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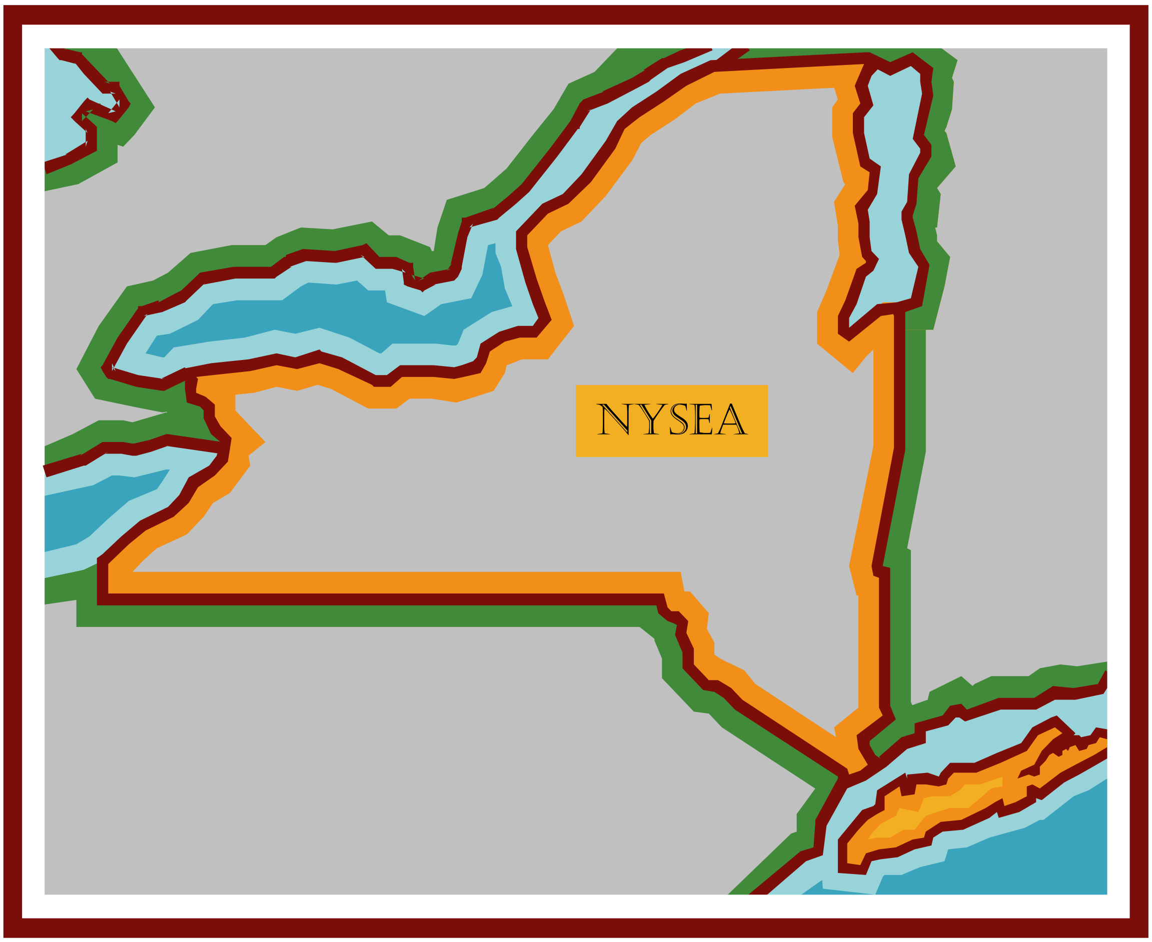

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CONTENTS

EDITORIAL

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The Impact of WVU Football and Basketball on Hotel Demand

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ABSTRACT

This paper uses daily hotel occupancy data for the Greater Morgantown area to estimate the effect of West Virginia University football and basketball games on hotel demand. Hotel demand is an important part of the economic activity generated by sporting events because hotel rooms are largely occupied by out-of-town guests. Their expenditures, therefore, are likely to represent new local economic activity. We look at the effect of Mountaineer football and basketball games on average daily room rates, revenue per room, demand, occupancy, and total revenue. We find large and statistically significant effects of Mountaineer football on hotel demand with very little evidence of crowding out. Our estimates for basketball, while statistically significant in a positive direction, are considerably smaller. WVU football games bring in approximately $360,000 in additional hotel revenue from each football home game, while WVU basketball games generate about $20,000 in additional hotel revenue.

INTRODUCTION

For some communities, tourism is the fuel the drives the local economy. For others, tourism plays a small, but economically meaningful role in their economy through increased employment and tax revenue due to the spending of out-of-town visitors. With the largest university in the state of West Virginia located in the greater Morgantown area (population approximately 140,000), it plays host to regular events such as motorcycle rallies, challenging marathons, visits from prospective parents and students, and of course the West Virginia University Mountaineers. The college town is visibly packed with economic activity on days where the 60,000 capacity Milan Puskar Stadium hosts a WVU football game.

While it may appear obvious that these mega sporting events (for a college town) bring many visitors from outside the greater Morgantown area; measuring the net impact of WVU football and basketball games can prove to be difficult due to offsetting behavior on behalf of local residents or tourists uninterested in major college athletics. For example, while spending on sports might increase due to a major athletic event that might largely reflect substitution from consumer spending elsewhere in economy (Coates and Humphreys, 2002). Similarly, convention organizers know that average daily room rates at hotels will be higher during major sporting events. Therefore, they may move conventions that attract out-of-town guests to weekends with away games, or to the spring. This redistribution of economic activity to other dates makes it difficult to estimate the economic impact of athletic events and is one reason why many *ex ante* analyses tend to overestimate the magnitude of the positive impacts compared to *ex post* analyses of the same event (Matheson, 2002).

This paper tests for economic impacts associated with WVU football and basketball games by examining daily hotel occupancy data on the Morgantown metropolitan area, provided by STR, a firm that specializes in metropolitan lodging data worldwide. In doing so, we build off a growing literature that uses changes in hotel demand to estimate part of the economic impact of sports (Depken and Stephenson, 2018), same- sex marriage legalization (Earhart and Stephenson, 2018), and political conventions (Heller et al., 2018). The current paper contributes to the existing literature on the economic impact of college sports on local economies. Baade et al. (2008) look at 63 metropolitan areas that host major college football from 1970 to 2004 and find no positive impact on employment or income levels. Most relevant to the current paper, they examine 42 small college towns and find football success *reduces* the growth rate of per capita personal income. Baade et al. (2011) investigate the effect of home football and basketball games at the University of Florida and Florida State University, finding that basketball has zero impact on taxable sales, while football increases taxable sales by $2 million per home game. Using monthly sales revenue, Coates and Depken (2009) find that, on average, increases in sales tax revenue due to the game is exactly offset by a reduction in local spending in other areas. Coates and Depken (2011) find, however, that a full season of major college football has the same impact on local sales tax revenue as a hosting the Super Bowl. Lentz and Laband (2009) conduct an MSA level analysis of college athletics and find that there exists a positive relationship between athletics revenues and employment in accommodations and food services industries.

Our analysis provides an estimate of the economic impact of WVU football and basketball on the Morgantown economy, which may be of interest to policymakers. A 2012 study conducted by WVU's Bureau of Business and Economics Research (BBER) found that each home football game generated approximately $1.6 million in revenue, creating an impact of over $11 million over the course of a season (Christiadi, 2012). The study employs a survey method in which local lodging, drinking, and eating establishment self-report their data for game days and non-game days to gauge the net impact. While our methodology cannot estimate the local impacts from dining and drinking, we hope to more precisely estimate the net effects of WVU football and basketball from out-of-town visitors using hotel demand. Given that very few Morgantown residents use hotels for home games, increases in room rates and occupancy rates related to these events provide a good estimate of out-of-area visitors. To preview our results, we find statistically significant evidence that WVU football and basketball home games increase hotel occupancy and revenue. Only football games, however, have an economically meaningful impact.

We proceed as follows. Section 2 discusses our daily hotel data and our empirical approach. In Section 3 we present our primary results, while Section 4 focuses on robustness checks. Section 5 concludes.

DATA AND EMPIRICAL APPROACH

This study uses nightly hotel data from Greater Morgantown for the period January 1, 2005 to December 31, 2017. The data were obtained from STR, a firm that compiles hotel occupancy data from the U.S. and other countries. This detailed data is then matched with WVU's football and basketball schedule. Our variables of interest are occupancy, average daily room rate (ADR), demand, revenue per available room

(RevPar), and total revenue. Occupancy represents the percentage of available rooms that are occupied, ADR is calculated by dividing the total revenue by the number rooms sold, and revenue per available room is found by dividing the total revenue by the number of available rooms. Summary statistics for these variables over the full period can be found in Table 1.

Table 1: Summary Statistics, Full Sample

Statistic	Mean	St. Dev.	Min	Max
Occupancy (%)	61.23	15.87	16.29	94.90
Average Daily Room Rate	80.55	11.29	56.44	130.64
Rev. Per Room	50.10	17.25	11.33	115.10
Supply	12,142.62	993.05	10,727.00	14,241.00
Demand	7,435.82	2,009.83	1,855.34	12,202.51
Revenue ($1,000’s)	612.80	227.79	121.58	1,535.55

N= 4,748.

The summary statistics tell us that on average about 61% of the rooms are full, and that the prices are typically around $80 per room per night, creating overall revenue averaging $612,000 (with a maximum in our sample of over $1.5 million). While these statistics are for the entire sample, Table 2 shows the summary statistics by days with a football game, basketball game, and no game. Importantly, these are all days of a sporting event, and thus include away games for each respective sport. From this table, it can be seen that the means for football are higher in every statistic used when compared to both basketball and days when there is no game. Basketball, on the other hand, appears to have lower amounts for each hotel statistic, although this is consistent with overall lower hotel demand during the winter months of basketball season.

To further illustrate the different behavior of the two sports relative to days when no game exists, Figures 1 and 2 track the revenue and average daily room rate over the period for football and basketball, respectively and compare home games with away games and no games. Here the magnitude of the difference in means for football games versus non-football games is clear. These differences in means suggest that home football games have a large impact on revenue, RevPar, ADR, occupancy, and demand, while basketball may have no effect or even, perhaps, a negative effect. In Section 3, we further explore the data using regression analysis to account for other possible explanations for these differences such as day of the week and seasonal effects.

Table 2: Summary Statistics by Sport

Statistic	Mean	St. Dev.	Min	Max
Football (N = 167)				
Occupancy (%)	68.80	15.23	20.22	93.58
Average Daily Rate	95.28	17.10	59.62	130.64
Rev. Per Room	67.19	23.18	13.70	115.10
Supply	12,243.54	1,000.30	10,840	14,092
Demand	8,423.26	1,963.19	2,249.04	11,979.42
Revenue ($1,000's)	827.58	305.41	152.33	1,522.41
Basketball (N = 456)				
Occupancy (%)	51.42	12.33	19.77	83.17
Average Daily Rate	77.98	10.54	57.33	127.91
Rev. Per Room	40.57	13.09	11.33	100.16
Supply	12,091.65	1,005.06	10,727	14,045
Demand	6,221.16	1,579.93	2,120.68	10,551.39
Revenue ($1,000's)	494.55	175.42	121.58	1,259.93
No Game (N = 4142)				
Occupancy (%)	61.97	15.82	16.29	94.90
Average Daily Rate	80.29	10.69	56.44	125.85
Rev. Per Room	50.46	16.76	11.83	113.23
Supply	12,144.63	991.36	10,727	14,241
Demand	7,526.05	2,002.08	1,855.34	12,202.51
Revenue ($1,000's)	617.25	221.94	131.36	1,535.55

Figure 1: Football ADR and Revenue

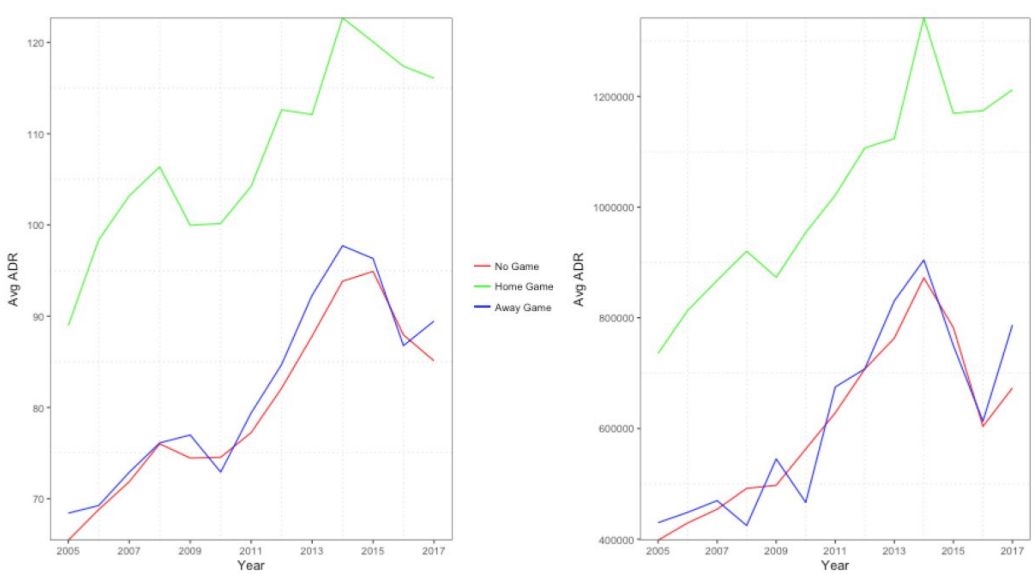
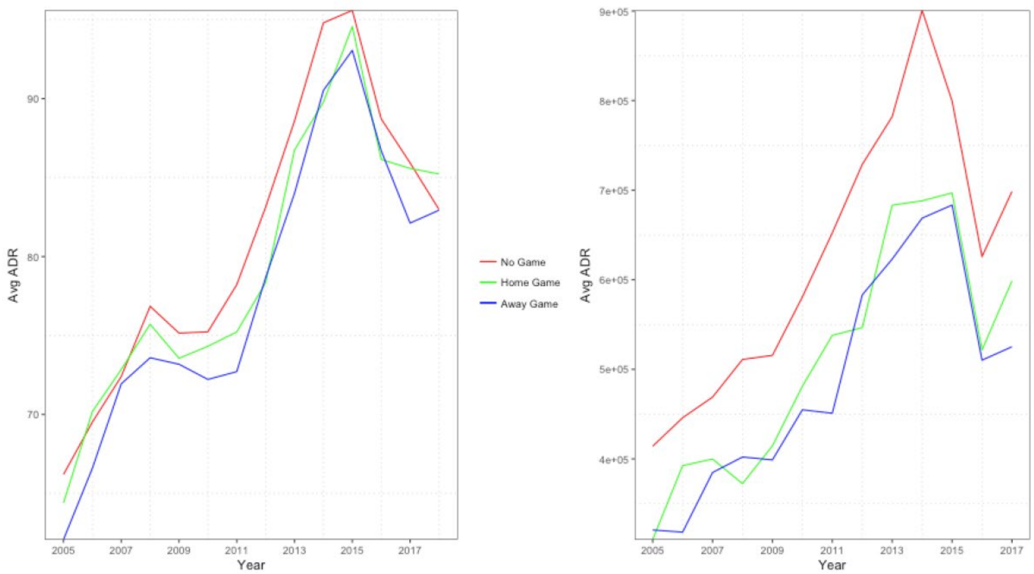


Figure 2: Basketball ADR and Revenue



EMPIRICAL RESULTS

Table 3: Football Regressions

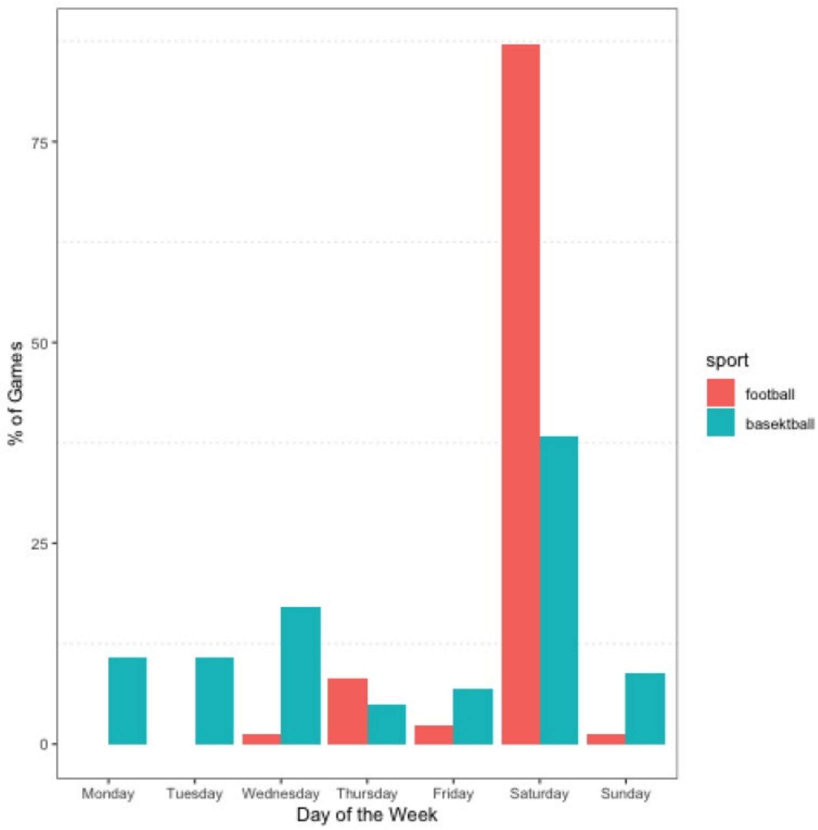
Dep. Variable	ADR		RevPar		Demand		Occupancy		Revenue	
Mean	80.54		50.09		7,435.81		61.22		612.8	
Football - 4	-0.117		0.577		75.831		0.381		9.779	
	(0.408)		(0.811)		(96.171)		(0.804)		(10.055)	
Football - 3	0.59		1.103		88.234		0.484		16.451	
	(0.411)		(0.818)		(96.975)		(0.811)		(10.139)	
Football - 2	1.439	***	1.326		52.648		0.288		18.15	*
	(0.415)		(0.824)		(97.729)		(0.817)		(10.218)	
Football -1	21.159	***	22.575	***	1023.701	***	8.165	***	280.781	***
	(0.416)		(0.827)		(98.095)		(0.820)		(10.256)	
Home Football	24.816	***	29.220	***	1499.803	***	12.121	***	360.789	***
	(0.423)		(0.841)		(99.734)		(0.834)		(10.427)	
Football +1	1.617	***	0.567		-43.529		-0.128		5.235	
	(0.389)		(0.773)		(91.714)		(0.767)		(9.589)	
Football +2	0.453		-0.342		-119.983		-0.981		-3.737	
	(0.389)		(0.773)		(91.648)		(0.766)		(9.582)	
Football +3	0.370		1.160		111.516		0.841		15.363	
	(0.382)		(0.760)		(90.096)		(0.753)		(9.420)	
Football +4	-0.036		1.03		135.189		1.092		13.19	
	(0.381)		(0.756)		(89.716)		(0.750)		(9.380)	
Day of Week FE	✓		✓		✓		✓		✓	
Week FE	✓		✓		✓		✓		✓	
Month FE	✓		✓		✓		✓		✓	
Year FE	✓		✓		✓		✓		✓	
R-Squared	0.943		0.904		0.901		0.888		0.915	

Note: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Intercept included but not reported. N = 4,748 for all models. The dependent variable is noted across the top of each model. The full sample mean of each dependent variable is provided above the regression results for reference.

