demand, including the impact of cross-border traffic and international economic factors on bus ridership.

The sample includes data from two U.S. border cities: Brownsville and Laredo, Texas. Several factors make these cities good candidates for a study of cross-border impacts on transit ridership. First, as of 2009, they jointly receive more than 6.6 million pedestrian crossings per year. That is 35% of total pedestrian crossings into Texas and 16% of all pedestrian crossings into the United States (BTS 2011). Second, the central municipal bus terminal in each city is within walking distance of the U.S.–Mexico border, making transit accessible to pedestrians crossing the border. Third, both cities have well-established public transportation systems dating back to at least 1978, making time series analysis feasible. There is a long-standing assumption that pedestrian border crossers form a substantial portion of transit ridership in these cities (TDHPT 1979). One purpose of the present effort is to determine whether, and to what extent, such claims can be quantified.

The demand for public transportation is affected by fares, the level of service, urban population, income, car ownership, and the price of substitutes such as automobile travel. These determinants of ridership, along with the impact of cross-border traffic, provide the basic inputs to the analysis of transit demand presented here. This section is followed by a review of relevant literature and a description of the methodology utilized. Empirical results are presented next, followed by a conclusion.

LITERATURE REVIEW

Much of the literature on determinants of the demand for public transportation concerns variables that are controlled by transit system administrators, such as the level of transit service or fares. Several studies show that changes in these variables can have a substantial impact on transit ridership (Kain and Liu 1999, Taylor et al. 2009). The demand for public transportation is, in most cases, inelastic with respect to fares (Pham and Linsalata 1991, Paulley et al. 2006), although fare elasticities may vary substantially across regions (Dargay and Hanly 2002). Some studies suggest that transit demand is also inelastic with respect to the level of service (Paulley et al. 2006), although other evidence points to an elastic relationship (Holmgren 2007).

The level of service is often measured by a variable such as vehicle revenue hours, which is correlated with the density of bus routes and the frequency of departures. Service, however, is inversely related to the time costs of using public transportation, i.e., the costs of time spent walking to boarding stations, waiting for transportation to arrive, and travelling to a given destination (Frankena 1978). The level of service is sometimes also construed as a measure of the supply of public transportation. Peng et al. (1997) address the issue of simultaneity between service and ridership by estimating separate transit supply and demand equations.

Other determinants of transit ridership are beyond the control of administrators. For example, the level of transit patronage is highly correlated with urban area population (Taylor et al. 2009). Income is also likely to influence the level of ridership, and while some evidence suggests that public transportation is an inferior good (Nizlek and Duckstein 1974), others imply that it is a normal good (Chiang et al. 2011). Bresson et al. (2004) argue that ridership is positively correlated with income but negatively correlated with car ownership. Since income and car ownership are also correlated, omitting a measure of car ownership from a ridership equation tends to introduce downward bias on the income elasticity estimate.

Another variable included in many transit demand equations is a measure of the price of substitutes for public transportation. Motor vehicles are the main substitute for transit. Frankena (1978) finds that, of the various costs associated with owning a vehicle, only gasoline price has a statistically significant impact on transit ridership. In a study of seven U.S. cities, Wang and Skinner (1984) find that the elasticity of ridership with respect to gasoline prices ranges from 0.08 to 0.80. More recently, Lane (2010) confirms that gasoline price fluctuations can significantly impact the demand for public transportation.

There is little literature that directly addresses the impact on public transportation of proximity to an international border. Dascher and Haupt (2011) briefly discuss some attempts to extend public transit across the German-Polish border to facilitate cross-border shopping, among other things. While cross-border shopping is mutually beneficial for both the shoppers and the businesses they frequent, other groups may disapprove of public infrastructure projects aimed at facilitating trans-boundary commerce. The model explains the relative strength of these opposing interests based on factors such as inter-regional mobility, centralization of decision making, and patterns of property ownership. The model does not, however, attempt to determine the extent and nature of demand for cross-border public transportation.

While there is only limited academic work directly relating border region dynamics to patterns of transit ridership, there is a rich literature concerning trans-boundary traffic flows, especially along the U.S.–Mexico border. Much of it centers on pedestrian border crossings, and pedestrians are probably the group of entrants most likely to access public transportation in border cities. Survey evidence presented by Charney and Pavlakovich-Kochi (2002) indicates that 83% of pedestrians returning from trips to Arizona state that their primary purpose was shopping as compared with 68% of motor vehicle passengers. Similarly, Ghaddar and Brown (2005) find that 85% of cross-border visitors to Texas come for the purpose of shopping. Based on a survey conducted at the U.S.-Mexico border south of San Diego, Herzog (1991) finds that pedestrians comprise at least 50% of the Mexican nationals who cross the border to shop for food, clothing, or articles for the home. None of these surveys explicitly asks pedestrians whether or not they use transit on the U.S. side of the border.