The effect of a home football game is significant across all dependent variables, as is the day before a home football game. This makes sense as many football games have a start time of noon on Saturday, suggesting we should expect an increase on the day before as visitors come in. The results are quite stark, with over a 50% increase in the revenue and revenue per available room. Average daily room rates increase as well by almost 30% while occupancy increases to about 75%.

Basketball tells a slightly different story. In addition to occurring during the winter months, these games are also more frequently scheduled on weekdays (see Figure 3). However, the inclusion of the fixed effects should account for any day of the week or seasonal differences in any of the dependent variables. The results for the regressions on Basketball can be seen in Table 4.

Figure 3: Percentage of Games by Day of the Week



These results are much less clear compared to football. With basketball, many variables for the leads and lags are also significant, most likely due to basketball games occurring more than once a week. The magnitude of the results is much smaller as well, with the revenue and revenue per available room increasing slightly, but not any amount that is economically significant. For comparison purposes, football had an impact of about $360,000 in additional revenue on game days, while basketball only brings in around $20,000. Even if the coefficients for all the leads and lags that are significant are taken into consideration, it is still only about one-third of the impact from football. Occupancy also sees little change. As mentioned previously, events may have the adverse effect of crowding out other visitors who would be coming to the area for other events. This most likely occurs because Morgantown is a destination city for little else than events hosted by the University, with only a few other events occurring throughout any given year. While it is clear that basketball would not have a crowding out effect due to the small impact of games on occupancy, football's impact is much greater. Even with this larger effect, occupancy remains around 75% on average, though there may be certain games, namely games against rivals or highly ranked opponents, where this crowding out effect may occur.

Table 4: Basketball Regressions

Dep. Variable	ADR		RevPar		Demand		Occupancy		Revenue	
Mean	80.54		50.09		7,435.81		61.22		612.8	
Basketball - 4	0.136		0.099		-7.079		-0.049		-1.622	
	(0.413)		(0.628)		(63.277)		(0.527)		(7.769)	
Basketball - 3	0.518		1.790	***	223.898	***	1.792	***	22.276	***
	(0.444)		(0.675)		(67.962)		(0.567)		(8.350)	
Basketball - 2	0.433		1.613	**	208.324	***	1.705	***	20.09	**
	(0.447)		(0.679)		(68.411)		(0.571)		(8.406)	
Basketball -1	-0.338		0.583		162.659	**	1.363	**	7.062	
	(0.434)		(0.660)		(66.436)		(0.554)		(8.163)	
Home Basketball	0.588		1.614	**	228.084	***	1.844	***	19.846	**
	(0.433)		(0.658)		(66.231)		(0.553)		(8.138)	
Basketball +1	1.466	***	2.610	***	207.997	***	1.706	***	31.823	***
	(0.438)		(0.666)		(67.106)		(0.560)		(8.245)	
Basketball +2	1.138	**	1.298	*	62.782		0.634		14.509	*
	(0.446)		(0.678)		(68.312)		(0.570)		(8.393)	
Basketball +3	0.902	**	0.863		31.873		0.360		9.453	
	(0.439)		(0.667)		(67.202)		(0.561)		(8.257)	
Basketball +4	1.135	***	1.672	***	108.446	*	0.877	*	20.839	***
	(0.417)		(0.634)		(63.842)		(0.533)		(7.844)	
Day of Week FE	✓		✓		✓		✓		✓	
Week FE	✓		✓		✓		✓		✓	
Month FE	✓		✓		✓		✓		✓	
Year FE	✓		✓		✓		✓		✓	
R-Squared	0.855		0.856		0.892		0.88		0.874	

Note: *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$. Intercept included but not reported. N = 4,748 for all models. The dependent variable is noted across the top of each model. The full sample mean of each dependent variable is provided above the regression results for reference.

The results presented in this section suggest that while including a number of fixed effects, sports games in the Morgantown area have a positive impact on hotel revenues and occupancy, suggesting that there is an increase in visitors from outside of the area. These results lend credence to the idea that these events are having the expected positive effect. However, to test if this is due to any sort of random chance, placebo regressions are run and included in the following section. Additionally, the impact of changing conferences is evaluated.

ROBUSTNESS CHECKS

To test the validity of the results, a placebo test was run using 86 randomly selected days to simulate home football games, 199 days for home basketball games, and 285 for the full sample of home football and basketball games. These results can be seen in Tables 5 and 6. The results show no significance on the random selection of days for both groups and across all variables. This would suggest that the positive effects occurring in tandem with home football and basketball games are not due to some random chance.

Table 5: Football Placebo

Dep. Variable	ADR	RevPar	Demand	Occupancy	Revenue
Random Days	0.32	-0.37	-67.3	-0.45	-6.49
	(0.55)	(0.84)	(84.70)	(0.71)	(10.40)
Day of Week FE	✓	✓	✓	✓	✓
Week FE	✓	✓	✓	✓	✓
Month FE	✓	✓	✓	✓	✓
Year FE	✓	✓	✓	✓	✓
R-Squared	0.85	0.86	0.89	0.88	0.87

Note: ∗∗∗ $p < 0.01$, ∗∗$p < 0.05$, ∗$p < 0.1$. Intercept included but not reported. N = 4,748 for all models. The dependent variable is noted across the top of each model. The Random Days variable captures a random assignment of 86 days, which is the number of home football games during our sample.

Table 6: All Home Games Placebo

Dep. Variable	ADR	RevPar	Demand	Occupancy	Revenue
Random Days	0.06	0.01	-2.46	0.02	-0.23
	(0.31)	(0.47)	(47.59)	(0.40)	(5.85)
Day of Week FE	✓	✓	✓	✓	✓
Week FE	✓	✓	✓	✓	✓
Month FE	✓	✓	✓	✓	✓
Year FE	✓	✓	✓	✓	✓
R-Squared	0.85	0.86	0.89	0.88	0.87

Note: ∗∗∗ $p < 0.01$, ∗∗$p < 0.05$, ∗$p < 0.1$. Intercept included but not reported. N = 4,748 for all models. The dependent variable is noted across the top of each model. The Random Days variable captures a random assignment of 285 days, which is the number of home games in our sample (199 basketball + 86 football).

To further extend the analysis, the hotels in the area received an exogenous shock following the 2011 football season when WVU switched from the Big East conference to the Big 12. To explore this effect, a binary variable is created that takes on the value of one if the season is 2012 or later, and zero otherwise. The regressions in Tables 7 and 8 test whether changing to a more competitive conference has had a significant effect on tourism within the area, for football and basketball, respectively. The interaction term between the home football variable and the Big 12 dummy variable is the variable of interest to deduce the differential impact of home football games after the conference switch. As can be seen, the change to the Big 12 conference appears to be associated with a decrease in the occupancy rate of the hotels. However, there is a positive and statistically significant increase in the revenue generated by these hotels after the change in conference. This coefficient suggests that the change of conference is associated with a modest increase of about $38,000 per game. Basketball, however, does not appear to be affected in any way by the move to the Big 12 conference.

Table 7: Big 12 Football

Dep. Variable	ADR		RevPar		Demand		Occupancy		Revenue	
Home Football	20.72	***	24.89	***	1365.36	***	12.11	***	282.58	***
	(0.64)		(1.05)		(114.40)		(0.95)		(13.02)	
Big 12	21.32	***	19.11	*	2732.80	**	13.93		317.12	***
	(6.61)		(10.85)		(1181.02)		(9.85)		(134.46)	
Home Football * Big 12	-0.70		-1.09		-207.36		-3.81	***	38.03	**
	(0.92)		(1.51)		(164.26)		(1.37)		(18.70)	
Day of Week FE	✓		✓		✓		✓		✓	
Week FE	✓		✓		✓		✓		✓	
Month FE	✓		✓		✓		✓		✓	
Year FE	✓		✓		✓		✓		✓	
R-Squared	0.85		0.86		0.89		0.88		0.87	

Note: *** p < 0.01, **p < 0.05, *p < 0.1. Intercept included but not reported. N = 4,748 for all models. The dependent variable is noted across the top of each model. The Big 12 dummy variable is equal to one if the year of the game is 2012 or later (the year WVU joined the conference).

Table 8: Big 12 Basketball

Dep. Variable	ADR		RevPar		Demand		Occupancy		Revenue	
Home Basketball	0.53		0.90		134.34	*	1.25	*	8.69	
	(0.51)		(0.78)		(78.37)		(0.65)		(9.63)	
Big 12	12.54		8.54		2167.36	*	9.31		187.68	
	(7.90)		(12.02)		(1210.71)		(10.10)		(148.78)	
Home Basketball * Big 12	-0.84		-0.77		-50.28		-0.82		-4.06	
	(0.74)		(1.12)		(1113.00)		(0.94)		(13.89)	
Day of Week FE	✓		✓		✓		✓		✓	
Week FE	✓		✓		✓		✓		✓	
Month FE	✓		✓		✓		✓		✓	
Year FE	✓		✓		✓		✓		✓	
R-Squared	0.85		0.86		0.89		0.88		0.87	

Note: *** p < 0.01, **p < 0.05, *p < 0.1. Intercept included but not reported. N = 4,748 for all models. The dependent variable is noted across the top of each model. The Big 12 dummy variable is equal to one if the year of the game is 2012 or later (the year WVU joined the conference).

CONCLUSION

This paper has investigated the impact of WVU home football and basketball games on hotel demand and revenue. These measurements are used to uncover the extent to which major collegiate sports contribute to the economy of a college town by bringing in external tourism dollars. The statistical significance for both football and basketball games in the demand, occupancy, revenue, and average daily room rates suggest that there is a positive impact of these games on tourism. The effects are economically meaningful for football games.

One concern of any small city holding large-scale events is that they may crowd out other visitors. We find that the percentage of rooms that are occupied increases during WVU football and basketball games. On average, however, this increase in the number of occupied rooms brings Morgantown hotels to about 75% of capacity. This suggests that while there may be individual games that induce crowding out, it is not seen on average. Our results are robust to the inclusion of numerous fixed effects as well a placebo test. Lastly, we find that the switch from the Big East to the Big 12 for WVU in 2012 appears to only have a modest impact on revenue received from football games, with no impact from home basketball games. This finding is perhaps unsurprising given the geographic distance from WVU to the rest of the Big 12 (Chatmon, 2016).

Our results suggest that there is an increase in hotel revenue of over $640,000 for each home football game. (This is including the increased revenue of $360,789 for the day of a home football game and $280,781 for the day before the game.) To put these results into context, we will use the local 6% lodging tax rate to estimate the impact these games have on tax revenue using back of the envelope calculations. Doing this suggests that there is an increase of over $38,000 in additional tax revenue for each home football game. Extending these numbers throughout an entire season of 7 home games, the impact is quite large. Hotels earn an additional $4.35 million in revenue, with over $261,000 in additional tax revenue collected annually. This result of $4.35 million is about one-third of the estimates of Christiadi (2012), who evaluated the impact of WVU home football games using a survey sent to local establishments. Basketball has a much smaller impact, with only $1,100 per game in additional tax revenue, which over the course of a season, does not reach the magnitude of a single home football game.

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Would Columbus Miss the Crew? Major League Soccer and Hotel Occupancy

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ABSTRACT

In late 2017, reports indicated that the Columbus Crew soccer team would relocate to Austin, Texas. This paper uses three years of daily data to examine the effect of the Crew, as well as other sports teams and events, on hotel room rentals, hotel room rates, and hotel revenues in Columbus. The results indicate that Ohio State University football games and a three-day rock music festival generate large increases in hotel room rentals, rates, and revenues, while the Crew, an NHL hockey franchise, and a minor league baseball team have small effects. The Crew ultimately stayed in Columbus under new ownership and with a large subsidy for a new stadium; nonetheless this paper yields insights on the overnight drawing power of various sporting and other events.

INTRODUCTION

On October 16, 2017, news broke that the operator of Major League Soccer's (MLS) Columbus Crew was planning to move the club to Austin, Texas in 2019 (Arace 2017, Moore-Bloom 2019). The Crew's average home match attendance of 15,439 ranked 20th among the MLS's 22 teams, and the team's management blamed its poor attendance in part to playing in an outdated stadium. With Columbus apparently unwilling to publicly fund construction of a more modern playing facility, the team thought Austin would provide greener pastures.

Crew fans were outraged at the possibility of losing their team, especially since the Crew were among the league's ten inaugural teams in 1996 and Columbus does not have NFL, NBA, or MLB franchises. Crew fans formed an organization called Save the Crew and undertook highly publicized activities pressuring Columbus and Ohio politicians to prevent the move. On October 22, 2017 Save the Crew held a large rally at city hall, and the October 31, 2017 home game featured fans chanting "Save the Crew" (Moore-Bloom 2019). Ultimately, Ohio Attorney General Mike DeWine sued MLS and the Crew's operator/investor Precourt Ventures under the provisions of Ohio's "Art Modell law" to prevent the Crew from moving (McCann 2018). The Art Modell law, named for the team owner who moved the Cleveland Browns to Baltimore in 1996, prohibits Ohio sports franchises that have played in taxpayer subsidized facilities from relocating unless they have government permission or provide an opportunity for the team to be sold to new owners who would keep the team in Ohio. Eventually, when this paper was at an advanced stage, MLS announced the Crew would be transferred to an investor/operator group including Cleveland Browns owners Dee and Jimmy Haslam and Crew team doctor Pete Edwards and would remain in Columbus on the condition of playing in a new, publicly subsidized stadium (Ferenchik and Rouan 2018).

Of course, a common concern in franchise relocations or the hosting of prominent sport events is the economic impact associated with the event or franchise. Like many team or events, the Crew's economic impact had been analyzed. An economic impact study released by the team in 2012 reported that the Crew had generated $384 million in additional spending between 1996 and 2011, with $160 million coming from outside Franklin County (home of Columbus).[1] The study also claimed that 69% of attendees came from outside Columbus with 20% coming from outside Ohio.

Economic impact studies of this sort are controversial in the sports economics literature. Research going back at least as far as Baade and Dye (1988, 1990) casts doubt on claims of large economic gains associated with sports events and franchises. In a now somewhat dated literature review, Coates and Humphreys (2008) found little evidence of large gains relative to the frequently sizable public subsidies associated with sports teams and events. Among the limitations noted about traditional economic impact studies such as the one performed for the Crew, one of the most important is the failure to differentiate between gross and net visitors (Porter 1999, Baumann et al. 2009). Studies that fail to account for any displacement of would be visitors overstate the gains associated with the event. In an effort to get a better grasp on marginal visitors associated with sports events, recent papers such as Collins and Stephenson (2016), Depken and Stephenson (2018), and Heller et al. (2018) use granular hotel occupancy data to estimate the net effect of sports and other events on hotel occupancy, a key component of visitors' economic impact.

This paper uses a similar approach to examine the effect of Columbus Crew matches on local hotel occupancy. While it appears the Columbus will retain the Crew, this paper remains valuable because, to our knowledge, it is first paper to analyze the relationship between MLS matches and hotel occupancy. Since the MLS continues to seek public subsidies as it expands to new cities, understanding the economic impact of MLS franchises has important public policy implications. In addition, the paper controls for other events in order to avoid omitted variable bias; it also estimates the effect of Columbus Blue Jackets National Hockey League (NHL) games, Ohio State University (OSU) football and basketball games, Columbus Clippers minor league baseball games, OSU graduation ceremonies, and a large rock music festival on hotel occupancy.

EMPIRICAL FRAMEWORK

We use three years of daily hotel occupancy data from the Columbus metropolitan area spanning December 1, 2014 to November 30, 2017 (1,066 observations) in the analysis. The data are obtained from STR, a firm that compiles data from hotels in the U.S. and many other countries.
Our regression model is as follows:

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[1] https://www.columbuscrewsc.com/post/2012/02/08/columbus-crew-nets-regional-economic-benefits

$$DEP_t = \beta_0 + \beta_1 EVENTS + \beta_2 UNEMPRATE + \beta_3 DAY + \beta_4 WEEK + \varepsilon_t.$$

where DEP_t is the number of hotel rooms rented on night t, the average daily rate (ADR) or price of the rooms rented on night t, or the aggregate amount of hotel revenue on night t. Descriptive statistics for the dependent variables are reported in Table 1.

Table 1: Descriptive Statistics

	Mean	Std. Dev.	Min.	Max.
Rooms	17,651	4,371	5,945	26,599
ADR	99.18	10.57	65.82	131.87
Revenue	1,792,012	585,573	401,093	3,507,765
Unemp Rate	4.86	0.42	4.1	5.7

The vector EVENTS contains dummy variables for Columbus Crew home matches, Columbus Blue Jackets home games, Columbus Clippers home games, home games for OSU football and basketball, OSU's spring, summer, and autumn commencement ceremonies, and the Rock on the Range music festival. OSU football games (21) are the least common event in our study while Columbus Clippers baseball games (220) are the most common. The Crew had 56 home matches during our study period, or roughly 18-19 per season. Counts for all of the included events are reported in Table 2. For Crew matches and OSU football games, we also include dummies for the night before scheduled games in order to capture visitors who arrive early. We also experimented with a dummy for two days before Crew matches and OSU football games, a dummy for the day after Crew matches and OSU football games, and dummies for the day before Blue Jackets games and OSU basketball games, but they are omitted because their inclusion has no effect on the estimation results reported below. Our not needing to include day after effects for any events and day before effects for only Crew matches and OSU football games indicates that there is little evidence that Columbus's sports events attract fans who come in advance of the events or stay beyond the events.

Table 2: Event Counts

Event	Number of Occurrences
Columbus Crew Home Matches	56
Ohio State Football Home Games	21
Ohio State Basketball Home Games	58
Columbus Blue Jackets Home Games	134
Columbus Clippers Home Games	220

Since the number of graduates likely differs for the commencement ceremonies, separate dummy variables are created for the spring, summer, and autumn commencements. A dummy for the night before each of the graduation exercises is also included since guests might arrive the day before the event. Rock on the Range is a Friday to Sunday music festival held in May of each year. To allow for possible spillover effects of the festival on hotel occupancy, we also include the Thursday before the festival and the Monday after the festival in the regression models.

To control for other factors affecting hotel occupancy, the unemployment rate (UNEMPRATE) is included to control for changes in macroeconomic conditions. The model also includes fixed effects for days of the week (DAY) since hotel occupancy can vary systematically across days of the week. Similarly, fixed effects for week in the year (WEEK) are included to control for systematic variation in hotel occupancy across seasons or because of holidays. (We also experimented with a trend variable but it had no qualitative effect on the results.)

ESTIMATION RESULTS

The hotel occupancy data are stationary so the model is estimated via OLS with Newey-West corrected standard errors to control for serial correlation. The estimation results are reported in Table 3.

Crew matches have no statistically significant effect on the number of hotel rooms let. The magnitude is small—an additional 273 rooms rented on match days and 499 rooms rented on the night before matches. These effects are smaller than would be expected based on the Crew's economic impact report. During the 2015-2017 seasons, the Crew averaged about 16,000 fans per home game so the economic impact report implies more than 3,000 would come from outside Ohio and would presumably need more than a few hundred rooms of lodging per night to accommodate them. Hotel data would fail to detect out of state visitors who stay with local friends or family members or in AirBnBs, though short term rentals account for a relatively small share of lodging relative to hotels. Another possible explanation for the discrepancy between our findings and the claims of the economic impact report would be if that report classified out of state OSU students who attended games as being out of state residents even though they are at least temporarily residing in Columbus. Likewise, the estimated coefficients on the Crew variables in the ADR regression are small (less than $1.15) and not statistically significant. The ADR results indicate that Crew games are not associated with increases in daily room rates, a result which buttresses the conclusion that Crew matches do not appreciably increase the demand for hotel rooms. With both the number of rooms let and the ADR effects associated with Crew matches being small and statistically insignificant, it is unsurprising that Crew matches also have no effect on hotel revenue.

Looking at the other events, OSU football has a significantly positive effect on room rentals, ADR, and hotel revenue. The number of rooms rented increases by about 3,650 over the two nights, a figure roughly one-third larger than Bonneau and Hall (this issue) find for West Virginia University home football games. OSU football games also see prices increase $7-10 per room and hotel revenue increase by more than $730,000. OSU basketball games are associated with increases of 357 rooms let and $55,000 in hotel revenue, though neither effect is statistically different from zero. Neither Blue Jackets hockey games nor Clippers baseball games has an economically or statistically significant increase in hotel rooms rented, ADR, or hotel revenue.

OSU's graduation ceremonies also have large effects on the number of hotel rooms let. The largest effect is associated with the autumn graduation which increases room rentals by about 7,360 room nights, ADR by $8-11 per night, and hotel revenue by $953,000. The spring commencement increases room rentals by about 4,000, ADR by roughly $5-9 per night, and hotel revenue by nearly $680,000. The summer graduation has the smallest effects but it still generates nearly 1,900 additional room nights and approximately $240,000 in additional hotel revenue.

Table 3: Estimation Results for Columbus Hotel Occupancy

	Dependent Variable		
	Rooms	ADR	Revenue
Day Before Crew	370.46	0.69	39,989
	(1.03)	(0.63)	(0.70)
Crew	151.44	-0.02	4,593
	(0.42)	(0.02)	(0.08)
Day Before OSU Football	1,319.84*	7.13**	259,091**
	(2.36)	(3.47)	(2.83)
OSU Football	2,152.60**	10.13**	434,279**
	(3.48)	(4.19)	(3.91)
OSU Basketball	374.74	0.89	57,975
	(1.19)	(0.88)	(1.25)
Blue Jackets	-91.11	-0.60	-23,959
	(0.49)	(1.02)	(0.87)
Clippers	-282.17	-1.74**	-62,353
	(1.23)	(2.67)	(1.84)
Day Before Spring Grad.	3,247.86**	9.14**	545,808**
	(4.81)	(4.59)	(4.67)
Spring Graduation	926.74	4.88*	150,058
	(1.89)	(2.23)	(1.87)
Day Before Summer Grad.	1,412.69**	2.92	195,034*
	(3.29)	(1.55)	(2.33)
Summer Graduation	532.94	1.55	57,650
	(0.87)	(0.96)	(0.70)
Day Before Autumn Grad.	2,548.42*	10.12**	301,311*
	(2.50)	(2.72)	(2.29)
Autumn Graduation	4,658.08**	8.56**	624,944**
	(5.49)	(2.77)	(5.99)
Rock on the Range			
Thursday	2,919.69**	1.98	327,076.75**
	(6.88)	(1.67)	(5.00)
Friday	6,150.75**	20.85**	1,110,373.55**
	(10.32)	(15.04)	(15.46)
Saturday	5,587.96**	22.01**	1,101,812.71**
	(7.56)	(13.14)	(11.22)
Sunday	5,923.12**	9.72**	695,358.61**
	(7.17)	(4.50)	(6.10)
Monday	1,850.73**	4.69**	250,883.10**
	(3.10)	(3.75)	(3.53)
Unemp. Rate	-982.31**	-4.34**	-180,673.49**
	(5.23)	(7.61)	(6.52)
Constant	16,513.03**	103.36**	1,887,098.46**
	(15.06)	(29.53)	(11.98)

Parentheses contain t-statistics derived from Newey-West standard errors. * and ** indicate statistical significance at the 5% and 1% level, respectively. The models also include fixed effects for days of the week and weeks of the year.

The Rock on the Range music festival has large effects on hotel room nights, ADR, and hotel revenue, and its effects begin the Thursday before the festival and continue through the Monday following the festival. Over the five days, the festival generates more than 22,000 room nights and nearly $3.5 million in additional hotel revenue. ADR increases by $21-22 for the Friday and Saturday nights of the festival and by smaller amounts on the other three nights.

Macroeconomic conditions also have a large effect on the Columbus hotel market; a one percentage point increase in the unemployment rate is associated with almost 1,000 fewer rooms rented per night, a $4 per night decrease in ADR, and a reduction of $179,000 per night in hotel revenue.

CONCLUSION

Contrary to claims by the Crew's economic impact report, this paper's results indicate that the Crew have little effect in attracting overnight visitors to Columbus. Hence, it seems that most Crew fans reside in Columbus or are close enough for day trips. While day trippers might increase economic activity in bars, restaurants, or souvenir shops, much spending associated with the Crew is likely redirected from other local activities. While some passionate fans would have missed the Crew, losing the team would have had little effect on overnight visitors to Columbus. Yet taxpayers in Ohio, Franklin County, and Columbus are now expected to pay at least $140 million for the new publically subsidized stadium (Bush 2019). The lack of overnight visitors to Columbus should also be of concern to cities such as nearby Cincinnati which is currently partially funding a stadium for an MLS expansion franchise through hotel taxes (Seitz 2018).

As previously noted, the overall economic impact of stadiums has long been questioned by economists (Coates and Humphreys 2008). However, the Crew's ability to obtain taxpayer subsidies by threatening to move to Austin highlights how MLS and other closed sports leagues exploit their monopoly power to restrict the supply of teams and foster competition among cities. Perhaps a new stadium will generate more overnight visitors and economic gain than the existing stadium, but if it doesn't much of the economic benefit of the new stadium will be internalized by the Crew's owners (Humphreys and Zhou 2015).

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Fan Reaction to Pace-of-Play Rule Changes: Game Duration and Attendance in Major League Baseball

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ABSTRACT

Major League Baseball introduced pace of play initiatives to speed up the game. Using an attendance model with individual per-game attendance as the dependent variable and game duration, measured by minutes per out, and control variables such as team quality, outcome uncertainty, weather, etc. as independent variables, we test whether the length of a game influences attendance. We find that, after the rule changes were announced, teams that played shorter games were rewarded by fans through higher attendance.

INTRODUCTION

In 2014, Major League Baseball (MLB) expanded its replay and challenge system during a game and announced proposed rule changes related to games' pace of play. The proposed rule changes were designed to add more excitement to the game by lessening time where no action is occurring on the field and to decrease the overall length of games. At the time, game length had increased by fifteen minutes since 2005 and nearly an hour since 1981. Some of the increased game length had to do with in-game advertising, but the main changes occurred due to increased substitutions in terms of pitchers and hitters throughout the game.

The proposed rule changes in 2014 consisted of a twenty-second pitch clock, enforcement of the batters' box rule, no-pitch intentional walks, a maximum of two minutes and five seconds in the break between innings, a two-and-a-half-minute maximum on length of pitching change delays, and a limitation on mound meetings. The proposed rule changes were hotly debated between MLB and the Players Association, and only some of the proposed changes have been implemented.

A timeline of actual pace-of-play rule changes in MLB is outlined below. In 2015, MLB stated that the batters' box rule in the rulebook would be more strictly enforced. Specifically, batters must keep one foot in the batters' box, with various exceptions that can occur. Penalties for non-adherence to this rule involved individual player "progressive discipline" from the league. Also, the league added timers to measure non-game action, specifically the break time between innings and duration of pitching changes. For 2015, the rules were a maximum of two minutes and twenty-five seconds for locally broadcast games and two minutes and forty-five seconds for a nationally televised game. Also, beginning that year, managers were able to initiate instant replay challenges directly from the dugout, instead of adding the time it would take for them to walk on the field to challenge the play.

The following season, 2016, pace of play rules were amended to add a thirty-second time limit for visits to the mound by pitching coaches or managers. In addition, the timers for break time in the game for between innings and pitching changes were reduced by twenty seconds to two minutes and five seconds for local broadcasts and two minutes and twenty-five seconds for national broadcasts.

In 2017, no pitch intentional walks, which allowed managers to simply call for an intentional walk from the dugout, were added. In addition, a thirty-second-time limit was placed on managers to challenge a play on the field. Also, although it contains many exceptions and has not been enforced regularly, the league stipulated a two minute limit for umpires to decide replays. For the 2018 season, pace of play rules were also adjusted. Mound visits by pitching coaches and managers are limited to six per game, with one additional mound visit per inning if the game goes to extra innings. In 2019, the number of mound visits were reduced to five per nine innings.

Although initial tests of the pace of play rules were positive, including a test in the Arizona Fall League in 2014 where game duration fell by ten minutes, the rules have not appeared to be overly successful at reducing the length of games in MLB. In 2014, the average length of a game was just over three hours (3:02). In 2015, the average length of games fell below three hours (2:56), perhaps showing some positive results of the rule changes. However, in 2016 the average length of a game rose back to three hours and in 2017 the average game length hit a record of three hours and five minutes. There are also substantial differences across teams within seasons.

Since the purpose of the pace of play rule changes was increasing the appeal of games to consumers, this study tests if the pace of play for different teams in MLB significantly influences attendance at games. Using four years of data, 2014-2017 on all MLB games played, an attendance model is specified to test if game duration, all else equal, tends to have any significant impact on attendance and, if so, by how much. Controlling for a variety of factors including team success, uncertainty of outcome, weather effects, days of the week, months of the season, opening day, etc., we test whether a variable called minutes per out is related to attendance. We calculate minutes per out from the game length information and game details on www.retrosheet.com (not all baseball games contain the same number of outs based upon if the home team wins the game without having to bat in the bottom of the ninth inning, games with extra innings, weather-shortened games, etc.). Through several specifications, we hope to determine MLB gate attendance is related to game duration or if pace of play is an issue with which the league should not be overly concerned.

The paper proceeds as follows. Section II contains a brief overview of the literature on baseball attendance. Section III describes the variables used in the regression model and shows the empirical results. The last section discusses the implications of the results and concludes the paper.

LITERATURE REVIEW

Prior to the implementation of the pace of play rules, Paul et. al. (2016) examined game duration as it related to attendance and found a statistically insignificant relationship between the two variables. Beyond that study, we did not find other research related to game duration and attendance. However, there are studies that have contributed to the regression model used in this paper. There is a wide range of published papers on attendance in baseball at both the game-by-game and season-by-season levels. Some key studies in the history of research on baseball attendance, with the key independent variables they introduced, are noted below.

To describe all the papers on baseball attendance would be beyond the scope of the literature for this study, but some important papers that influenced our choice of regression models in this study are noted below. Demmert (1973) studied the roles of televised games, team quality, and availability of substitute sports in baseball attendance. Noll (1974) introduced the role of population, income per capita, recent team success, and star players. The impact of a new stadium on attendance has been studied (Coates and Humphreys, 2005; Depken, 2006) to ascertain the magnitude and sustaining impact of the investment into a new stadium for the local team. Other key independent variables introduced by different studies in the literature include the expected probability of winning a championship (Whitney, 1988), the role of salary structure of a team on attendance (Richards and Guell, 1998), turnover in team rosters (Kahane and Shmanske, 1997), and interleague play (Butler, 2002; Paul et al., 2004).

Many of the key research papers related to baseball attendance in general are discussed in formal literature reviews on the topic that have appeared in a variety of journals. Studies by Schofield (1983), MacDonald and Rascher (2000), and Villar and Guerrero (2009) have highlighted the literature surrounding the topic and each gives a perspective for the differences in these studies by era.

REGRESSION MODEL OF GAME DURATION AND MLB ATTENDANCE

To examine the role of game duration on attendance, we used an ordinary least squares (OLS) regression model. For each MLB team, all home games are included in the sample for the four seasons following the introduction of the new pace of play rules in 2014. Attendance for each game was gathered from the compiled box scores as reported on www.retrosheet.com.

The key variable of interest for this study is the game duration, measured as minutes per out. This variable was constructed by taking the number of minutes in each game played (taken from www.retrosheet.com) and dividing it by the number of outs in each game. In baseball, the assigned duration of the game is 9 innings, with 3 outs per half-inning for a total of 54 outs. Not all games consist of 54 outs, however, as when the home team is leading after the visitor bats in the 9th inning, the home team has won the game and does not bat, resulting in a game with 51 outs. If the home team takes the lead in the bottom of the 9th inning the game ends immediately so the game may have between 51-53 outs. When the road team wins in a 9-inning game, there will be 54 outs. When games are tied after 9 innings, the game goes into extra-innings, resulting in more than 54 outs. Also, there are sometimes weather-shortened games, which have fewer than 54 outs. Minutes per out therefore allows the pace of the game to be compared more evenly across all games. Minutes per out is calculated for each home game, but the variable used in the regression model is a running average throughout the season for each team. Specifically, the minutes per out was represented as the home team's average for their home games entering that day's game. If fans do not enjoy longer baseball games, minutes per out should have a negative and statistically significant impact on game attendance.

Team performance is likely to influence attendance for MLB teams. To control for past performance, the win percentage from the previous season for the home team is included in the regression model. Additionally, current season home and visiting win percentages entering the game were also included as independent variables in the regression model to account for within-season team performance. It is assumed that fans prefer successful teams to unsuccessful teams; therefore, home win percentage should have a positive and significant effect on attendance. The impact of the visiting team win percentage is ambiguous, as less successful visiting teams should make it more likely for the home team to win, which home fans are likely to find favorable, but successful visiting teams likely have better players and more superstars, which fans of the home team likely would enjoy viewing in person as well. If fans want to see successful visiting teams, either because there are some fans of the road team in attendance, home team fans hope to see their team beat other quality teams in the league in person, or baseball fans in general enjoy seeing talented players and teams, this variable should have a positive effect on attendance.

A dummy variable for the home opener was included as an independent variable in the regression model to account for the popularity of the festivities surrounding opening day in baseball. Historically, fans have flocked to opening day, even in cold-weather cities, as the start of the baseball season is met with much fanfare. If fans enjoy attending the first game of the season, this variable should have a positive and significant effect on attendance. There are other holidays where baseball has also become a tradition and likely see significant increases in attendance. These days are Memorial Day, July 4th, and Labor Day. Dummy variables for these holidays are also included in the regression model.

Days of the week and months of the season dummy variables were included in the regression model to control for any systematic patterns in attendance. In terms of weekdays, we expected weekends to be more popular than weekdays for attendance due to the opportunity cost of fans' time. The reference category for the days of the week is Wednesday. In terms of months of the season, outside of opening day, early season games are often less popular than games in the summer. Therefore, we would expect lower early season attendance. Summer months are likely to be most popular and late-season games could be popular in places where playoff races are still active, and less popular in cities where teams are not in playoff contention.

Betting market data from www.covers.com was included in the regression model to account for uncertainty of outcome and expected scoring. The odds on the games were converted into win probabilities and the win probability (and in some specifications, its square) are included to account for expectations of game outcome. The betting market total, in terms of the numerical value stipulated in the wagering market, was included in the regression model to account for expected scoring to control for the possibility that fans like higher scoring games. The total has multicollinearity issues with game duration as games that are expected to be higher-scoring typically take longer to play. Therefore, model specifications with and without the total included in the model are presented.

Weather variables likely affect fans' decisions to attend games, but also could influence how long a game takes to play as adverse conditions could lead to longer games. A wide variety of variables are available and archived on www.weatherunderground.com. We used this resource to capture daily information on temperature, humidity, wind speed, and precipitation (in inches). We would generally expect temperature to have a non-linear relationship with attendance as warmer weather is preferred to cold, but temperatures that are too high are also uncomfortable when choosing to attend an outdoor sporting event. We also allow for a non-linear relationship with humidity and attendance, although we expected that higher humidity would decrease attendance. There are some multicollinearity issues between temperature and humidity, so regression model results with both variables included and with only temperature are presented. Wind could also make the game less enjoyable to attend, so we also expected a negative relationship with this variable. Precipitation should also detract from attendance, therefore the greater the amount of precipitation, the fewer fans we expected to see in attendance for baseball games.

Fans' desire to attend baseball games can change over time for reasons other than pace of play (e.g., the presence of new stars or a changes in the availability and attractiveness of substitute activities). Including year dummies can, therefore, be an important control for unobservable changes over time, but could also pick up the rule changes administered in each season. With this in mind, a separate set of regression models were run, and the results are shown in the paper as a form of robustness check. In the models where yearly effects are included, dummy variables for the individual seasons were incorporated into the regression (all results compared to the reference category of 2014). In all models shown, home and road team dummies are included. The summary statistics for the non-binary variables are included in Table I.

Table I: Summary Statistics

Variable	Mean	Median	Standard Deviation
Attendance	30,264.93	30,434.00	9,758.60
Minutes	185.53	182.00	27.93
Outs	53.59	54.00	5.05
Minutes Per Out	3.46	3.44	0.39
Home Win Probability	0.54	0.55	0.08
Total	8.20	8.00	1.06
Temperature	70.51	72.00	10.92
Humidity	64.62	66.00	14.19
Wind Speed	6.83	7.00	3.74
Precipitation (in)	0.09	0.00	0.28

OLS regression results are shown in Table II and Table III below. The sample includes 9684 MLB games over four seasons (2014-2017), including all games played except where data on attendance or game duration were missing from the records. Each result is presented using Newey-West HAC-consistent standard errors and covariances due to initial results indicating heteroskedasticity and autocorrelation issues. In both tables, *-notation denotes statistical significance at the 10% (*), 5% (**), and 1% (***) levels.