Other analysts consider how economic conditions in Mexico influence the volume of pedestrian and vehicle traffic at ports of entry into the U.S. Fullerton (2000) finds that movements in the real peso-dollar exchange rate influence the volume of bridge traffic from Ciudad Juárez into El Paso. De Leon et al. (2009) conducted a similar study, but consider pedestrian traffic across the international bridges separately from motor vehicle traffic. Pedestrian crossings are positively related to Mexico's industrial production index and negatively related to bridge tolls. There is also a correlation between pedestrian crossings and the real exchange rate, but the direction of this relationship is uncertain. This ambiguity may arise because the data do not distinguish the nationality of border crossers. Thus, it is difficult to disentangle the effects of a shift in the exchange rate, which is likely to be different for Mexican and U.S. nationals.

Giermanski (1997) notes that cross-border shopping makes an important contribution to the economies of Texas border cities. Coronado and Phillips (2007) estimate that actual retail sales in Texas border cities exceed predicted sales by $1.9 billion. The surplus is attributed to Mexican nationals who cross the border into the U.S. to shop. Results of a mail survey conducted by Patrick and Renforth (1996) indicate that, between 1994 and 1995 when the Mexican peso lost more than half of its value relative to the dollar, retail sales in Texas border cities dropped by 42%. This supports the argument that cross-border shopping is an important feature of the border economy, at least in Texas.

While the literature on traffic across the U.S.-Mexico border does not specifically address public transportation, it provides insight into the motivations and economic impact of pedestrians who cross the border. One objective of this study is to quantify the link between pedestrian traffic through international ports of entry and public transportation networks along the Texas-Mexico border. Because Texas has several large border cities with well-developed public transportation systems, and receives 45% of all pedestrian entrants into the United States, the Texas-Mexico border region is well suited to a study of trans-boundary impacts on transit ridership (BTS 2011). In order to take advantage of potential border region covariances, the analysis is conducted using a seemingly unrelated regression estimator (Zellner 1962). Tests are also carried out for potential endogeneity among the variables utilized.

METHODOLOGY

Data

Laredo's public transportation system, El Metro, traces its origin to 1976. Its counterpart in Brownsville, the Brownsville Urban System, owes its existence to the city's acquisition of two privately-owned bus lines in 1978 (TDHPT 1979). Public transit in both cities is concentrated in fixed-route bus transportation, but demand response services are also provided (NTD 2011). The data on fare, service, and ridership in these cities are collected from Texas Transit Statistics (TTS), a report issued annually by the Texas Department of Transportation (TDT 1983-2009). Ridership, the dependent variable, is measured by the number of unlinked passenger trips per year and the level of service is measured by vehicle hours per year.

The calculation of fare is somewhat more complicated. While transit agencies often offer travel passes or fare discounts to particular groups of riders, the agencies that form the basis of this study do not provide historical information on each fare category. Therefore, the average fare is calculated by dividing annual farebox revenue by the number of passenger trips each year. The TTS figures for Laredo vehicle hours, fare revenue, and ridership in 2005 are exceptionally low; between 48% and 54% lower than the 2004 figures. Yet, data from the Laredo City Budget Department show that El Metro's operating expenditures and fare revenue steadily increased during that period. To resolve this contradiction, the service, fare and ridership data for 2005 are taken from the National Transit Database (NTD 2011), which provides figures similar to those of TTS in the other years for which the two datasets overlap.

Average fare data are adjusted for inflation using the Consumer Price Index (CPI, 1982-1984 = 100) from the website of the Federal Reserve Bank of St. Louis (FRB 2011). The real price of unleaded regular gasoline is obtained from the website of the Energy Information Administration (EIA 2011). Data on unemployment are from the website of the Bureau of Labor Statistics (BLS 2011) for years since 1990 and from the Texas Labor Market Review for prior years (TEC 1983-89). Population data are retrieved from the website of the Bureau of Economic Analysis (BEA 2011). The real peso-dollar exchange rate index is obtained from the Border Region Modeling Project at the University of Texas at El Paso (UTEP 2011). The number of pedestrian border crossings is provided to the authors by U.S. Customs and Border Protection.

The number of vehicle registrations is taken from the website of the Texas Department of Transportation (TDT 2011) from 1996-2007 and from the District and County Statistical Data reports (TDT 1987-95) and the Texas Almanac (Dallas Morning News 1984-86, 2008-09). The vehicle registration data are not available for Brownsville or Laredo for 1983 and 1994. Vehicle registration data are, however, available for El Paso, Texas, for both years through the Border Region Modeling Project (UTEP 2011) and the El Paso vehicle registration data are highly correlated with those for Brownsville and Laredo. The number of registered vehicles in El Paso and a time trend are used as regressor variables to allow estimating missing observations (Friedman 1962, Fernandez 1981).