There are eight model specifications shown for validation and to allow for some exploration of the interaction of some variables that may be multicollinear. The major difference between the results presented in Tables II and III relate to the inclusion of weather-related variables and season dummies. In Table II neither of these categories of independent variables are included in the model, while in Table III both weather variables and year dummies are included. Across the different specifications shown in each table, the key differences revolve around the way the home team win probability and the total were treated within the respective models. For Table II, the first specification includes only the home team win probability and not the total. The second specification includes the home team win probability and its square, but not the total. The third specification includes the home team win probability and the total, while the fourth specification includes the home team win probability, its square, and the total.

In Table III, in the first specification (model V), the home team win probability and both temperature and humidity and their squares are included in the model. Specification II in Table III (model VI) includes everything in model V, but also includes the square of the home team win probability variable. Model VII in Table III includes everything in model V except for the betting market total, while the last specification (model VIII) includes everything in specification I except humidity and humidity squared.

Table II: Game Duration and MLB Attendance – 2014-2017 Seasons

Variable	I	II	III	IV
Intercept	20,763.28***	24,9441.20***	21,626.68***	25,841.15***
	(11.7420)	(9.1288)	(12.1538)	(9.4105)
Minutes Per Out	-1,297.88***	-1,359.14***	-930.25**	-990.20**
	(-2.9792)	(-3.1121)	(-2.0736)	(-2.2026)
Lagged Win %	20,376.85***	20,290.04***	20,136.64***	20,047.80***
	(18.8828)	(18.7874)	(18.6528)	(18.5574)
Home Win %	6,523.81***	6,506.20***	6,449.40***	6,431.04***
	(10.5647)	(10.5383)	(10.4506)	(10.4227)
Visitor Win %	4,680.64***	4,682.54***	4,525.57***	4,527.13***
	(7.5495)	(7.5446)	(7.3148)	(7.3093)
Home Opener	16,026.61***	16,026.33***	16,100.00***	16,099.36***
	(21.9442)	(21.9857)	(22.1210)	(22.1623)
Memorial Day	5,657.94***	5,652.83***	5,724.36***	5,719.56***
	(6.5907)	(6.5867)	(6.6911)	(6.6875)
July 4th	3,427.28***	3,415.26***	3,426.07***	3,413.95***
	(4.0856)	(4.0805)	(4.0688)	(4.0634)
Labor Day	3,261.66***	3,291.94***	3,293.84***	3,324.55***
	(4.6013)	(4.6524)	(4.6634)	(4.7153)
Sunday	5,335.56***	5,341.29***	5,356.16***	5,362.03***
	(28.4922)	(28.5243)	(28.5655)	(28.5994)
Monday	-1.141.66***	-1,141.02***	-1,129.27***	-1,128.61***
	(-5.2295)	(-5.2269)	(-5.1696)	(-5.1670)
Tuesday	-414.19**	-415.87**	-407.72**	-409.39**
	(-2.2296)	(-2.2388)	(-2.1971)	(-2.2063)
Thursday	566.57***	566.01***	570.75***	570.21***
	(2.6696)	(2.6678)	(2.6905)	(2.6887)
Friday	4,545.50***	4,546.87***	4,541.34***	4,452.71***
	(24.9319)	(24.9324)	(24.9154)	(24.9163)
Saturday	7,714.21***	7,715.70***	7,729.19***	7,730.76***
	(41.1666)	(41.1778)	(41.2720)	(41.2852)
March	1,395.95	1,422.22	1,067.57	1,093.13
	(0.6359)	(0.6470)	(0.4842)	(0.6205)
April	-2,929.99***	-2,912.53***	-3,022.04***	-3,004.88***
	(-15.6212)	(-15.5154)	(-15.9912)	(-15.8884)
May	-2,002.49***	-1,985.54***	-2,057.44***	-2,040.61***
	(-11.3777)	(-11.2788)	(-11.6762)	(-11.5783)
July	1,084.16***	1,076.57***	1,087.76***	1,080.13***
	(6.0213)	(5.9798)	(6.0376)	(5.9961)
August	-302.69*	-319.10*	-310.34*	-326.92*
	(-1.7483)	(-1.8415)	(-1.7939)	(-1.8880)
September	-2.044.07***	-2,087.12***	-2,063.68***	-2,107.16***
	(-11.5226)	(-11.7049)	(-11.6482)	(-11.8308)
October	-2.198.79***	-2,249.97***	-2,255.47***	-2,307.31***
	(-3.5601)	(-3.6469)	(-3.6471)	(-3.7348)
Home Win Probability	6,457.02***	-8,254.18	6,459.29***	-8,366.80
	(7.4445)	(-1.1141)	(7.4515)	(-1.1260)
Home Win Probability2		13,664.38**		13,771.31**
		(1.9906)		(2.0000)
Total			-219.78***	-220.82***
			(-3.3155)	(-3.3343)
Home Team Dummies	Yes	Yes	Yes	Yes
Road Team Dummies	Yes	Yes	Yes	Yes
R-squared	0.7220	0.7222	0.7225	0.7227
Adjusted R-squared	0.7197	0.7199	0.7202	0.7203

Table III: Alternate Specifications of Game Duration and MLB Attendance – 2014-2017 Seasons

Variable	V	VI	VII	VIII
Intercept	10,527.88***	14,485.87***	9,033.83***	10,554.87***
	(3.0161)	(3.5961)	(2.6114)	(3.2993)
Minutes Per Out	-972.02*	-1,013.12*	-1,141.68**	-951.65*
	(-1.7753)	(-1.8504)	(-2.0902)	(-1.7372)
Lagged Win %	20,222.73***	20.154.40***	20,508.23***	20.357.33***
	(18.7665)	(10.6844)	(19.0538)	(18.7932)
Home Win %	6,546.21***	6,535.91***	6,632.30***	6,542.36***
	(10.6964)	(10.6844)	(10.8317)	(22.8987)
Visitor Win %	4,603.63***	4,612.71	4,761.73***	4,617.95***
	(7.5034)	(7.5115)	(7.7541)	(7.5403)
Home Opener	16,475.99***	16.478.79***	16,404.11***	16,452.32***
	(22.8477)	(22.8826)	(22.6004)	(22.8987)
Memorial Day	5,467.28***	5,460.05***	5,420.60***	5,378.58***
	(6.2680)	(6.2622)	(6.1905)	(6.2496)
July 4th	3,398.65***	3,386.68***	3,388.84***	3,424.94***
	(4.0202)	(4.0138)	(4.0157)	(4.0290)
Labor Day	3,263.74***	3,292.44***	3,249.33***	3,228.26***
	(4.6089)	(4.6583)	(4.5812)	(4.5532)
Sunday	5,283.23***	5,288.09***	5,260.68***	5,306.29***
	(28.3420)	(28.3727)	(28.2624)	(28.4474)
Monday	-1,162.81***	-1,161.96***	-1,173.87***	-1,150.59***
	(-5.3526)	(-5.3489)	(-5.4047)	(-5.2905)
Tuesday	-406.05**	-407.95**	-410.27***	-407.45**
	(-2.2050)	(-2.2156)	(-2.2259)	(-2.2113)
Thursday	594.61***	593.26***	588.08***	590.81***
	(2.8265)	(2.8211)	(2.7940)	(2.8090)
Friday	4,562.34***	4,563.42***	4,565.08***	4,557.31***
	(25.2072)	(25.2061)	(25.2181)	(25.1481)
Saturday	7,714.17***	7,714.73***	7,696.15***	7,722.54***
	(41.4684)	(41.4767)	(41.3600)	(41.4918)
March	1,108.60	1,114.43	1,259.41	1,431.94
	(0.5026)	(0.5050)	(0.5752)	(0.6531)
April	-2,445.41***	-2,427.26***	-2.408.54***	-2,265.51***
	(-10.3350)	(-10.2438)	(-10.1946)	(-9.8720)
May	-1,816.08***	-1,798.03***	-1,790.31***	-1,729.92***
	(-9.4577)	(-9.3592)	(-9.3122)	(-9.0650)
July	996.05***	989.42***	1,007.35***	966.35***
	(5.4154)	(5.3806)	(5.4835)	(5.2725)
August	-402.02	-417.32**	-383.46***	-458.05***
	(-2.2843)	(-2.3691)	(-2.1778)	(-2.6140)
September	-2,062.83	-2,104.13***	-2,051.55***	-2,098.99***
	(-11.5770)	(-11.7478)	(-11.5068)	(-11.7749)
October	-1,684.83***	-1,733.63***	-1,673.55***	-1,776.34***
	(-2.7620)	(-2.8456)	(-2.7496)	(-2.8827)
Home Win Probability	6,382.37***	-7,869.45	6,390.90***	6,406.09***
	(7.4585)	(-1.0780)	(7.4599)	(7.4746)
Home Win Probability^2		13,237.01**		
		(1.9622)		
Total	-250.78***	-240.08***		-249.51***
	(-3.2551)	(-3.1137)		(-3.2350)
Temperature	306.62***	305.93***	305.78***	300.93***
	(4.6494)	(4.6351)	(4.7092)	(4.3477)
Temperature^2	-2.03***	-2.03***	-2.06***	-1.93
	(-4.2700)	(-4.2597)	(-4.3810)	(-3.8736)
Humidity	32.48	31.51	31.58	
	(1.0400)	(1.0083)	(1.0179)	
Humidity^2	-0.45*	-0.44*	-0.44*	
	(-1.8472)	(-1.8075)	(-1.8217)	
Wind Speed	-69.05***	-69.03***	-68.08***	-65.78***
	(-3.7693)	(-3.7647)	(-3.7192)	(-3.5871)
Precipitation	-464.17**	-456.33**	-472.56**	-854.73***
	(-2.5210)	(-2.4711)	(-2.5635)	(-5.0211)
2015	-130.24	-139.83	-158.71	148.32
	(-0.7401)	(-0.7946)	(-0.9020)	(-0.8441)
2016	-206.53	-252.15	-361.21**	-233.32
	(-1.2967)	(-1.5705)	(-2.3769)	(-1.4644)
2017	-29.38	-74.87	-286.24*	-104.28
	(-0.1739)	(-0.4394)	(-1.8843)	(-0.6171)
Home Team Dummies	Yes	Yes	Yes	Yes
Road Team Dummies	Yes	Yes	Yes	Yes
R-squared	0.7266	0.7267	0.7261	0.7259
Adjusted R-squared	0.7240	0.7241	0.7236	0.7233

Results are analyzed for the models included in Table II first, where weather-related variables and season dummies were not included in the models. The key variable of interest in this study, the minutes per out per game, has a negative and statistically significant effect on attendance at the 1% level in specifications I and II and at the 5% level in specifications III and IV. Teams which played longer games, all else equal, were found to attract fewer fans to the ballpark since the new rules on pace of play were initiated. For each additional second added per out in a game, the number of fans decreased by around 16-22 fans, across the four specifications (the coefficient on minutes per out divided by 60).

Team performance also affects attendance for MLB games. Home team performance from the previous season is positive and statistically significant at the 1% level in all four specifications. Both home team and road team performance, measured by the win percentage of the team, have positive and significant effects at the 1% level on attendance. The coefficient on home team performance is larger than road team performance, as would be expected, but the relatively large coefficient on road team win percentage implies that fans also pay attention to the quality of the road team when choosing games to attend.

In terms of the scheduling of games during the season, special events and holidays including the home opener, Memorial Day, July 4th, and Labor Day have positive and significant effects on attendance. For the days of the week, weekend days have a statistically significant effect on attendance. Saturday is the most popular day for attending games, followed by Sunday, and then Friday. Thursday also has a positive and significant effect on attendance compared to the reference day category of Wednesday. Monday and Tuesday have negative and statistically significant results compared to the reference category of Wednesday. The monthly dummies revealed that July is the best month for MLB attendance, followed by the reference category of June. Both early season and late season games have fewer fans at games with these monthly dummies having negative signs and being statistically significant.

The uncertainty of outcome involving the individual game, measured by the home team win probability calculated from the betting market odds on the game, was shown to have a positive and significant effect on attendance. Fans of MLB appear to prefer less outcome uncertainty as they attend more games where the home team is expected to win. The second specification shows the result of an alternative model which included both the home team win probability and the home team win probability squared, and the squared term was shown to dominate with a positive and significant effect on attendance. Given the coefficients on the variables and the range of home team win probabilities in the sample, this again illustrated the fans' desire for less uncertainty of outcome in games.

The betting market total, where included in the model, has a negative and statistically significant effect on attendance. This reinforces the results for minutes per out, as higher scoring games generally take longer to complete. Specifications I and II, which do not include the total in the model, reveal the minutes per out variables are still negative and significant even without controlling for total. Although scoring has been shown to be popular for viewing games on television in different sports, fans do not appear to enjoy attending higher-scoring baseball games.

Table III presents alternative models including weather-related variables and season dummy variables. As noted previously, these variables are related to game duration as poor weather conditions could lead to longer games and seasonal dummies could be capturing the rule changes made in relation to pace of play for that individual season. Even with the possibility of multicollinearity issues, however, the variable of interest, minutes per out, continues to have a negative and statistically significant effect on attendance.

Turning to the control variables added in Table III, weather-related variables have statistically significant effects on attendance. Temperature has a non-linear relationship with attendance, as temperature has found to have a positive coefficient but temperature squared has a negative coefficient. Humidity squared has a negative effect on attendance, indicating high humidity discourages fans from attending games. Wind speed and precipitation also have negative and significant effects on attendance as fans stay away from games on days with precipitation and/or high winds. Dummy variables for seasons are statistically significant only in the model that did not include the betting market total; in this model the 2016 and 2017 have a negative effect.

CONCLUSIONS

Starting in the 2014 season, MLB instituted new rules to increase the pace of the game and shorten the overall length of games. Given the opportunity cost of time to fans, which has risen over time due to technology and additional entertainment options available on-demand, league officials hoped to offer an improved product, prevent the loss of fans who may abandon the game due to its increasing time commitment, and attract new fans to the game.

This study investigates whether the length of a MLB game affects game attendance. To properly account for the varying length of baseball games, we calculate minutes per out for each game and create a running average for each home game for the first four seasons following the rule changes. This variable is included in linear regression models with per game attendance as the dependent variable and independent variables including measures of home and visiting team quality, uncertainty of outcome, expectations of scoring, timing of the game in terms of weekday, month, and holidays/opening day, and detailed weather effects across different specifications

The regression model results indicate that MLB fans prefer to watch better teams play (both home and road team quality had positive coefficients and were statistically significant), prefer less uncertainty of outcome (would prefer to see the home team win), attend more games on weekends, and are impacted by inclement weather (precipitation, cold temperatures, etc.). In terms of the key variable of interest of the study, the minutes-per-out proxy for the pace of play, the variable had a negative effect on attendance, thereby indicating that fans appear to reward teams that play faster-paced games.

A particularly surprising aspect of this result is not that fans may prefer shorter games, as the pace of play rule changes were instituted for this reason, but how fans reacted after the focus on pace of play in the media compared to before. Paul et al. (2015) examined the role of pace of play on attendance in the seasons before the announced and initiated rule changes (2011-2013) and found statistically insignificant results for the minutes per out variable, suggesting fans did not seem to attend fewer games due to increasing duration of the contests. The strong negative reaction to the minutes per out variable post-rule changes, however, suggests a form of behavioral bias on the part of MLB fans. It appears that MLB fans may behave according to an observational selection bias, where after the media attention and league focus on pace of play, fans paid more attention to the length of games, choosing to attend fewer in situations where the minutes-per-out increased and more when minutes-per-out decreased. An observational selection bias is the effect of noticing things now that were not noticed before and assuming their frequency has increased. It appears the focus on pace of play may have become a self-fulfilling prophecy, as fans reacted to a condition which has always been there (baseball being a game that takes longer to play than other substitute sports) after league and media attention brought it to the forefront of their thoughts.

Given this result post-rule change compared to pre-rule change, it is difficult to speculate if the positive fan reaction to teams who play shorter games is merely transitory, due to the behavioral bias, or has brought out a lasting, more permanent effect on MLB attendance. Much may depend on how many new fans the sport attracted due to the focus on pace of play, who may be much more sensitive to this variable, compared to fans with a much longer attachment to baseball. Those long-term fans, due to the behavioral bias that appears to be occurring, may only care about pace of play if it is a hot topic, and perhaps be less sensitive to this variable in the long-run.

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Waving the White Flag: The Attendance Effects of Trading Away Talent by Contending Teams

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ABSTRACT

This paper analyzes the attendance effect of a White Flag trade, one in which a team seems to give up when it still has a chance to make the playoffs. We hypothesize that the attendance should drop following a White Flag trade; however, neither of the two examples of White Flag trades analyzed in the paper has a significant decrease in attendance. These results suggest White Flag trades do not harm attendance, but properly accounting for confounding variables such as properly accounting for team success or for the effects of the baseball strike of 1994 remains a concern.

INTRODUCTION

There are occasions when baseball teams make trades seemingly at odds with their immediate on-field success. One particular case is the White Flag Trade, which occurs when a team in playoff contention trades away star players mid-season, seemingly surrendering (or waving the white flag, hence the name). In general, trades such as these receive negative press and would be expected to have a negative effect on the team's standing with its fans, leading to fewer ticket sales. In this paper, the effect of these White Flag trades is tested to determine whether they affect individual-game attendance.

We have two examples of this type of trade to analyze. In 1997, the Chicago White Sox traded three major players (Wilson Alvarez, Danny Darwin and Roberto Hernandez) to the San Francisco Giants for six minor leaguers in a trade that coined the term White Flag trade. More recently, in 2008, the Oakland A's traded Rich Harden and Chad Gaudin to the Cubs for four young players. In both of these cases the teams were in contention for the playoffs at the time of the trade. The White Sox record at the time of their trade was only 53-53, but they were only three games back in the AL Central. Similarly, the Oakland A's in 2008 were 49-41, five games back and in second place. Furthermore, they were even closer in the wild card race, trailing the Boston Red Sox by only 3.5 games.

Silver (2008) examined the White Sox trade in 1997, but also examined a trade by the 2002 Cleveland Indians who he describes as one game back. However, on the date of their trade they were actually 36-41, seven games back and in third place. He does not explain why he chooses to include this trade, and since this team had a losing record and was many games back, we exclude this Indians trade.

We estimate the impact of the White Sox and A's White Flag trades on the teams' individual game attendance. Silver (2008) used regression analysis to model projected annual per-game attendance in the years following the trades by the White Sox and by the Indians, which he then compared to the actual per-game attendance for the teams. His analysis of the White Sox does show a decline from expected attendance following the trade, but does not indicate the significance of the results or provide evidence that the decrease is due to the trade. Furthermore, there could also be an impact on attendance in the season in which the trade is made which an annual attendance model would miss, but that individual-game attendance analysis would account for.