Prior research indicates that unobserved social and economic factors may collectively exert long-term upward or downward pressure on transit patronage. The impact of such factors can be incorporated into a regression equation by means of a deterministic time trend (Wang and Skinner 1984, de Rus 1990, Romilly 2001, Lane 2010, Gkritza et al. 2011). Time is therefore included as an exogenous variable to capture any secular trend in ridership that is not explained by the other independent variables. A trend variable is not only of interest in its own right, but is useful in controlling for the effect of time on other explanatory variables in the equation. A regression equation that includes a trend variable yields the same parameter estimates that would be obtained from a regression in which all the variables are de-trended (Lovell 2008). De-trending is useful when the effect on ridership of short-term variations in an explanatory variable is obscured by the trend component of that variable. Excluding a time trend is sometimes found to substantially affect parameter estimates in transit demand equations (Dargay and Hanly 2002, Lane 2010).

Time series data, such as those analyzed here, are often non-stationary. Augmented Dickey-Fuller tests indicate that several of the variables are stationary in level form, while others are non-stationary. Faced with a similar situation, in which both stationary and non-stationary variables form part of a regression equation, Gkritza et al. (2011) utilize a linear time trend in conjunction with annual indicator variables to avoid spurious estimation results. As noted above, a time trend is also utilized in the present study.

Summary statistics are presented for all of the variables except for the time trend in Table 1 and 2. The sample period begins in 1983 and ends in 2009. The average ridership level in Laredo over the sample period is about 4.1 million trips per year compared with an average of 1.7 million trips

Table 1: Brownsville Summary Statistics

Variable	Mean	Standard Deviation	Minimum	Maximum	No.
Passenger Trips	1,657,263	126,191	1,333,719	1,909,292	27
Real Fare (cents)	34.938	3.955	29.016	48.495	27
Vehicle Hours	67,728	11,954	46,434	83,714	27
Real Gasoline Price (cents)	175.185	45.032	124	301	27
Unemployment	13,338	2,467	8,548	16,566	27
Registered Vehicles (000's)	178.233	42.097	118.561	261.453	27
MSA Population	310,827	51,285	238,878	396,371	27
Border Crossings (000's)	3,436.222	599.435	2,546.720	5,036.891	27
Real Exchange Rate Index	100.969	15.735	78.764	145.238	27

Table 2: Laredo Summary Statistics

Variable	Mean	Standard Deviation	Minimum	Maximum	No.
Passenger Trips	4,059,538	512,666	3,155,122	5,012,758	27
Real Fare (cents)	31.086	3.851	21.423	37.479	27
Vehicle Hours	135,843	39,465	88,830	186,304	27
Real Gasoline Price (cents)	175.185	45.032	124	301	27
Unemployment	6,604	1,635	4,251	11,010	27
Registered Vehicles (000's)	96.946	30.713	54.515	159.624	27
MSA Population	172,567	41,662	115,419	241,438	27
Border Crossings (000's)	4,405.490	759.044	3,112.505	6,674.293	27
Real Exchange Rate Index	100.969	15.735	78.764	145.238	27

in Brownsville. Laredo annually registers 4.4 million pedestrian border crossings on average, while 3.4 million visitors walked through the ports of entry into Brownsville each year. Thus, the ratio of pedestrian border crossings to transit passenger trips is roughly 2:1 in Brownsville but closer to 1:1 in Laredo.

Figures 1 through 4 depict the movement of the ridership series for Brownsville and Laredo in comparison to the movement of the real exchange rate index and pedestrian border crossing series. The data for Brownsville do not provide clear evidence of a positive correlation between ridership and pedestrian border crossings (Figure 1) but they do seem to indicate a negative relationship between ridership and the real exchange rate index (Figure 2). In Laredo, on the other hand, there appears to be some positive correlation between ridership and pedestrian border crossings for most of the period (Figure 3) but it is more difficult to discern a negative relationship between ridership and the real exchange rate index (Figure 4).

Figure 1: Brownsville Ridership and Pedestrian BorderCrossings

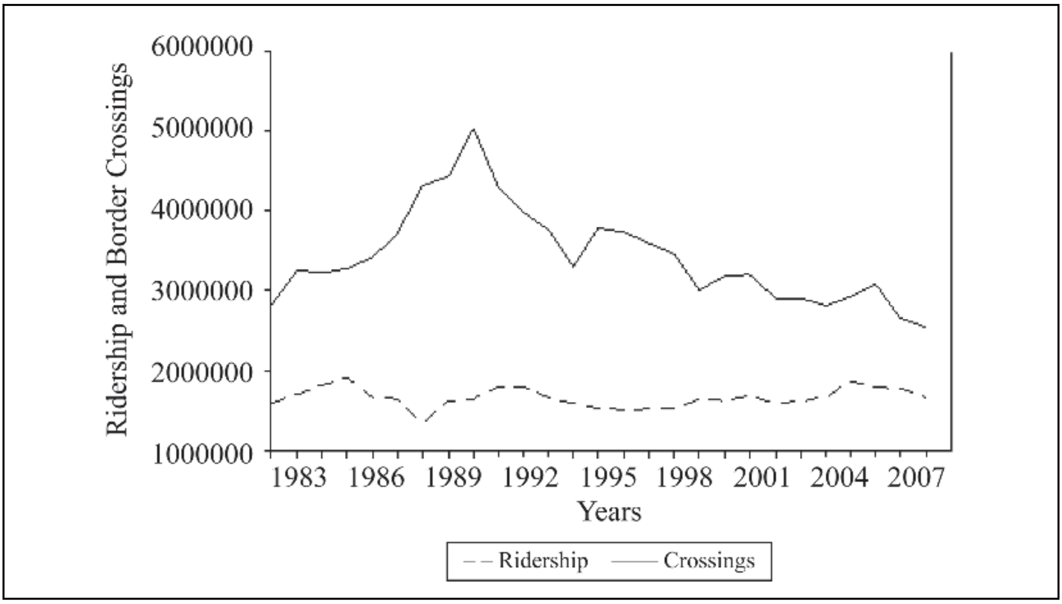


Figure 2: Brownsville Ridership and the Real Exchange Rate Index

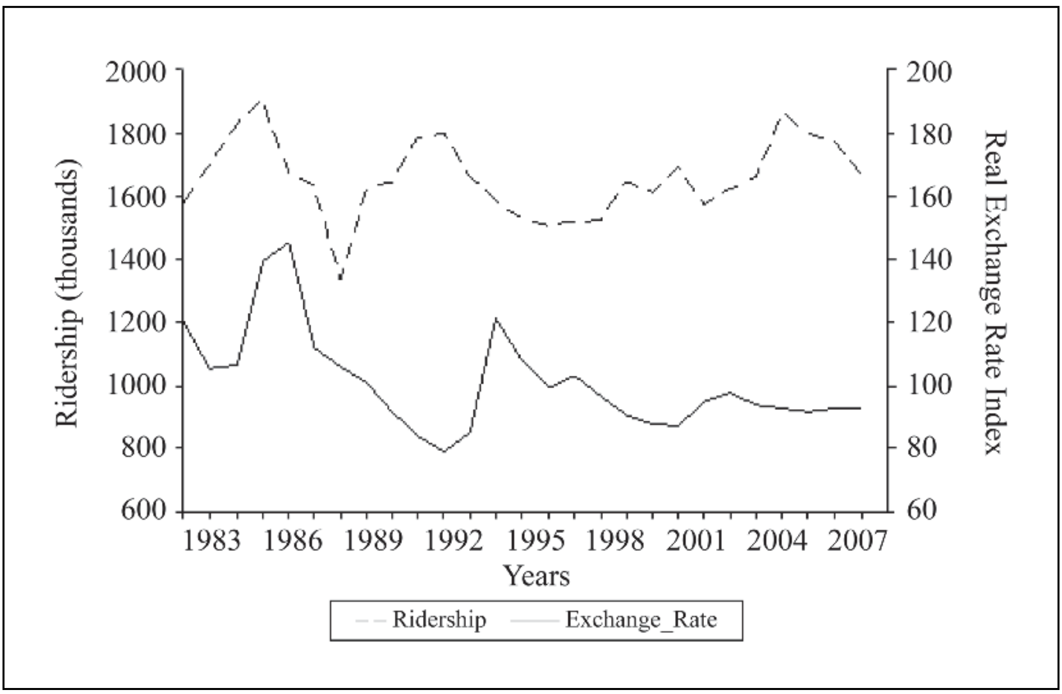


Figure 3: Laredo Ridership and Pedestrian Border Crossings

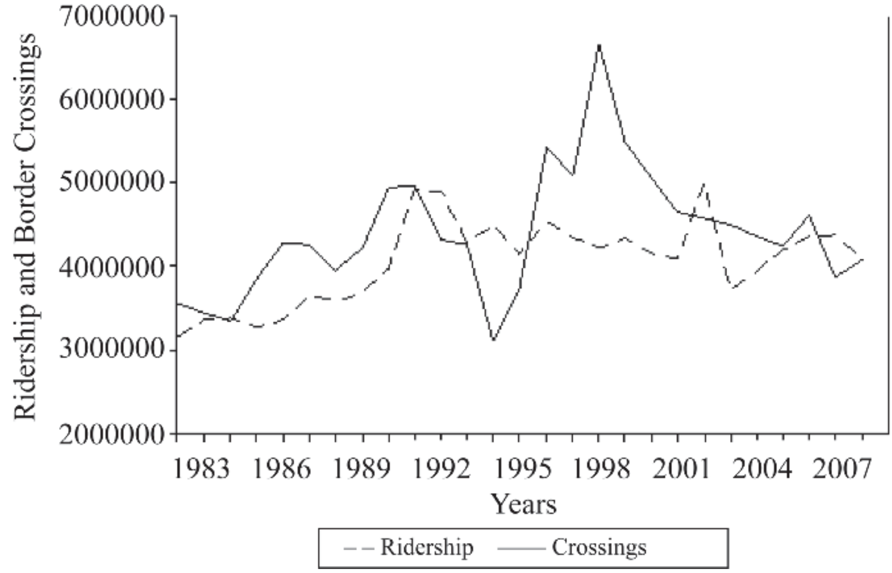
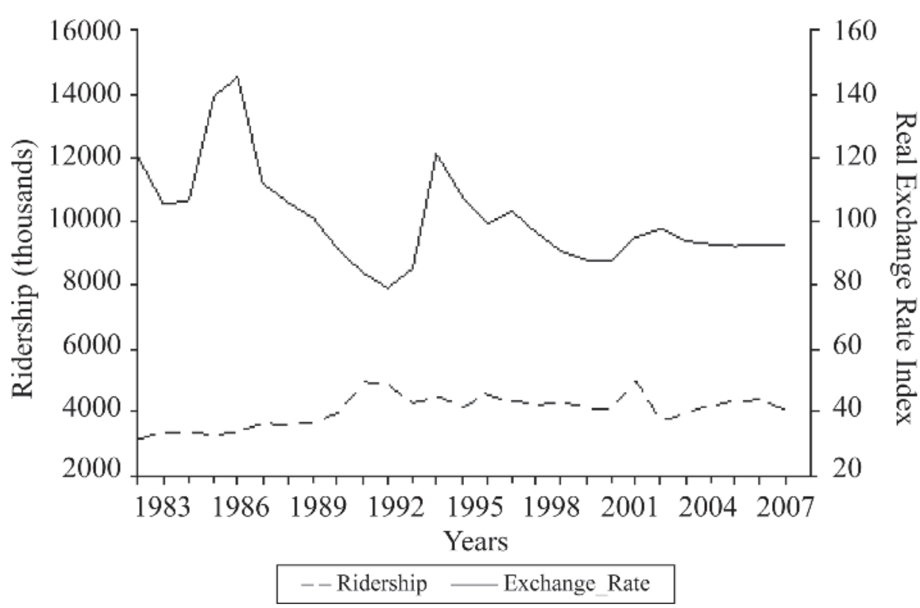


Figure 4: Laredo Ridership and the Real Exchange Rate Index



Analytical Approach

Although the magnitudes of the parameter estimates are likely to vary between Brownsville and Laredo, there is no á priori reason to believe that the signs of these coefficients will also vary. Thus, the following discussion applies equally to both cities. The basic model employed in the analysis of ridership can be expressed in implicit form as follows:

(1) $R_t = F(F_t, S_t, GP_t, U_t, V_t, P_t, T_t, PC_t, XR_t)$

where R_t represents transit ridership, F_t is the average fare, S_t is the level of service (vehicle hours), GP_t is the price of gasoline, U_t is the unemployment rate, V_t is vehicle registrations, P_t is population, T_t is a time trend, PC_t is pedestrian border crossings, and XR_t is the real exchange rate index. No hypothesis is advanced in the discussion that follows regarding the marginal effect of the time trend on ridership. That is because there is not enough information available to predict how multiple unobserved factors may collectively impact long term trends in ridership. The other independent variables are hypothesized to have the following marginal effects on ridership:

(2) $\frac{\partial R}{\partial F} < 0, \ \frac{\partial R}{\partial S} > 0, \ \frac{\partial R}{\partial GP} > 0, \ \frac{\partial R}{\partial U} > 0, \ \frac{\partial R}{\partial V} < 0, \ \frac{\partial R}{\partial P} > 0, \ \frac{\partial R}{\partial PC} > 0, \ \frac{\partial R}{\partial XR} < 0$