MODEL

Presumably, a team will make a mid-season trade of a veteran for prospects if the benefits outweigh the costs. There seem to be two primary costs associated with making this type of trade: a lower probability of making the playoffs and lower attendance. This paper examines the attendance effects. As to the playoff impact, Krautmann and Ciecka (2009) estimate that making the playoffs has an expected value of $11 million.

Of course, the team also receives benefits from the trade. It gets to lower its costs by removing the salary of the traded players. It also will potentially improve its team in the future by receiving prospects who will help the team in subsequent seasons. Increasing the probability of making the playoffs in future years by a substantial amount could outweigh the cost of lowering the probability of making the playoffs this year even after discounting the future value.

Regardless of the reason for the trade, our purpose is to examine the impact on attendance. We therefore model the attendance of a team:

$$\textit{Attendance} = \textit{Constant} + \beta_1 * \textit{Measure of Team Quality} + \beta_2 * \textit{White Flag Trade} + \beta_3 * X + \varepsilon \qquad (1)$$

An important determinant of attendance is team quality; we account for this factor by including a measure of how the team is doing during the year. However, even after controlling for team quality, a White Flag trade could cause attendance to drop because fans expect the team to do worse. We therefore include a White Flag trade dummy variable for all of the games following the trade for the remainder of the season.

X includes other standard factors that affect a team’s attendance. Dummy variables are included for the year, the month, the day of the week, and the time of day (night or day). A variable for opening day has also been included due to the atypically large showing seen by most teams. To account for potentially larger than normal crowds we include indicator variables for Memorial Day, Labor Day and Independence Day. For both White Sox and A’s attendance, the Augmented Dickey Fuller test rejects the presence of a unit root so we estimate the model as an ARMA (1,1).

One additional concern might be that the player who is traded away may be a superstar. Even if the returning players were providing equal value, the attendance may drop because the traded player is likely to be better known and thus more marketable. Previous studies have shown some effect from star players on attendance at sporting events (Berri, Schmidt and Brook, 2004; Brandes, Franck and Nuesch, 2008;

Mullin and Dunn, 2002; Treme and Allen, 2011). However, this concern should not apply to the two white flag trades we consider; while all of the veteran players traded away were key contributors, none of them were considered superstars.

DATA

The data come from Retrosheet.org. For each team we examine attendance for five years of home games: the year of the trade, the preceding two years and the following two years.

The choice of team success variable warrants some discussion. Win percentage is often used when looking at yearly attendance, but varies drastically early in the season. As an alternative we use Davis's (2009) approach of games above or below .500 which provides a steady movement up and down. As a robustness check we re-estimate the model with winning percentage. As can be seen in the A's results below, winning percentage does not account for team success very well. We therefore include a third model with a modified winning percentage. Meehan et al. (2007) dropped the first ten observations, but we do not have enough data to drop observations, so our similar solution is to set the winning percentage equal to .500 for the first twenty games and then set it at the actual value afterward. This adjustment should help to alleviate the early season fluctuations.

Summary statistics for the key variables for the White Sox are presented in Table 1. The White Flag Trade variable is for the games after the trade for the remainder of the 1997 season. However, since year-indicator variables are included, the variable is functionally the same as an indicator variable for all games after the trade. Attendance and Games Above .500 are shown for the full sample, as well as for samples representing before and after the trade (within the year and the following years). Table 2 presents similar data for the A's sample. Both teams have lower average attendance after the trades than before them.

Table 1: Summary Statistics for White Sox

Variable	Sample	Obs	Mean	Std. Dev.	Min	Max
White Flag Trade	Full	383	.0679	NA	0	1
Attendance	Full	383	20551	6840	746	44249
Games Above .500	Full	383	-3.731	8.408	-18	19
Attendance	Pre-Trade	202	22407	6260	746	44249
Games Above .500	Pre-Trade	202	-1.119	9.312	-17	19
Attendance	Post-Trade	181	18479	6880	8980	44153
Games Above .500	Post-Trade	193	-6.646	6.083	-18	4

Table 2: Summary Statistics for A’s

Variable	Sample	Obs	Mean	Std. Dev.	Min	Max
White Flag	Full	403	.0769	NA	0	1
Attendance	Full	403	20599	7300	8874	36067
Games Above .500	Full	403	-.6129	7.798	-17	27
Attendance	Pre-Trade	210	23124	7072	10128	36067
Games Above .500	Pre-Trade	210	3.286	7.495	-11	27
Attendance	Post-Trade	193	17852	6524	8874	36067
Games Above .500	Post-Trade	193	-4.855	6.113	-17	9

RESULTS

The results for the White Sox are presented in Table 3. The three columns are the same except for the method for controlling for team success. The estimated effect on the White Flag trade dummy variable is negative as hypothesized in all three specifications. However, even though the estimated effect is large—a decrease of more than 5% relative to mean attendance—it is not statistically different from zero in any of the three specifications.

There are a number of factors that might be influencing the results and causing the data to give insignificant results. First of all, the Games Above .500 variable for the White Sox is low relative to what was found by Davis (2009) in which the coefficient averaged about 228. The regression equations used here are not the same and the time period shorter than what was used by Davis, but failing to account for team success could bias the White Flag Trade variable, although it is unclear in which direction. The White Sox performed about as well after the trade (.491 winning percentage) as they did beforehand (.500 winning percentage). The alternative specifications for team success, winning percentage and winning percentage after 20 games, are not any more successful accounting for winning. They also did not have a substantial impact on any of the other coefficients.

Aside from the uncharacteristic Games Above .500 results, another factor that could be influencing the results is the baseball strike of 1994. With the cancellation of many major league games, including the World Series, many baseball fans were angry and likely not to attend as many games following the strike (see Schmidt and Berri, 2002, 2004). Since the data used in the White Sox analysis is immediately following this strike, it is possible that there is still an effect on attendance from the strike that is not accounted for by the year indicator variables. To attempt to reduce the impact of the strike, Table 4 presents the results of a re-estimation of the model with 1995 and 1999 dropped from the analysis. The results are essentially the same as the results from the full sample—the estimated coefficient on the White Flag trade variable is negative and large but not statistically significant.

Table 3: White Sox Results

	(1)		(2)		(3)	
Variable	**Coefficient**	**P-Value**	**Coefficient**	**P-Value**	**Coefficient**	**P-Value**
White Flag	-1671.45	0.370	-1374.65	0.456	-1273.51	0.492
Games Above .500	124.99	0.178				
Winning %			5214.87	0.487		
Win % After 20					3896.61	0.747
Sunday	8220.71	<0.001	8252.81	<0.001	8179.38	<0.001
Monday	3413.01	<0.001	3450.32	<0.001	3422.07	<0.001
Wednesday	1259.56	0.137	1272.34	0.138	1274.33	0.134
Thursday	1117.28	0.282	1167.45	0.260	1154.79	0.267
Friday	3320.42	<0.001	3308.02	<0.001	3315.72	<0.001
Saturday	9748.37	<0.001	9767.91	<0.001	9726.73	<0.001
April	-9693.68	<0.001	-9227.10	<0.001	-9608.90	<0.001
May	-5968.69	<0.001	-5655.38	<0.001	-5831.19	<0.001
June	-2086.79	0.028	-1967.99	0.039	-1990.42	0.041
August	-2521.42	0.071	-2869.88	0.030	-2882.27	0.031
September	-4521.66	0.002	-4814.57	0.001	-4839.50	0.001
October	-7416.80	0.147	-7878.18	0.147	-7823.25	0.139
Night	2466.86	<0.001	2480.29	<0.001	2422.10	<0.001
1995	5151.42	<0.001	4924.08	0.001	4575.21	0.003
1996	2378.86	0.173	3645.94	0.005	3637.21	0.019
1997	7604.77	<0.001	7751.84	<0.001	7550.43	<0.001
1998	1220.36	0.431	665.42	0.636	512.67	0.724
Opening Day	19066.97	<0.001	19178.59	<0.001	19112.58	<0.001
Memorial Day	3761.34	0.757	3725.16	0.774	3767.48	0.762
Labor Day	-1163.00	0.717	-1005.16	0.748	-994.08	0.752
4th of July	-1770.76	0.379	-1674.39	0.416	-1676.56	0.414
Constant	15537.1	<0.001	12462.32	<0.001	13291.01	0.027
MA (1)	0.045	0.670	0.042	0.695	0.040	0.706
AR (1)	0.372	<0.001	0.375	<0.001	0.377	<0.001

Table 4: White Sox Results (1996-1998)

Variable	Coefficient	P-Value
White Flag	-1618.87	0.397
Games Above .500	135.14	0.241
Sunday	6727.81	<0.001
Monday	3556.16	0.001
Wednesday	1541.09	0.123
Thursday	1521.15	0.210
Friday	3389.69	0.001
Saturday	8899.45	<0.001
April	-8018.10	<0.001
May	-4177.93	0.026
June	784.81	0.553
August	-770.59	0.637
September	-2616.44	0.115
Night	1290.26	0.130
1996	972.62	0.724
1997	6221.06	<0.001
Opening Day	19468.33	<0.001
Memorial Day	2917.74	0.764
Labor Day	-1744.63	0.569
4^{th} of July	6725.56	0.312
Constant	16218.0	<0.001
MA (1)	0.048	0.772
AR (1)	0.377	0.005

Table 5 has estimation results for the Oakland A's. As with the White Sox, there is no statistically significant evidence that White Flag trades harm attendance; the estimated coefficients are actually positive in two of the three specifications. The results suggest that the A's simply did not experience a negative direct effect on attendance from their trade, though it is possible that they did have a negative impact from the trade, but that the impact is all accounted for through the team's performance. The A's saw their winning percentage drop from .544 before the trade to .366 after. Given the coefficient in Column 3, such a drop in winning percentage cost them about 6600 fans a game, and a similar impact from the Games Above .500, depending on the win total, can be deduced from Column (1).

Table 5: A's Results

	(1)		(2)		(3)	
Variable	**Coefficient**	**P-Value**	**Coefficient**	**P-Value**	**Coefficient**	**P-Value**
White Flag	1494.66	0.534	-444.23	0.855	1660.64	0.500
Games Above .500	215.88	0.007				
Winning %			718.97	0.894		
Win % After 20					37118.63	0.016
Sunday	7040.89	<0.001	6968.22	<0.001	6908.46	<0.001
Monday	-767.44	0.409	-756.05	0.421	-808.44	0.385
Wednesday	5955.74	<0.001	5981.08	<0.001	5873.51	<0.001
Thursday	58.58	0.965	-54.33	0.967	-114.65	0.932
Friday	3216.58	0.001	3181.11	0.001	3080.53	0.001
Saturday	7179.20	<0.001	7111.62	<0.001	7055.26	<0.001
April	-2176.68	0.121	-2120.44	0.159	-2182.64	0.115
May	-906.44	0.473	-857.54	0.517	-659.50	0.601
June	-489.23	0.699	-384.52	0.767	-398.50	0.754
August	310.55	0.826	131.78	0.927	131.04	0.925
September	-2210.39	0.157	-1810.37	0.259	-2230.63	0.150
October	-3963.25	0.903	-4050.89	0.873	-4948.19	0.845
Night	940.96	0.285	887.62	0.309	906.83	0.294
2006	5193.10	<0.001	6839.42	<0.001	5581.28	<0.001
2007	6179.81	<0.001	5915.74	<0.001	5816.16	<0.001
2008	1306.53	0.406	2227.89	0.169	644.85	0.707
2009	1816.20	0.241	-81.98	0.958	1932.36	0.717
Opening Day	17592.79	0.001	17562.96	<0.001	17641.71	<0.001
Memorial Day	4036.40	0.207	3482.20	0.289	4427.96	0.168
Labor Day	2318.97	0.851	2695.86	0.708	2465.00	0.793
4^{th} of July	-2346.16	0.728	-2312.68	0.721	-2466.28	0.719
Constant	14161.68	<0.001	13740.95	<0.001	-4208.89	0.591
MA (1)	-0.023	0.879	-0.044	0.753	-0.018	0.904
AR (1)	0.360	0.008	0.403	0.001	0.353	0.009

Most of the other parameters have expected coefficients. A standing promotion to have $2 tickets on Wednesdays likely explains the A's unusually large Wednesday coefficient (Saldivar, 2009). As with the White Sox, the holidays do not have a significant impact, which may be due to the rareness of these events in our samples. Since we have only five years included in the analysis and half the games are on the road, each holiday will only occur around two or three times during the sample.

CONCLUSION

The findings of this study suggest that the white flag trades had no significant impact on attendance for either the White Sox or the A's. The White Sox results are the closest to the expected result of lost attendance following the trade but are still not significant. There may be an indirect impact through reduced team success, at least for the A's. The A's saw a drop in team success following the trade, and lower attendance. However, it is not clear if the trade had an impact on the team or if it would have performed worse even without the trade. It may be that fans will not lessen their attendance just from the team signaling that they are not trying to win, but will stop attending if the team actually stops winning.

Possible explanations for the lack of an impact on attendance include the possibility that the negative impact on team quality is cancelled out by increased media attention from the trade. The fans may also be acting rationally and not expecting the teams to compete with or without the trades.

Lastly, a selection bias may exist in that teams who expected a large loss in attendance from these types of trades simply do not make them. This limitation should also make us reluctant to generalize too strongly to other teams considering trades. However, the addition of a second wild card in 2012 financially devalues some post-season appearances, since some playoff teams are no longer guaranteed even one home playoff game. A reduction of the financial incentive of contending could make more of these trades rational.

ENDNOTES

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Collector Preferences for Hall-of-Fame Baseball Player Picture Cards 1981-2010

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ABSTRACT

Better understanding consumer behavior in collector markets may provide evidence of the broader pervasiveness of attitudes towards race and ethnicity. This paper examines the effects of race and ethnicity on prices in guides published between 1982 and 2010 for cards picturing Hall of Fame baseball players between 1956 and 1980. Using a hedonic pricing model, we estimate 750 models, account for card scarcity, and control for performance with the career Wins Above Replacement statistic. We find evidence that race negatively affects the listed prices, and this effect persists after 1992, a period for which prior studies have not identified consumer discrimination.

INTRODUCTION

In this paper, we examine consumer preferences for players inducted into the National Baseball Hall of Fame and Museum (Hall of Fame) who appear on picture cards issued by the Topps Company for the baseball seasons between 1956 and 1980. A voluminous amount of prior research has reviewed racial and ethnic preferences of collectors of sports memorabilia. As a hobby, identification of collector preferences may inform how a society values some traits and allocates its resources. Almost all published empirical research concludes that there is little evidence of collector discrimination for players pictured on baseball cards by the early-1990s. Using a simple hedonic pricing model, we estimate 750 models to find evidence of collector discrimination, including since the 1990s.

The effects of discrimination are difficult to empirically detect. Becker (1971) identifies employer discrimination, employee discrimination, and customer discrimination as sources of prejudice which may result in differential treatment. Two individuals with identical characteristics (gender, age, experience, skills, etc.) may be negatively affected in their relative labor productivity by racial or ethnic characteristics. Employers may give in to their customers' preferences and pass on the higher costs of discrimination in the price that the customer pays. Whether these forms of discrimination are systemic or cultural, individuals negatively affected by discrimination may suffer from fewer opportunities and lower expected lifetime compensation.

Performance statistics for players in Major League Baseball (MLB) are widely available. Researchers have extensively studied racial and ethnic salary differentials in MLB player compensation controlling for performance. By the 1980s, their empirical models show statistically inconclusive or insignificant results for racial or ethnic discrimination.[ii]

Scholars have turned their attention to whether collectors exhibit a preference for racial or ethnic characteristics when making choices to purchase in the secondary market for sports player memorabilia. As the market for memorabilia is independent of the player and cannot affect the player's performance, the observed prices reflect collectors' preferences when controlling for players' performances. When consumers' attitudes towards race and ethnicity are reflected in their preferences when purchasing baseball memorabilia, they may affect mediums in which players appear and/or perform, including advertisements, magazine covers, ballots, and votes for the MLB All-Star Game and Hall of Fame inductions.[iii]

The retail baseball card market today is like that which existed between 1956 and 1980, when Topps was the sole nationwide distributor of baseball cards sold with a confectionary product.[iv] In the retail market, player cards are offered to collectors in a package including a known number of cards of unknown content, with an average per card retail price that ranges from about 15 cents to more than $1. The variability in retail prices should reflect the various costs incurred to produce, including licensing, design, marketing, manufacturing, etc. Once a card is produced and sold at retail its production cost is sunk and no longer relevant to its secondary market price. The forces of supply and demand from consumers determine a card's secondary market price.

The benefits of an active secondary market enable a consumer to assemble a collection through purchases which reflect the collector's preferences. This secondary market is attractive for study due to its numerous transactions and relative pricing transparency. Secondary market exchanges can occur via face-to-face, online, phone, text, or mail contact. Transactions may be observed and recorded at shows and auctions. COMC.com, eBay.com, and other websites and forums offer the means for anyone to offer his cards to everyone in an auction or sale format. Pricing information is readily available online, in at least one monthly magazine, and in at least one annual guide.[v]

This study focuses on prices between 1982 and 2010, a period during which baseball card collecting rose into a national fascination.[vi] During the 1980s, the hobby exploded into the mainstream with millions of collectors, thousands of dealers, perhaps a store in every large town, and multitudes of weekend conventions. The price appreciation of old baseball cards was widely observed in the newspapers and magazines.[vii] Eventually, the hobby exhibited characteristics of a mania, with books and magazines which offered advice to collectors for how to invest or speculate in baseball cards.[viii] These publications emphasized that long-term values could rise or fall, but their relative values were assured only for the cards picturing players who became enshrined in the Hall of Fame.[ix] Many of these views continue today, although the size of the retail and secondary markets has decreased since the 1980s. More recently, the spread of third-party grading and the introduction of registries which rank the average condition of a card collection have altered the way collectors assemble their collections.[x]

During the 1980s and 1990s, newspapers and magazines reported perceptions of collector biases for race and ethnicity when purchasing baseball memorabilia. Writers observed that serious baseball card collectors were generally white, male, middle age, and middle class.[xi] They prominently noted the price differentials observed between cards which pictured white and black players.[xii] As an explanation for collectors' racism, Bloom (1997) suggests that collectors of baseball memorabilia revealed their childhood experiences with race and gender. Therefore, secondary market card prices reflected the narrower culture and racial preferences of white, middle class males.