Fares are expected to be negatively related to ridership. An increase in the level of service, measured by vehicle hours, is anticipated to stimulate utilization of transit. Because travel by automobile is a substitute for travel by bus, higher gasoline prices are also expected to increase transit patronage. A higher level of unemployment, which is usually associated with lower income levels, is predicted to increase ridership. Car ownership, represented by vehicle registrations, is predicted to vary inversely with transit patronage. Holding other factors constant, a larger population is anticipated to expand transit usage. Pedestrian border crossings are hypothesized to be positively related to ridership. Finally, an increase in the value of the Mexican peso relative to the U.S. dollar is hypothesized to increase cross-border shopping and transit usage by Mexican nationals. An increase in the value of the peso relative to the dollar is reflected in a decrease of the real exchange rate index because the latter is defined as pesos per dollar in inflation-adjusted terms. Therefore, a negative relation is expected to exist between the real exchange rate index and ridership.

Because the number of vehicle hours per year can be construed as a measure of the supply of public transportation, the presence of this variable in a public transportation demand equation raises the prospect of simultaneity. Average fare may also be endogenous because it is calculated as the ratio of total fare revenue to total ridership. Changes in this ratio may reflect changing patterns of ridership rather than the actual changes in ticket prices (de Rus 1990). Alternatively, the average ticket price may depend, to some extent, on the total volume of ridership. To determine whether simultaneity exists in either average fares or vehicle hours, artificial regression tests are conducted for the Brownsville and Laredo ridership equations (Davidson and MacKinnon 1989).