EVIDENCE FOR COLLECTOR DISCRIMINATION

Significant evidence of collector discrimination has been identified in price guides up to 1992, and very little after 1992. Nardinelli and Simon (1990) find evidence in prices from 1989 that consumers discriminate against Latin hitters and black pitchers in the 1970 Topps series. Anderson and La Croix (1991) use samples of the 1960-1961 and 1977 Topps series, and several years of price guides (1982, 1983, and 1985), to find consumer discrimination against black players but not against Hispanic players. Using a continuous measure for consumer perception of race or ethnicity, Fort and Gill (2000) evaluate the impact of race and ethnicity on prices from a 1987 guide for the 1987 Topps series, and conclude that consumers discriminate against African American hitters and pitchers, as well as Hispanic pitchers. Reviewing pricing for the 1960-1969 Topps series, Burnett and Van Scyoc (2004, 2013) and Van Scyoc and Burnett (2013) find evidence for collector discrimination against African American and Hispanic hitters in the 1992 guide, none in the 2008 guide, and evidence against African American hitters in the 1981 guide. They also find no evidence for collector discrimination in the 1986 series using the 2008 guide.[xiii] Scahill (2005) repeats the Nardinelli and Simon (1990) study for the 1970 Topps series using almost 20 years of price guides. Scahill observes that the race variable is negative and significant for some years, primarily between 1984 and 1991, but between 1992 and 2001 Scahill identifies no evidence that race significantly affects prices.

Other studies review the impact of race and ethnicity for memorabilia and rookie cards. Reviewing memorabilia prices in a guide from 1995-1996, Sharpe and La Croix (2001) conclude there is not sufficient evidence that collectors discriminate.[xiv] Reviewing April 1992 prices and race/ethnicity for players with Topps rookie cards issued between 1984 and 1990, Gabriel, Johnson and Stanton (1995) observe no significant evidence of customer discrimination. Burge and Zillante (2014) find that for cards issued between 1986 and 1993, African American players have an estimated 15 to 19 percent premium above that of comparable white ones.[xv]

Several studies examine Hall of Fame player rookie card prices. A sample with 1990 prices shows a very weak impact of race and ethnicity (Regoli, 1991). Another sample across a longer time period does not identify racial or ethnic discrimination (Regoli, 2000). Regoli (2000) concludes that collectors evaluate white Hall of Fame players by their career statistics but evaluate African American players for their more subjective Hall of Fame career achievement. However, Hewitt, Munoz, Oliver, and Regoli (2005) find a Hall of Fame player's rookie card price is not affected by his race, when controlling for career performance and card scarcity. Replicating the Hewitt, Munoz, Oliver, and Regoli (2005) study with corrected data, Findlay and Santos (2012) confirm the conclusion by Hewitt, Munoz, Oliver, and Regoli (2005), then include ethnicity and find it does not affect prices. Primm, Piquero, Regoli, and Piquero (2010) collect a sample of rookie cards for 66 players who were elected into the Hall of Fame via a vote conducted by the Baseball Writers' Association of America (BBWAA). They find that the April 2003 price is affected by performance and card availability but neither race nor ethnicity. Primm, Piquero, Piquero, and Regoli (2011) review rookie card prices for those players who received at least one vote in the balloting for induction into the Hall of Fame. They find that race rather than skin tone affects price.

THE SAMPLES AND VARIABLES

Prior studies generally use all cards issued in a single year or examine star cards from several years or for several manufacturers.[xvi] We record a sample of players inducted into the Hall of Fame for 25 Topps series and review almost 30 years of price data from a single annual source, rather than a sample from a single series, rookie cards from multiple series, or prices from a single period. Instead of several different performance measures, our specification includes a single career performance variable, Wins Above Replacement (WAR), which allows us to combine the player sample within each series. We explain further below.

This study focuses on Hall of Fame baseball players because enshrinement is an independent stamp of approval.[xvii] Compared to almost all other baseball cards, Hall of Fame player cards are the most collectible and relatively liquid.[xviii] Markets for cards of active players are volatile. Collectors evaluate recent events and project current and future performance to forecast career statistics sufficient for consideration to be honored within the Hall of Fame. Collector perceptions and expectations about career performance can drastically change as the player ages, changes teams, suffers an injury, etc. Because retired players' career statistics are well-known and final, the impact on prices due to speculative demand is reduced. As of 2019, 232 former MLB players are enshrined.

Topps has produced baseball cards series continuously since 1951. Following its purchase of its competitor Bowman, it had a monopoly on the distribution of licensed cards distributed with gum between 1956 and 1980.[xix] In this study, we focus on the years of issue between 1956 and 1980, when no nationally distributed substitutes were available.

The Official Price Guide to Baseball Cards was a comprehensive list published annually beginning in 1981 and ceasing after 2010.[xx] These guides state their prices reflect actual market pricing from observed transactions. Guides are popular because they reduce collector uncertainty regarding prices.[xxi] We record Hall of Fame baseball player card prices for the highest grade cards, beginning with the 1982 guide and concluding with the 2010 guide. Rather than select a specific Topps series year, we use an extensive range, 1956 through 1980.[xxii] Pricing information is recorded for those Hall of Fame baseball players on single image, regular Topps manufactured gum cards, and appear on a Topps card prior to the 1981 season and for the years after the players are inducted.

The sample consists of regular series cards picturing Hall of Fame players on their non-rookie cards issued by Topps between 1956 and 1980, using the guides between the years 1982 and 2010 for the years following when the player was enshrined. The initial issued Topps card – the so-called “rookie card” – is not included, as some collectors specialize in collecting Hall of Fame player rookie cards.[xxiii] Players who do not appear as an active player on a Topps card after 1955 are not included within the sample. Cards picturing managers, executives, umpires, coaches, and other non-players are not in the sample. Cards picturing All-Stars, statistical leaders, groups, milestone events, or a team are not included in the sample.

For example, Mike Schmidt is a member of the 1995 Hall of Fame class and first appeared on a Topps card in 1973. For our study, we include the Topps issues picturing Schmidt between 1974 and 1980, exclude card prices from the 1995 and previous years’ samples, and include the prices for these cards in the 1996 and following years’ samples. Alternatively, Ralph Kiner is last pictured in a 1955 Topps card (before 1956) and was inducted with the 1975 class. No cards picturing Kiner are in the samples. Ron Santo and Ted Simmons are examples of players active sometime between 1956 and 1980 but are not in our samples. Santo is first pictured on a Topps card in 1961 but is a member of the 2012 class (after 2010), while Simmons is first pictured on a Topps card in 1971 and a member of the 2020 class. Similarly, Harold Baines and Lee Smith are not in our samples as members of the class of 2019 (after 2010), and they is first pictured on a baseball card in 1981 and 1982 (after 1980).

Given these restrictions, Table A in the Appendix shows the 74 players – 49 hitters and 25 pitchers – who are possible within the yearly samples. These are players who were active at some point between 1956 and 1980. We record approximately 15,000 prices between 1982 and 2010 for the Topps cards issued between 1956 and 1980.

The observed card price is transformed into a premium, the ratio of the player’s card price to the common card series price. Consider two greats, Willie Mays and Mickey Mantle, who played center field for New York City teams during baseball’s golden age. In the 1991 price guide, the classic 1957 card for Mr. Mays was priced at $180 while Mr. Mantle’s card was $750. Topps printed the 1957 picture cards in separate series and sold each series separately. Series printed early in 1957 were distributed early in 1957 while series printed later in 1957 were distributed later in 1957. Often, series were printed in different quantities. The 1957 series is listed as printed in five series. Mays is pictured on card number 10, and a common card in the series which includes Mays is listed at $5. Mantle is pictured on card number 95, a different series, and a common card in that series is listed at $4. For Mays, the premium is $180 divided by $5 or 36.00. The Mantle card’s premium is 187.50. Since Hall of Fame players are the all-time greats, their picture cards are priced higher than common cards and their premiums exceed 1. If the listed common price reflects relative series scarcity, this ratio controls for scarcity differences within and between issues.

The different premiums for different Hall of Fame players reveal collectors have different demands for these players. These differences in premiums can be explained by measures of productivity or quality of players’ careers. The measures include WAR and the number of Most Valuable Player (MVP) and Cy Young awards.[xxiv] Rather than batting average, slugging percentage, on base percentage, runs batted in, wins, strikeouts, earned runs average, etc., we use the single performance variable career WAR. The interpretation for WAR is the number of wins that the player contributes to his club above that of a replacement player, generally viewed to be like the team's AAA player at the appropriate position. The meaning of WAR is similar to the economic concept opportunity cost. The replacement player is the relevant comparison because the average player is already under contract to an MLB club. To give up the MLB player for a replacement player means that the club would lose the MLB player's contribution to its wins. The statistic is computed differently for pitchers and position players, but the interpretation is the same.

To identify differences in collector preferences between position players and pitchers, as well as to account for some differences in measurement, we use a dummy variable taking the value of 1 if the player is a pitcher and 0 otherwise.

We measure collector preferences for player picture cards by including a dummy variable to identify the race and ethnicity of the player. For the race variable, we use Ashe (1993) to identify African-American players. The assignment of race and ethnicity is provided in Table A in the Appendix. If the player is African-American, the variable takes the value of 1. For the ethnic variable, the variable takes the value of 1 if the player is a native of a Latin-American country or has a Latin surname.

Table 1 lists the summary measures of the entire sample. Of the 74 players in the sample, 21 are African American, 6 are ethnically Hispanic, and 47 are White. A total of 72 MVP or Cy Young awards were received by the players in the sample. The WAR statistic has a maximum value of 156.2 (Willie Mays) and a minimum of 21.3 (Monte Irvin), a sample mean of 71.39, and a sample median of 66.50.

Table 1: Summary Statistics

	African-American	**Hispanic**	**White**	**Pitcher**	**Awards**	**Career WAR**
Observations	21	6	47	25	72	74
Maximum	1	1	1	1	4	156.20
Minimum	0	0	0	0	0	21.30
Mean	0.28	0.08	0.64	0.34	0.96	71.52
Median	0.00	0.00	1.00	0.00	1.00	67.20
Standard Deviation	0.45	0.27	0.48	0.48	1.05	26.86

Note: The 74 players in the sample are listed in Table A in the Appendix. See the text for explanation of the variables.

Table 2: Summary Statistics by Race and Ethnicity, All Players and All Position Players

Measure	Maximum	Minimum	Mean	Median	Standard Deviation
			WAR		
All Players (74)	156.20	21.30	71.52	67.20	26.86
White (47)	128.10	24.50	70.31	67.20	24.83
African American (21)	156.20	21.30	75.67	67.40	33.30
Hispanic (6)	94.50	50.30	66.43	59.40	17.59
Position players (49)	156.20	21.30	73.70	67.80	28.66
White (25)	128.10	36.20	73.24	69.05	24.60
African American (19)	156.20	21.30	74.44	64.40	34.85
Hispanic (5)	94.50	50.30	67.10	55.70	19.58
			Awards		
All Players (74)	4	0	0.96	1	1.05
White (47)	4	0	1.02	1	1.15
African American (21)	3	0	0.95	1	0.92
Hispanic (6)	1	0	0.50	0.5	0.55
Position players (49)	3	0	0.90	1	0.98
White (25)	3	0	0.93	1	1.05
African American (19)	3	0	0.84	1	0.90
Hispanic (5)	1	0	0.60	1	0.55

Table 2 provides a refined review of the summary measures for WAR and Awards, by race and ethnicity, for all 74 Hall of Fame players, and for the 49 position players (non-pitchers). For all 74 players, African American players on average have WARs above the mean while Hispanic and White players on average have WARs below the mean. The same pattern holds when looking only at the 49 position players (non-pitchers) or when looking at the median WAR rather than mean WAR.

METHODOLOGY

Nardinelli and Simon (1990) author the initial published empirical research for the impact of race or ethnicity on baseball card prices. In contrast to measuring baseball player behaviors, they noted measuring consumer behavior for cards is relatively easy because prices are recorded in guides. Their model connects consumers' entertainment preferences for a card with the secondary market for cards and links consumer racial attitudes to prices. They assume player productivity and consumer discrimination are independent. A player's productivity should be unaffected by the secondary market for cards, and therefore any collector discrimination for cards should not affect their prices.

When consumers receive satisfaction from players' baseball cards, their preferences affect secondary market pricing. The level of satisfaction is a linear function of players' career performances. As a player's career performance increases, all else equal, a collector's utility increases when acquiring the player's card. As the collector's utility increases, the collector's willingness to buy increases.

As each series of baseball cards has a common card price, cards of players with higher career productivity are priced higher, and the card price is the sum of the common price and the premium. The premium for a specific player's card is the difference between a player's card price and the common card price. The premium is determined by the maximum value of either 0 or the level of collectors' satisfaction due to the player's career productivity. By transitivity, the price difference is equal to this maximum value. When the lower bound is not truncated, the price difference may be estimated by ordinary least squares (Gabriel, Johnson, and Stanton, 1995, 221).[xxv] All else equal, a preference for an attribute of a baseball player is expected to increase his card price and premium. If consumers discriminate, a player's race or ethnicity will be a significant determinant of the premium, and the estimated coefficients on the racial or ethnic variables will be statistically significant.

Card prices are affected by performance. Nardinelli and Simon (1990) include measures for hits, doubles, triples, home runs, walks, stolen bases, at bats, seasons, post season games, and position, but not runs batted in (RBI). Other studies include RBI. We use the single career performance measure, WAR. Hewitt, Munoz, Oliver, and Regoli (2005) and Findlay and Santos (2012) also use single measures for performance which are similar to WAR.

For each series, the price model controls for player productivity using the career numbers of WAR and the number of MVP or Cy Young awards. All else equal, player productivity and price are expected to be positively correlated. A dummy variable to identify race and ethnicity of the player is included in the model. If the estimated coefficient for the race/ethnicity variable is statistically significant, then price differences may be attributed to discrimination. The model to be estimated is:

$$\ln\frac{P_{it}}{P_{ct}} = \beta_0 + \beta_1 AA_{i_i} + \beta_2 \text{Hispanic}_i + \beta_3 \text{Pitcher}_i + \beta_4 \text{Awards}_i + \beta_5 \ln WAR_i + \varepsilon_{it}$$

Where $\ln\frac{P_{it}}{P_{ct}}$ is the natural logarithm of the ratio of player *i*'s card price to common price, during price guide year *t*, β_0 is a constant, AA_i takes the value 1 if player *i* is African American, Hispanic_i takes the value 1 if player *i* is from a Spanish speaking country or has a Latin surname, Awards_i is the number of MVP or Cy Young awards received by player *i*, Pitcher_i takes the value 1 if player *i* is a pitcher, $\ln WAR_i$ is the natural logarithm of player *i*'s WAR, ε_{it} is a stochastic error term assumed to be normally distributed with mean 0 and known variance 1, for each player *i* in series from guide year t, and β_j, $j = 0 \ldots 5$, are the regression parameters to be estimated.

Since the sample sizes are small, typically not more than 30 observations, and the degrees of freedom are scarce, a limited number of independent variables is desirable. A single performance variable such as WAR offers a parsimonious specification.[xxvi]

ESTIMATED COEFFICIENTS

Table 3 summarizes the numbers of regressions providing estimates of the model equation. A sample includes the cards of Hall of Fame players in a Topps series for a single price guide year. We estimated 656 samples which had enough degrees of freedom using ordinary least squares estimation. Of those 656 542 estimated models, almost 82 percent, are statistically significant using an F-test at the 10 percent level of significance. In 429 models, the estimated coefficients of WAR for the WAR variable are statistically significant, as well as for the estimated coefficients for the Awards variable in 419 models. Few Hispanic variables, 5, have a statistically significant estimate for the coefficient.

In 104 estimated models, the estimated coefficient for the African-American variable – 19.19 percent of statistically significant models – is statistically significant and negative. While these results are meaningful, the effect is not large. The estimated impact for price due to an African American player picture on the card is expected to be about two cents. Across a sample of 100,000 cards, all else equal, all collectors impute a total reduction of $2,000 in the value of this card compared to a card picturing a White player. If there are ten African American players pictured in a year's series for Topps and in the Hall of Fame, then the cumulative impact is an estimated reduction of $20,000.

Overall, as seen in Table 3, roughly four of five estimated models are at least statistically significant at the 10 percent level of significance; of these, almost one in five of these have a statistically significant and negative sign on the coefficient for the African-American variable.

Table 3: All Topps Series, Numbers of Estimated Regressions, Statistically Significant Regressions, and Statistically Significant Regression Coefficients, 1982-2010 Price Guides

Topps Series Year	Regressions		Negative Coefficients		Positive Coefficients	
	Total Estimated	**Statistically Significant**	**African-American**	**Hispanic**	**WAR**	**Awards**
1956	28	28	0	0	28	28
1957	29	29	2	0	29	29
1958	29	29	21	0	29	29
1959	29	29	0	0	27	29
1960	29	27	0	0	27	27
1961	29	23	3	0	23	14
1962	29	29	0	0	29	28
1963	29	28	0	0	28	27
1964	29	27	0	0	14	27
1965	29	28	0	0	20	27
1966	29	29	25	0	4	29
1967	29	27	13	0	26	23
1968	29	28	11	0	28	27
1969	28	26	19	0	18	23
1970	28	12	1	0	12	6
1971	28	27	1	1	24	7
1972	27	26	0	0	20	17
1973	27	22	1	0	13	4
1974	26	15	0	0	12	3
1975	19	8	0	0	3	4
1976	19	14	0	0	11	1
1977	19	10	1	0	2	5
1978	19	8	2	0	2	1
1979	19	6	3	0	0	3
1980	19	7	1	4	0	1
Total	656	542	104	5	429	419
Percent Regressions	100.00	82.62	15.85	0.76	65.40	63.87
Percent Significant Regressions		100.00	19.19	0.92	79.15	77.31

Note: Level of significance at 10 percent

EMPIRICAL EVIDENCE OF CONSUMER DISCRIMINATION

For the Topps series between 1982 and 2010, our estimated models overwhelmingly are statistically significant using a career performance measure of WAR and an achievement measure for numbers of MVP and Cy Young awards. We have identified evidence of consumer discrimination against cards picturing African American Hall of Fame players, in almost one in five statistically significant estimated models.

To better understand the performance of the model and the prevalence of collector discrimination, in Table 4, we combine the one-year sample periods into five-year ranges and summarize. More than 74.81 percent of the models estimate a significant regression prior to the 1976-1980 Topps series, with the highest estimation significance – 99.31 percent – observed within 1956-1960. Less than half, 47.37 percent, of the models estimated are statistically significant within 1976-1980. Based on the model estimation, the WAR statistic generally does not explain pricing within 1976-1980.

Table 4: All Topps Series in Five-Years Range, Percentage of Total Statistically Significant Regressions and Coefficients, 1982-2010 Price Guides

Years of Topps Series	Statistically Significant Regressions	Statistically Significant Coefficients			
		African-American	Hispanic	WAR	Awards
1956-1960	99.31	16.08	0	97.90	99.30
1961-1965	93.10	2.22	0	84.44	91.11
1966-1970	85.31	56.56	0	72.13	88.52
1971-1975	74.81	2.04	1.02	73.47	35.71
1976-1980	47.37	20.09	8.89	33.33	24.44

Note: Level of significance at 10 percent

Roughly one in six statistically significant estimated models has a statistically significant African-American coefficient that has a negative sign in the five-year Topps series between 1956 and 1960, and about one in five between 1976 and 1980. For whatever the reason, the 1958 Topps series appears to illustrate collector discrimination. More than half (56.56 percent) of the models in the 1966-1970 range are shown to have a statistically significant coefficient with a negative sign. To further examine our evidence, Table 5 maps the 104 statistically significant and negatively signed estimated African-American coefficients.