Next, separate regression equations are estimated for both cities by the method of ordinary least squares. If there is evidence that the disturbance terms in the regression equations for each city are correlated, more efficient coefficient estimates may be obtained using the seemingly unrelated regression parameter estimation method. Zellner (1962) shows how knowledge of the covariance between error terms in separate equations can be incorporated into parameter estimates through a generalized least squares (GLS) procedure. For a system of seemingly unrelated equations, the GLS estimator in matrix form is as follows:

(3) $\hat{\beta}_{GLS} = (X' H' HX)^{-1} X' H' Hy = (X' \Sigma^{-1} X)^{-1} X' \Sigma^{-1} y$

If the covariance between the disturbances of equations i and j in identical time periods (the contemporaneous error covariance) is represented by σ_{ij}, then Σ is equivalent to the following when there are only two equations:

$$(4)\quad \Sigma = \begin{bmatrix} \sigma_{11}\,I & \sigma_{12}\,I \\ \sigma_{21}\,I & \sigma_{22}\,I \end{bmatrix}$$

In practice, the error variances and cross-equation covariances are unknown and must, therefore, be estimated. If the error terms exhibit substantial contemporaneous covariance across equations, the seemingly unrelated regression procedure will be used.

The seemingly unrelated regression model as presented does not allow for serially correlated error terms (Zellner 1962). To correct for first to 12th-order serial correlation, Gkritza et al. (2011) incorporate autoregressive parameters into a seemingly unrelated regressions model. In the current case, where there are only 27 observations and nine explanatory variables for each equation, correcting for higher order serial correlation in this manner would result in an unacceptably small number of degrees of freedom. To determine whether the error terms are autocorrelated, Q-statistics are calculated for the regression residuals and these are compared against critical values of the Chi-square distribution (Pindyck and Rubinfeld 1998).

There are various reasons why ridership in Brownsville and Laredo may be affected by the same unobserved factors. Both cities are located in Texas and their transit systems are thus affected by similar state regulations and funding procedures. They are affected by many of the same climatic conditions. Both cities also share a similar demographic profile that includes a larger share of foreign-born individuals than the average for Texas or the United States (U.S. Census Bureau 2011). Some research shows that recent immigrants are more likely than others to use mass transit (Heisz and Schellenberg 2004). While those factors may be individually insignificant, taken as a whole they may potentially generate substantial error covariance between the Brownsville and Laredo equations.

EMPIRICAL ANALYSIS

To determine whether average fare and vehicle hours are endogenous, they are each regressed on instrumental variables as well as the exogenous variables in the demand equation (Davidson and MacKinnon 1989). For Brownsville, real earnings per employee in the local transportation sector are used as an instrumental variable for vehicle hours. The instruments for fare are the real composite price of fossil fuels and the net public operating cost of transit. Bus driver wages and fuel prices are important costs that transit operators face (Frankena 1978), while net public operating costs reflect governmental allocation of funds to transit operations. Public funding of transit and input prices both influence the level of service provided and the setting of fares but they are not likely to be affected by ridership levels. The residuals from these regressions are statistically insignificant when included as independent variables in the ridership equation. Thus, it is not possible to reject the null hypothesis that both variables are exogenous.

The procedure is repeated for the Laredo variables. The instruments chosen for fare are the number of cooling degree days per year and a one-year lag of cooling degree days. The number of cooling degree days per year affects fares indirectly by influencing costs related to air conditioning, breakdowns, equipment malfunctions and the like. The instrument for vehicle hours is the total number of state and local employees in Laredo. As mentioned above, labor represents an important cost for transit providers, and this cost affects decisions to expand or contract service. Neither temperature patterns nor government employment levels are likely to be correlated with the error term in a transit demand equation. As with the Brownsville variables, the null hypothesis of exogeneity cannot be rejected.

The residual correlation coefficient for the Laredo and Brownsville equations is 0.650. Given substantial error correlation between equations, a seemingly unrelated regression method is used to produce the parameter estimates shown in Tables 3 and 4.The variables are mean-centered and, with the exception of the time trend in the Brownsville equation, all variables are also logarithmically transformed prior to estimation. The coefficients are very similar to the elasticities derived from regressions performed on the variables in level form. The variables included in the regression equation are all statistically significant for at least one of the two cities. Although more parameter estimates are statistically significant in the Brownsville regression output, the coefficient of determination is much higher in that for Laredo. As discussed below, this may be due to multicollinearity among several of the explanatory variables.

The elasticities of demand with respect to fare are estimated to be –0.45 for Brownsville and –0.41 for Laredo. These figures are very similar to the short-run elasticity estimates reported in a national study by Pham and Linsalata (1991) and in an analysis of multiple elasticity studies by Paulley et al. (2006). Both of these estimates suggest that ridership levels in these cities are considerably more responsive to changes in fares than what is implied by the traditional rule-of-thumb elasticity of –0.33 (Cervero 1990).