Table 5: Statistically Significant African American Regression Coefficient, 1982-2010 Price Guides, 1956-1980 Topps Series

Year	57T	58T	61T	66T	67T	68T	69T	70T	71T	73T	77T	78T	79T	80T	Total
1982															
1983															
1984															
1985							X			X					2
1986				X											1
1987		X		X											2
1988		X		X											2
1989		X		X											2
1990		X		X	X		X								4
1991		X		X			X	X	X						5
1992		X		X			X				X	X	X		6
1993		X		X			X						X	X	5
1994				X			X								2
1995		X		X			X								3
1996				X			X								2
1997				X			X								2
1998		X		X	X		X								4
1999	X	X		X	X	X	X						X		7
2000	X	X		X	X		X								5
2001		X	X	X		X	X								5
2002		X		X	X	X									4
2003		X		X	X	X									4
2004		X		X	X	X									4
2005		X		X	X	X	X								5
2006		X		X	X	X	X								5
2007		X		X	X	X	X								5
2008		X		X	X	X	X								5
2009		X	X	X	X	X	X								6
2010		X	X	X	X	X	X					X			7
Total	2	21	3	25	13	11	19	1	1	1	1	2	3	1	104

Note: Level of significance at 10 percent

Table 5 shows collector discrimination appears particularly in the Topps series for 1966 and 1969 and prior to the price guide years of 1999. For the Topps series of 1966, 1967, 1968, and 1969, collector discrimination is especially seen following the price guide years following 1999. We have identified evidence of collector discrimination. Between 1956 and 1980, 14 of 25 Topps series (as shown) have at least one statistically significant negative coefficient for the African-American variable. By 1990, the coefficient is significant and negative for more than two Topps series.

Scahill (2005) identified 1992 as the beginning of the period when race or ethnicity have no significant effect on prices. We have seen differently. For the price guides published between 1992 and 2010, Tables 6 and 7 repeat the summarizations shown in Tables 3 and 4, beginning with the 1992 price guides. We limit the information shown in Table 6 for the years with statistically significant coefficients for race or ethnicity.[xxvii]

Table 6 shows that the model estimates with statistical significance with a slightly higher proportion (83.79 percent to 82.27), with a higher proportion of WAR coefficients with statistical significance (82.91 percent to 79.01). Similarly, a higher proportion of estimated coefficients with a negative sign is observed for the African-American variable (21.36 percent to 19.15).

Table 6: All Topps Series, Numbers of Estimated Regressions, Statistically Significant Regressions, and Statistically Significant Regression Coefficients, 1992-2010 Price Guides

Topps Series Year	Regressions		Negative Coefficients		Positive Coefficients	
	Total Estimated	Statistically Significant	African-American	Hispanic	WAR	Awards
1957	19	19	2	0	19	19
1958	19	19	16	0	18	19
1961	19	19	3	0	19	14
1966	19	19	19	0	19	19
1967	19	19	12	0	19	19
1968	19	19	11	0	19	19
1969	19	19	16	0	16	19
1977	19	10	1	0	2	5
1978	19	8	1	0	2	1
1979	19	6	3	0	0	3
1980	19	7	1	4	0	1
Total	475	399	85	4	330	282
Percent Regressions	100.00	84.00	17.89	0.84	69.47	59.37
Percent Significant Regressions		100.00	21.30	1.00	82.71	70.68

Note: Level of significance at 10 percent

That the model estimates statistical significance of collector discrimination with greater frequency beginning with the 1992 guide is confirmed in Table 7. Notably, compared to Table 4, proportionally more estimated African American coefficients are statistically significant for the five-year periods 1956-1960, 1966-1970, and 1976-1980 Topps series, but fewer (or the same) are statistically significant for 1961-1965 and 1971-1975. Contrary to conclusion in Scahill (2005), we find evidence of collector discrimination after 1992, particularly in the Topps series for 1958, 1966, 1967, 1968, and 1969. Table B, in the Appendix, lists the Hall of Fame players pictured on card appearing in a Topps series between 1966 and 1970.[xxviii]

Table 7: All Topps Series in Five-Years Range, Regressions Percentage Statistically Significant, and Statistically Significant Coefficients, 1992-2010 Price Guides

Years of Topps Series	Regressions	Coefficients			
		African-American	Hispanic	WAR	Awards
1956-1960	100	18.95	0	98.95	100
1961-1965	100	3.16	0	94.74	94.74
1966-1970	94.74	64.44	0	91.11	91.11
1971-1975	75.79	0	0	72.22	47.22
1976-1980	47.37	13.33	8.89	26.67	24.44

Note: Level of significance at 10 percent

CONCLUDING OBSERVATIONS

The simple hedonic pricing model statistically explains the variation in Hall of Fame player card prices in Topps series between 1956 and 1980. By limiting the performance variable to WAR, we increase sample sizes and the degrees of freedom available in estimation. There is buyer discrimination for some card series but not others.

This study provides evidence that since 1992, until at least 2010, collectors sometimes discriminate against African American players who have been inducted in the Baseball Hall of Fame and appear on Topps picture cards issued in series between 1956 and 1980. The comprehensive identification of this persistence since the early 1990s differs from most published scholarship which study prices and collector perceptions of race and ethnicity. For many Topps card series, particularly 1958, 1966, 1967, 1968, and 1969, this discrimination means collectors have less apparent interest in Topps cards picturing an African American Hall of Fame player which results in a lower price, compared to collectors who have more interest in purchasing a Topps card picturing a Hall of Fame White or Hispanic player at a higher price. While the effect on an individual card price is very small, the cumulative impact for the total surviving examples of the card may be thousands of dollars; when multiplied by the number of affected players in a series, and then by the number of series, the total effect may be considerably large. Such an estimated cumulative reduction in expected value by collectors could reflect a broader monetary devaluation for identical social contributions made by affected groups.

One could also speculate why this discrimination persists from the 1990s into the new millennium. Collecting picture cards may be a hobby no longer enjoyed by large numbers of young persons.[xxix] The current collector base may be aging and thus card prices most likely reflect the preferences of this specific demographic rather than those of the broader population. As the relative size of those born prior to 1945 shrinks, the relatively few collectors from the Millennial Generation could result in a concentration of collectors from the Baby Boom Generation. Collectors from this generation may be less aware of their biases and unconsciously prefer players who look more like themselves.

Other possibilities may relate to our specification or changes in how the game is played. Our model may be too simple. It may, for instance, omit key explanatory variables or important control variables, such as a likeability Q-Score. The style and strategy of play has changed during the periods during the baseball card series which affect career, and play has also changed during the price guide years. These changes could affect consumer perceptions of prior performance when compared with recent performance. The end of the recent steroid era coincident with "money ball," the development and adoption of new performance measures, and broader acceptance of sabermetrics may reflect a transition which reflects changes in consumer preference for the game's style and strategy. As a result, collector preferences may be in transition and the market be in adjustment during our sample periods.

Nonetheless, our results reflect that of another study for the BBWAA ballots and racism. Jewell, Brown and Miles (2002) find discrimination against retired African American and Hispanic players. For whatever reason, our evidence shows this discrimination does not end for those players who are able to achieve election. We identify more recent evidence of collector discrimination against African American players who are enshrined at the Hall of Fame and appear on cards issued by Topps between 1956 and 1980.

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◊ I thank Andrew Turner (SUNY Oneonta alumnus, 2010 B.S. Economics) for assisting me with the collection of the price data between 1994-2008. Any errors in the data recorded and used remain my own.

END NOTES

1. This paper is based upon McAvoy (2019). "WAR, Race, and Ethnicity: Collector Discrimination for Hall-of-Fame Player Baseball Cards." This paper was first presented at the Twenty-Ninth Cooperstown Symposium on Baseball and American Culture, 2017. It was selected for publication in *The Cooperstown Symposium on Baseball and American Culture, 2017-2018* anthology by McFarland & Company, Publishers, Inc. This paper further extends and revises McAvoy (2013).
2. For an extensive literature review up to the time of its publication, see Kahn (1992).

3. Minority representation on the covers of *Sports Illustrated* has increased, but less than the increase in minority participation rates (Primm, DuBois, and Regoli, 2007). When assigning numbers, Topps did not consider race (Regoli, Primm, and Hewitt, 2007).
4. MLB granted Topps the exclusive rights to manufacture and distribute series of cards through 2025 (Mueller, 2018). Topps has had this exclusive right since 2011.
5. Professional Sports Authenticator, a prominent third party grading company, maintains an online price guide and records of prior auction sales. See "SMR Online" and "Auction Prices Realized," psacard.com (site accessed August 1, 2018).
6. Jamieson (2010) provides a comprehensive history of the hobby from its rise to national prominence to its decline and fall.
7. This continues to be true (Seideman, 2018).
8. Examples include Chadwick and Ray (1989), Green and Pearlman (1989), Rosen (1991), and Stewart (1993).
9. See Green and Pearlman (1989), Kirk (1990), and Stewart (1993). Election by the Baseball Writers' Association of America is considered most prestigious.
10. Interestingly, these trends have been observed in other hobbies such as coins and stamps. It is easy to use a smart phone and apps to quickly compare what is offered online and view the prices recorded in recent transactions. Third party grading and registries have resulted in very high prices for top condition scarcities.
11. Chadwick and Ray (1989) claim 99.5 percent of collectors were white in the late 1980s.
12. Rosenblatt (1990) provides a very good example of the public perception. Green and Pearlman (1989) and Chadwick and Ray (1989) provide hobby writers' perceptions. Rosen (1991) provides a dealer's view. Kiefer gives the hobby editor's view (1988). Regoli (1991, 2000) provides academic views.
13. They do not adjust for series scarcity.
14. Sharpe and La Croix (2001) find evidence collectors prefer memorabilia of African American baseball players.
15. The premium is the difference between the prices at initial release and 17 years later (Burge and Zillante, 2014).
16. Scahill (2005) is an exception and uses guides from 1979-1984, 1986-1997, and 1999-2001.
17. At present, no player active from 1956-1980 appears on the BBWAA ballot. The Veterans Committee now selects players from our sample period. Jack Morris and Alan Trammell, in 2018, were the most recent players inducted from this era. For information for how players are selected by the Veterans Committee, see the Baseball Hall of Fame and Museum, "Eras Committees," baseballhall.org (site accessed August 1, 2018).
18. Liquidity refers to the time and cost to convert a baseball card into money. As collectibles, baseball cards are not as liquid as financial instruments (checking account balances, stocks, bonds, etc.).

19. Topps defended its licenses, using courts against companies which infringed its rights. In 1981 the courts held for the Fleer Corporation which had contested Topps' claims to its rights. Besides Fleer, other issues or manufacturers produced baseball memorabilia between 1956 and 1980, including pictures on cereal boxes, bread labels, bottle caps, plastic cups, statues, cans, and others. The Topps series differed in terms of the size of each series, showing nearly every player who appeared on a team roster. Topps purchased licenses from the MLB and MLBPA, and it contracted individually with players for the rights to distribute their pictures. Only Topps offered a near complete series of picture cards during the season which showed the player in his team uniform and with his club name.
20. Hudgeons authors the 1981-1985 guides, and Beckett authors them after 1985. These pocket guides began their publication in 1981 and ceased in 2010. eBay, google.com searches, and online subscription databases have largely replaced the printed price guide.
21. Beckett attributes the success of his guide because it, "… is complete, current, and valid. … the most accurate, with integrity." See, "How to Use This Book" (Beckett, 1993). Stewart (1993) describes how the price guide developed into an authoritative resource.
22. Beckett (1985-2010). The price recorded is based upon the title cover year. The 2002-2003 22nd edition is excluded since it has a 2003 copyright. Prior to the 22nd, the copyright year preceded the title year by one year. After the 22nd, the copyright year matched the title year.
23. Excluding these rookie cards removes the difficulty of determining the rookie premiums identified by Burge and Zillante (2014).
24. WAR and numbers of awards are recorded from <baseball-reference.com>. We use the WAR 2.2 measure. Alternative measures of WAR are available, but all are understood to have similar relative player career rankings.
25. The price difference may be estimated by a Tobit technique, if the lower bound is truncated (Nardinelli and Simon, 1990). Hall of Fame player cards are priced above the common price. All sample prices have a multiple greater than 1 and estimated coefficients of the performance variables are positive. We proceed with the ordinary least squares estimation method.
26. Because of small sample sizes Anderson and La Croix (1991) delete insignificant variables when neither fit nor estimated ethnic coefficients are significantly affected. We remove neither explanatory nor control variables.
27. Available from the author upon request.
28. I thank an anonymous referee for observing that 1966-1970 may be perceived to be an era of heightened racial sensitivity and suggesting that some of these players may be perceived by collectors as controversial.
29. The Millennial Generation has a reputation for enjoying consumption of experiences rather than things.

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Determinants of Division I NCAA Soccer Participation

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ABSTRACT

Previous research has shown the importance of industrialization and immigration in the sport of men's soccer in the early 20th century. We test economic determinants for soccer and offer new evidence on the relationship between socioeconomic factors and labor supply for NCAA soccer participation. Exploiting a sample of 30,935 Division 1 college athletes, and the counties where these athletes attended high school, we observe income incentives and agglomeration economies at work. NCAA soccer players tend to come from higher income counties, but we also find that regions tend to develop sport-specific specializations over time, specifically, the Northeastern states.

INTRODUCTION

Soccer is a seemingly inexpensive sport. All you need is a ball, a field, and a few players to make a go of it. Yet most Americans regard it as a mere caricature of suburban prosperity in the United States. More perplexing, the U.S. Men's National Team has never been a serious contender on the world stage, and all of the income and marketing power in the country has achieved only a fraction of the revenues of the major U.S. sports. Since the NCAA is the ultimate proving ground for aspiring athletes in the U.S., we turn there for a logical explanation.

Sports teams and organizations share their knowledge directly and indirectly through the competitive landscape. Youth travel soccer is decidedly market-oriented. Indeed critics repeatedly level the "pay-to-play" argument against the U.S. Soccer Federation for the perennially weak showings of the U.S. Men's Soccer Team on the world stage. The argument assumes community-funded soccer would attract better athletes away from the more popular sports. An alternative view recognizes soccer's growth in the U.S., driven by market forces, as entrepreneurs have worked to provide a good that is under-supplied in many communities across the country. Forced to compete for resources and players in towns where traditional American sports are deeply embedded in the community culture, soccer has gained through entrepreneurial pursuits (Cuadros 2006). Add to this the influence of Title IX, through restrictions and incentives on college athletic programs, and the economic analysis of the collegiate sports is compelling.

Urban and regional economic theory provides a convenient lens for analyzing spatial phenomena. Youth sports apparently gain from regional concentration, similar to firms clustering in cities to benefit from labor pooling, knowledge spillovers, and the sharing of intermediate inputs, collectively referred to as agglomeration economies. Initial clusters of specific sport hotbeds may happen by accident, as Bigalke (2018) demonstrates for men's soccer and the Northeast with the 1930 U.S. World Cup team. The urban and regional literature has repeatedly documented agglomeration for places including Dalton, Georgia and carpet making, banking in Charlotte, technology in Silicon Valley and many more. Over time, the benefits of agglomeration tend to increase the concentration and economic value-added (Henderson 2003, Rosenthal and Strange 2001).

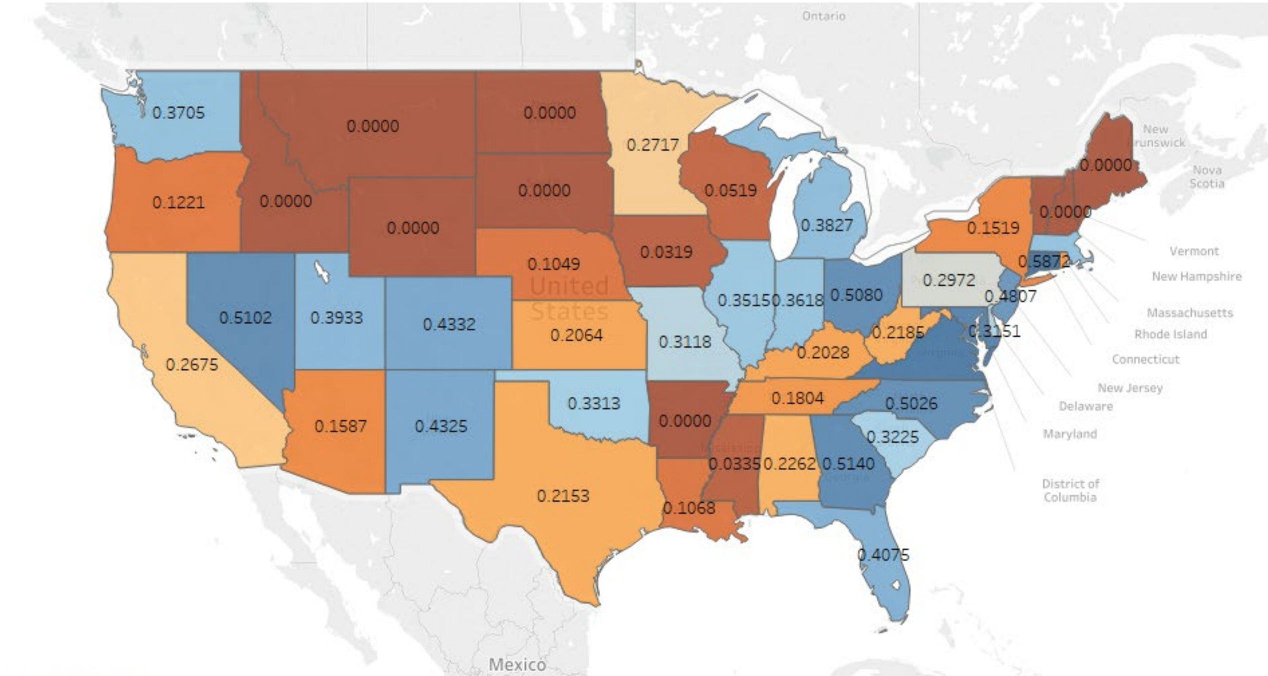
We examine Division 1 soccer participation through a sample of over 22,000 male athletes and over 8,000 female athletes on rosters at FBS schools in 2018. We limit the empirical analysis to major team sports based on data availability (football, baseball, basketball, soccer, and hockey for men, and soccer, softball, volleyball, and basketball for women) and find regional clustering consistent with economic theory. Our findings reinforce much of the work expounded in Kuper and Szymanski's (2014) *Soccernomics,* but we provide evidence suggesting a much weaker climb toward prominence in men's soccer than they suggest for the U.S. among the world's elite soccer nations. Examining the collegiate soccer player pool reveals important spatial and economic attributes of areas producing relatively higher numbers of soccer athletes at major universities and emphasizes a pecuniary tradeoff in favor of other major U.S. sports. The remainder of the paper proceeds as follows. We discuss the background for the research and economic intuition in Section 2 and present the empirical analysis in Section 3. We conclude with the results discussion in Section 4.