While the estimates presented here suggest that demand does respond to fare changes, it is also somewhat inelastic with respect to these changes. This implies that fare increases could generate additional revenues that might serve to either increase the level of service or reduce the level of public transit subsidies. However, it is important to remember that the estimates are short-term elasticities. While increasing fares might generate additional revenues in the short term, Paulley et al. (2006) caution that long-term bus fare elasticities may be close to unity. If this is the case, increasing fares may not lead to substantial revenue increases in the long run.

Table 3: Brownsville Ridership

Sample Period: 1983 to 2009

Dependent Variable: Ridership $_t$ (total number of unlinked passenger trips per year)

Variable	Coefficient	Std. Error	t-Statistic	Probability
Fare $_t$	-0.453189	0.114440	-3.960054	0.0004
Vehicle Hours $_t$	0.602066	0.125445	4.799439	0.0000
Gasoline Price $_t$	0.365247	0.078142	4.674150	0.0000
Unemployment $_t$	0.255233	0.078078	3.268956	0.0025
Vehicle Registrations $_t$	-0.603354	0.260029	-2.320335	0.0265
Population $_t$	4.382249	1.266791	3.459330	0.0015
Time Trend $_t$	-0.074575	0.025177	-2.961992	0.0055
Pedestrian Crossings $_t$	0.649275	0.180369	3.599701	0.0010
Exchange Rate $_t$	0.042229	0.102897	0.410399	0.6841
R-Squared	0.566830	Mean Dependent Variable		0
Adjusted R-Squared	0.337505	S.D. Dependent Variable		0.077137
Standard Error of Regression	0.062785	Durbin-Watson Statistic		2.054868
Sum of Squared Residuals	0.067013			

The elasticity of ridership with respect to transit service is estimated to be 0.60 in Brownsville. More surprisingly, the service coefficient for Laredo is negative and insignificant. The negative coefficient persists regardless of whether vehicle miles or vehicle hours are used to measure the level of service. While it seems counterintuitive that an increase in transit service would not result in increased ridership, Small (1997) notes that expanding transit service into suburban areas often does not yield substantial increases in ridership.

Consistent with the hypothesis that travel by bus and travel by automobile are substitutes, a 10% increase in the price of gasoline is found to increase transit usage by 3.7% in Brownsville and 2.0% in Laredo. These results, like those obtained by Wang and Skinner (1984), indicate that a change in gasoline price has a significant positive effect on ridership, but that this relationship is inelastic. If automobiles are substitutes for transit, it is not surprising to find that the number of registered vehicles is negatively correlated with ridership in Brownsville, although the relationship is less statistically significant in Laredo. Brownsville's elasticity of demand with respect to car-ownership is –0.60, which is smaller than the composite elasticity estimate of –1.48 reported by Holmgren (2007).

A 10% increase in unemployment is expected to cause a 2.6% increase in ridership in both Brownsville and a 1.5% increase in Laredo. This positive relationship is consistent with evidence in Nizlek and Duckstein (1974) that ridership moves in tandem with short-term cycles of unemployment. Additional regressions are performed to examine whether changes in personal income affect ridership levels in Brownsville and Laredo. Because income does not appear to explain much of the variation in ridership, it is excluded from the final specification.

Although transit ridership in Brownsville exhibits a secular downward trend, population growth in that city is associated with increased ridership. The population coefficient for Laredo is

Table 4: Laredo Ridership

Sample Period: 1983 to 2009

Dependent Variable: Ridership_t (total number of unlinked passenger trips per year)

Variable	Coefficient	Std. Error	t-Statistic	Probability
Fare_t	-0.411995	0.077357	-5.325866	0.0000
Vehicle Hours_t	-0.160948	0.202795	-0.793649	0.4329
Gasoline Price_t	0.198367	0.109467	1.812118	0.0788
Unemployment_t	0.145462	0.070783	2.055032	0.0476
$\text{Vehicle Registrations}_t$	-0.364547	0.205281	-1.775849	0.0847
Population_t	0.416503	0.489169	0.851451	0.4005
Time Trend_t	0.113669	0.064900	1.751448	0.0889
$\text{Pedestrian Crossings}_t$	0.164885	0.070609	2.335175	0.0256
Exchange Rate_t	-0.328210	0.081748	-4.014872	0.0003
R-Squared	0.892667	Mean Dependent Variable		0
Adjusted R-Squared	0.835843	S.D. Dependent Variable		0.128211
Standard Error of Regression	0.051946	Durbin-Watson Statistic		2.020353
Sum of Squared Residuals	0.045873			

not statistically significant, perhaps due to multicollinearity. The correlation coefficient between population and the time trend is 0.996, and the coefficient on population is statistically significant at the 10% level when time is omitted from the specification. Multicollinearity may also exist between these two variables and vehicle registrations because the correlation coefficient is 0.964 in both cases. This multicollinearity may explain why the coefficient of determination is relatively high for Laredo despite the apparent statistical insignificance of several explanatory variables.

A 10% increase in pedestrian border crossings is expected to increase transit ridership by 6.5% in Brownsville and by 1.6% in Laredo. Although the transit systems in both cities are accessible to pedestrian border crossers, on average there are twice as many such pedestrians per transit trip in Brownsville as in Laredo. The larger flow of cross-border pedestrian traffic relative to transit ridership partially accounts for the much larger elasticity of demand with respect to border crossings in Brownsville. One reason that more pedestrians cross into Brownsville than Laredo is that Matamoros, the Mexican city south of Brownsville, is much larger than Nuevo Laredo, the city situated opposite Laredo.

Border region transit operators could potentially capitalize on the positive relationship between border crossings and transit ridership by facilitating access to transit by visitors crossing through the ports of entry. For example, it may be possible to attract additional pedestrian border crossers by increasing the density of routes or the frequency of service near ports of entry. Bus terminals and other transit nodes can be located near border crossings to facilitate speedy access to a broader array of potential destinations within the city, including shopping centers. In order to attract additional passengers from Mexico, border area bus operators could follow the examples of numerous regional businesses and allow fares and passes to be purchased using pesos (Yoskowitz and Pisani 2007, Muñoz et al. 2011). Lastly, transit operators should take cross-border economic conditions into account when setting the levels of fare and service.

The exchange rate does not register a statistically significant impact on ridership in Brownsville. In Laredo, on the other hand, a strong negative relationship exists between the exchange rate and transit ridership. This shows that transit patronage in Laredo declines when the peso is weak relative to the dollar. The most likely explanation for this phenomenon is that Mexican nationals respond to an increase in the exchange rate by curbing cross-border shopping excursions and, therefore, reduce their utilization of Laredo's transit system. This is consistent with previous findings that shopping motivates the largest number of border crossings by pedestrians and others on the United States–Mexico border (Charney and Pavlakovich-Kochi 2002, Ghaddar and Brown 2005).

The negative relationship between the real exchange rate index and Laredo transit ridership raises the question of whether an improvement of economic conditions in Mexico is associated with more transit trips in Laredo. Although a weak peso may tend to increase manufacturing employment in Nuevo Laredo (Cañas et al. 2007), it also tends to reduce dollar-denominated wages in Mexico, and sharp peso depreciations often correspond to economic downturns. If residents of Mexico make fewer transit trips in Laredo during economic downturns and increase their utilization of the transit system during economic recoveries, this implies that transit is a normal good for those cross-border visitors. In contrast, the coefficient on the unemployment variable implies that transit is an inferior good for residents of Laredo. Border region transit authorities should consider the possibility that local and cross-border transit users may respond differently to cyclical changes in local economic performance.

The validity of the foregoing interpretation of the parameter estimates depends on the assumption that the residuals are not auto-correlated. Although the respective Durbin-Watson statistics are 2.02 and 2.05, the sample size is too small to definitively accept or reject the null hypothesis of no first-order serial correlation. Residual Q-statistics are calculated for up to 12 lags using both Box-Pierce and Ljung-Box methodologies (Lütkepohl and Krätzig 2004). The null hypothesis that the residuals are not autocorrelated up to lag 12 cannot be rejected at the commonly acceptable significance level of 0.05, implying that the error terms are not serially correlated.

CONCLUSION

The demand for public transportation in Brownsville and Laredo follows many, though not all, of the patterns observed in other regions. For example, ridership is inelastic with respect to both the level of service and fares, as in many other areas. Although some empirical evidence indicates that proportional increases in the level of service and fares will substantially increase ridership, this does not seem to hold for Laredo or Brownsville. Increasing fares and service proportionally appear to have small net effects on Brownsville ridership and a negative effect on Laredo ridership. The regression results also indicate that, as in other regions, public transportation and travel by private automobile are substitutes. Increases in gasoline prices and decreases in vehicle registrations generate positive, but less-than-proportional, changes in transit ridership in Brownsville and Laredo.

After controlling for other factors affecting transit demand, cross-border pedestrian traffic has a positive effect on transit ridership in Brownsville and Laredo. This relationship suggests that public transportation provides enhanced mobility to pedestrian visitors from Mexico, many of whom cross the border for the purpose of shopping. Similar to the retail sector, Laredo transit patronage increases when the peso appreciates and decreases when the purchasing power of the peso falls. In combination, these results suggest that public transit helps connect pedestrian border crossers with destinations such as retail outlets located on the north side of the international boundary.

While existing studies describe the role of public transportation in providing access to jobs for people who do not own an automobile, there has been little research into the role of transit in facilitating cross-border shopping. Additional research is needed to establish whether transit can directly benefit border city retail sectors by transporting foreign shoppers. It is also important to know whether the inspection process at ports of entry has a direct bearing on modal choice. There are times when vehicle wait times are longer than pedestrian wait times at Texas ports of entry. In response, some border crossers opt to walk across the border. Public transit routes and schedules can be modified to reflect this general pattern.

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