BACKGROUND

Sports participation studies typically focus on sport as a leisure or fitness activity. Consumers maximize utility across a set of leisure and labor alternatives and allocate time and income to the consumption of sport according to relative prices of available goods (Kokolakakis et al. 2012, 2014; Humphreys and Ruseski 2006, 2007). Other studies have examined the benefits of sports participation in high school in terms of college completion and income generation in the future (Barron, Ewing, and Waddell 2000). Sports economists have examined other more traditional microeconomic problems through the lens of sport, including anti-trust and industrial organization as well as labor markets more generally (Leeds and von Allmen, 2016; Downward, Dawson, and Dejonghe 2009). Our study focuses exclusively on NCAA sanctioned competitive sport participation. It does not apply to the European markets in general but makes a unique contribution to the sports economics literature. We believe this is the first study to examine the determinants of American collegiate soccer participation from a microeconomic framework and a local market geographical perspective.

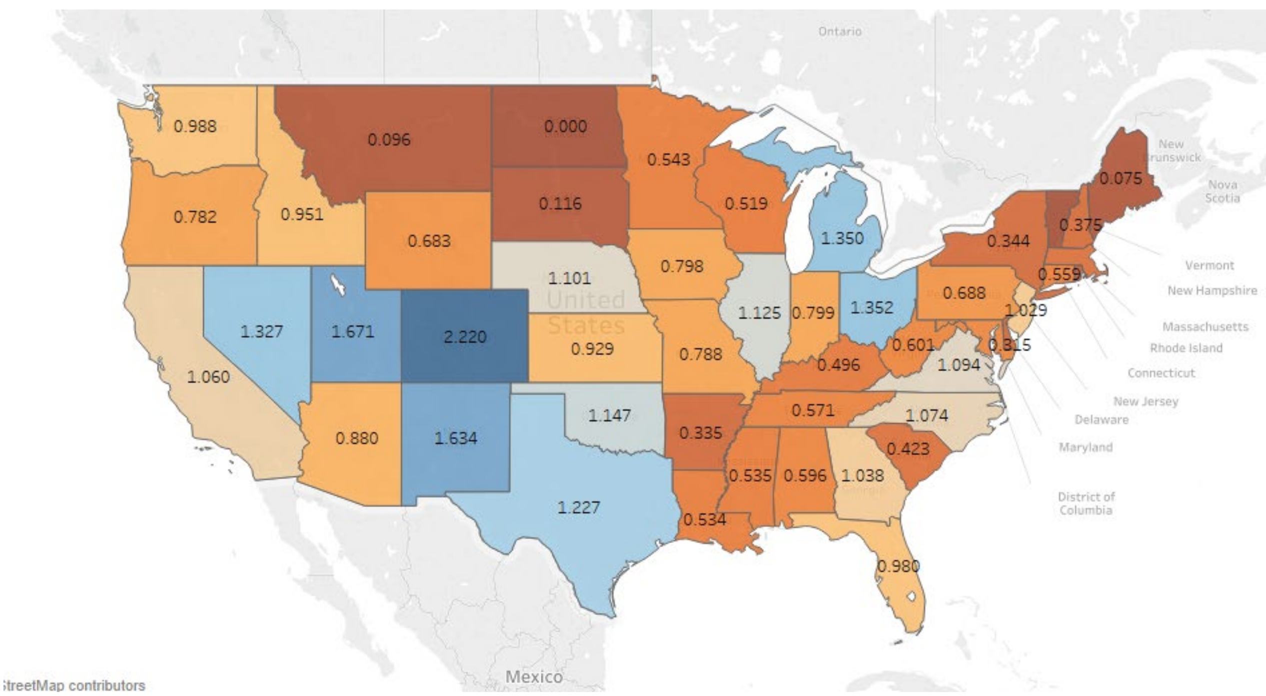
Given the relatively low popularity of collegiate and professional soccer in the United States, the decision for an athlete of high ability to choose to play soccer, over football or baseball, for example, would seem an inferior one. However, at the margin we assume the representative player-household chooses the best option relative to a host of limiting economic and spatial factors, including the local resources available and the competitive environment inclusive of other sports. We expect to find that certain areas are good places for soccer players to develop, and these places likely see more public and private resources go to soccer and have more people experienced with the sport. See Figures 1 and 2 for evidence of the variation in the number of soccer players per capita from each state playing D1 soccer (as measured by number of players in our dataset appearing on rosters at FBS hosting universities per 100,000 citizens).

Figure 1: Sending NCAA Men's Soccer per 100,000 Population



Source: 2018 NCAA Division 1 FBS School Rosters

Figure 2: Sending NCAA Women's Soccer per 100,000 Population



Source: 2018 NCAA Division 1 FBS School Rosters

Earning a college degree is challenging without the additional demands of playing an NCAA sanctioned competitive sport, yet for athletes the economic and psychic rewards are alluring. Young athletes and their families invest significant resources to increase the probability of receiving a Division 1 athletic scholarship. Assuming a constant supply of athletes, this probability increases with the number of scholarships available for each sport. Beyond college, some sports offer potentially lucrative professional opportunities, increasing the implicit value of the scholarship. This potential for earnings as a professional athlete varies across sports and is relatively low for men's soccer in the United States, in comparison to other sports. For women however, professional soccer may be relatively more attractive, given the smaller number of alternative opportunities for women in professional sport in the United States.

Sanderson and Siegfried (2015) provides an insightful discussion of the monetary incentives for colleges and universities with Division I sports programs and the vast financial differences across sports. For the male soccer athlete, the financial rewards in the form of scholarships available and potential for earnings as a professional athlete after college are much smaller when compared with the other major men's professional sports. Additionally, resources for youth soccer may reflect or partially contribute to the incentive differences. Regional disparities in median household income and educational attainment may skew the participation levels across men's team sports. These factors increase the relative value of an athletic scholarship for male athletes in the Southeast.

Other regional disparities may affect the environment for youth sport and individual development as well. The data show the South and West produce the most soccer athletes. Climate is an obvious factor in favor of these two regions, enabling more practice time for the outdoor sports. The availability of abundant, inexpensive land may also promote the development of large-scale recreational and athletic facilities, whereas higher-valued commercial and residential uses compete for land in more densely populated metros, such as those of the Northeast. Culture is another external factor that is difficult to measure. Soccer has greater economic and cultural significance in other countries. In the United States, soccer may have greater popularity where global influences due to immigration have been greater over time. This popularity should translate indirectly to ability, increasing the competitive environment where one's skills are refined. Ultimately, how these and other factors collectively affect an area's sports performance environment is an empirical question. In Table 1.A. we show states that send the greatest number of soccer athletes to FBS programs. Table 1.B., shows soccer athletes as a share of major FBS team sport athletes in our sample. The greater proportion of soccer athletes in the Northeast suggests culture, partly due to immigration and partly due to income and education, affects outcomes. However, fundamental economic factors may best predict NCAA soccer participation. In Section 3 we present our data, empirical models, and results.

Table 1.A: Soccer Players by State

Men's Soccer		Women's Soccer	
Top 5 States	**Soccer Players**	**Top 5 States**	**Soccer Players**
California	130	California	431
Florida	101	Texas	367
Texas	78	Florida	201
North Carolina	69	Ohio	148
Georgia	60	Illinois	140

Source: 2018 NCAA Division 1 FBS School Rosters

Table 1.B: Soccer as a Share of Total Athletes by State*

Men's Soccer			Women's Soccer		
Top 5 States	**Soccer Concentration**	**Soccer Players**	**Top 5 States**	**Soccer Concentration**	**Soccer Players**
Connecticut	13%	24	Colorado	65%	131
Washington	11%	35	New Jersey	60%	91
Massachusetts	11%	32	Utah	57%	52
New Jersey	10%	50	New Mexico	53%	36
Colorado	10%	29	Massachusetts	53%	30

*States with at least 1 million population producing at least 15 soccer players.
Source: 2018 NCAA Division 1 FBS School Rosters
Men's sports included in this dataset are Football, Baseball, Basketball, Soccer, and Hockey
Women's sports included in this dataset are Soccer, Softball, Volleyball, and Basketball

EMPIRICAL ANALYSIS

Data

We obtained individual soccer player data from 2018 team roster web pages available on the websites of the 130 universities hosting FBS football programs. Rosters for each sport provide position-specific information for each player as well as their hometown and high school of record, including zip code. With this geographical information, we add county level economic and demographic data. Where reporting was incomplete, such as no zip code reported, we attempted to match the player's hometown with their corresponding high school to determine precise zip code. The National Center for Education Statistics (NCES.ed.gov https://nces.ed.gov/ccd/pubschuniv.asp) provides a database of all U.S. public and private high schools with addresses. For approximately 10% of the athletes on rosters, we could not reliably match reported information to the high schools in our database. Reasons for a failed match include the following: student attended high school in another country, student was home schooled, no high school was reported, or the name of the high school provided on the website was either not present in the database or no unambiguous match was present.

Our economic and demographic indicators come from the 2016 American Community Survey 5-year estimates for counties. Excluding U.S. territories, as well as Kalawao County, Hawaii (which is an isolated peninsula that has served as a leper colony since 1870, and is now solely populated by the surviving colony members, along with caregivers and support personnel), and Loving County, Texas, with a population of 74. Both of these counties had incomplete data on housing values, and perhaps not surprisingly, had no representation of athletes on rosters of any sport in our dataset. This leaves 3,140 counties remaining in the dataset. We include measures for income, population, population density, education, race, and age. A binary variable for each county captures roster information from Bigalke (2018) that indicates whether a county is in one of the five states that produced U.S. Team soccer players for the 1930 FIFA World Cup in Uruguay. These states were New York, New Jersey, Pennsylvania, Massachusetts, and Missouri. Ten of the sixteen players on the 1930 team came from these states, with the remaining six players coming from Great Britain. The 1930 team was the most successful U.S. Men's World Cup soccer team in history. This indicator attempts to capture a cultural or local scale effect that income or education may not explain. Finally, a count variable for the number of FBS schools in each state captures an element of demand for athletes and visibility of the sport and opportunities at the Division 1 level. For complete data source information, definitions, and summary statistics, see Tables 2 and 3.

Modeling and Results

We model NCAA soccer participation at the county level to understand where soccer athletes come from and what factors tend to support greater numbers of soccer athletes from a given region playing NCAA Division 1 soccer at FBS schools. Our modeling approach needs to incorporate individual motivation as well as the regional economic and demographic environment. Ideally, we could observe household level characteristics over time, following the approach of Farrell and Shields (2002). Their study of leisure sport participation exploits data on individuals in households in England and models the unobserved propensity to participate in leisure sports. They estimate a random effects probit model of the following form:

$$S^*_{ih} = x'_{ih}\beta + v_{ih} \text{ for } i = 1,2,\ldots,n \text{ and } h = 1,2,\ldots,H, \quad (1)$$

$$V_{ih} = \alpha_h + \mu_{ih} \quad (2)$$

and

$$S_{ih} = \begin{cases} 1 & \text{if } S^*_{ih} > 0, \\ 0 & \text{otherwise} \end{cases}$$

where S is observed sports participation for the i^{th} individual from household h. Exogenous, observable factors include demographic characteristics, region aspects, and individual health. The composite error term, v_{ih}, captures unobservable household preferences toward leisure sports.

Table 2: Variable Definitions and Sources

Variable	**Definition and Source**
Men's soccer players	Number of men's soccer players from county on any FBS NCAA 2018 roster
Women's soccer players	Number of women's soccer players from county on any FBS NCAA 2018 roster
Gender population 18-24 (male)	Percent of county population ages 18-24, male
Gender population 18-24 (female)	Percent of county population ages 18-24, female
Uruguay 1930	Binary indicator for county in NJ, NY, PA, MA, MO (calculated from Bigalke 2018)
Population	County population from 2016 American Community Survey 5-year estimate
Median HH Income	County median household income from 2016 ACS 5-year estimate
Median Home Value	County median value of owner-occupied housing from 2016 ACS 5-year estimate
Pct. HS Grad	Percent of county population age 25+ with HS diploma or higher (2016 ACS/5-year)
Pct. Bachelor's Degree	Percent of county population age 25+ with Bachelor's degree or higher (2016 ACS/5-year)
Population per square mile	County population density calculation as a function of land area (sq. mi.)
Percent Black	Percent of county population that is Black race (2016 ACS/5-year)
Percent Hispanic	Percent of county population that refers to a person of Cuban, Mexican, Puerto Rican, South or Central American, or other Spanish culture or origin regardless of race (2016 ACS/5-year)
FBS schools	Number of FBS football program schools in state as of 2018 (NCAA)

Table 3: Summary Statistics of Dependent and Independent Variables

Variable	Obs.	Mean	Std. Dev.	Min	Max
Men's soccer players	3140	0.31	1.39	0	21
Women's soccer players	3140	0.94	4.20	0	85
Gender population 18-24 (male 1000’s)	3140	5.1	16.6	1	531.7
Gender population 18-24 (female 1000’s)	3140	4.9	16.0	0	516.6
Uruguay 1930	3140	0.09	0.28	0	1
Population (1000’s)	3140	102.2	328.4	.289	10,105.7
Median HH Income ($1000’s)	3140	$ 49.7	$ 13.1	19.3	$ 129.6
Median Home Value ($1000’s)	3140	$ 141.3	$ 85.1	18.7	$ 995.9
Percent HS Grad	3140	86.2	6.5	41.3	98.9
Pct. Bachelor's Degree	3140	21.2	9.3	4.9	78.1
Population per square mile	3140	224	1,284	0.03	49,105
Pct. Black	3140	9	14.5	0	87
Pct. Hispanic	3140	4.9	8.0	0	74
FBS schools	3140	3.7	3.1	0	12

Since we do not have household data or individual player characteristics other than those previously noted, we could not feasibly estimate the probability of an individual’s decision to play NCAA soccer with the methodology of Farrell and Shields (2002). Instead, we model the county propensity to supply more or fewer NCAA soccer athletes. There are several methods we could employ, but the two most tractable include a logit model estimation of the share of athletes per county and an ordinary least squares regression of the number of players per county as a function of a mix of observable right-hand side variables. The OLS model is straightforward and convenient to interpret. Because of Title IX, there are significant differences in participation for male and female soccer, with 60 men’s soccer programs and 130 women’s soccer programs at FBS schools. For this reason, we use separate models for male and female participation. We also control for state fixed effects and cluster the standard errors in each model at the state level. We report the OLS results here, but the findings from logit estimation are consistent with these results and available upon request.

The general model predicting the number of county male soccer players takes the following form and is otherwise identical for female soccer players:

$$SC_{is} = x'_i\beta_j + f_i\gamma + w_i d + u_i \text{ for } i = 1,2,\ldots,n \text{ and } s = 1,2,\ldots,S \quad (3)$$

where SC_{is} is the total number of male soccer players coming from county *i* of state *s* and ranges from 0 to 21. For females the count ranges from 0 to 85. We also tried models in which we scaled the dependent variable by county population, creating a players per capita model. The applicability and interpretation is not as intuitive. Ultimately, soccer clubs and college recruiters are interested in the richness of a soccer-producing area, in terms of the number of players. The per capita measure does not translate readily into recruiting demand. For example, the highest per capita county for male soccer players is in Nebraska, with only one soccer player but less than 5,000 people in the county. This is not a meaningful statistic. We have estimated these models and can provide the results upon request. The coefficient estimates are not fundamentally different from our absolute number model results.

In the following models, we have 3,140 county level observations. State level fixed effects control for unobservable differences across counties within a state. Analogous to Farrell and Shields (2002), each county is technically a member of a state, and a state may induce better or worse conditions for the development of NCAA soccer athletes. Our estimation clusters the standard errors by state to ensure accurate variance calculations for inference testing. The vector *x* includes economic and demographic variables. The variable *f* represents the number of FBS programs in a state and ranges from 0 to 12. Finally, *w* is the binary indicator for whether the county is in one of the five states that provided players for the 1930 FIFA Men's World Cup team.

The first set of results examines the supply of NCAA soccer players controlling for the general population of their home county. In Table 4, we find that a doubling of population produces a near doubling of soccer players per county. The coefficient on population, converted to population in hundreds of thousands, is 0.29. The mean population for a county in our sample is 102,231, while the mean number of NCAA-bound soccer players per county is 0.31. Thus, an increase of 100,000 people produces an additional 0.29 players, almost double the mean of 0.31. We find a similar effect for females. The coefficient on population is much stronger, 0.97, per 100,000 population, but the mean number of female players is greater at 0.94. Population density has a negative effect, but it is minor in magnitude. This is an interesting result compared with population's positive effect. It is consistent with scarcity of land supply that forces youth development organizations further from the central city, out to the suburbs. Higher income in the suburbs may also contribute to the negative density effect.

Income has a positive effect on NCAA soccer player production. For males, increasing county median household income \$10,000 yields 48 percent more soccer players, an elasticity value of 2.4. The effect is even greater for females. The coefficient of 0.64 yields an elasticity of 3.4. Similarly, we find a more college-educated population is consistent with greater soccer player production. For males, the coefficient on *Pct. Bach* is 0.023. A one unit increase in the percentage of the population with at least a bachelor's degree, an increase of 5 percent at the mean, yields 7 percent more soccer players. For females, the elasticity effect is similar. We control for adult population with at least a high school diploma or equivalent in the same models. Holding constant the population with a bachelor's degree or more, the *Pct. HS Grad* variable captures the difference in counties with less college attainment, since each of these variables has no limit in the upper bound of education, only a bottom. An increase in a county's population with no more than a high school diploma has a negative effect, but it is only significant in the women's soccer model. Women's soccer players at the D1 level from a county fall 8 percent for a 1-unit increase in the share of the adult population with only a high school diploma or equivalent.

We also control separately for the Hispanic and Black percentages of county population. Similar with the results for high school education, the percent of population of Hispanic origin has no significant effect in the men's model but a negative and significant effect in the women's model. We suspect some multicollinearity between the measures for high school education and Hispanic origin.[2] One plausible reason for the difference between the coefficients on percentage of Hispanic population in the men's and women's models could relate to the difference in culture across the populations. Hispanic countries have a culture that is less interested in female athletics in general, perhaps indicative of overall having less equality for females. Thus, we may reasonably expect that this cultural element would lead to lower female participation in soccer among those populations. There is likely less cultural distance between the American Hispanic population and Hispanic countries. We also expect that this coefficient difference is not obtained with the black percentage as there may be a much more significant cultural distance between African Americans and African cultures as compared to Hispanics and Central- and South-American countries.[1]

We also controlled for the states that produced male players for the 1930 FIFA World Cup competition in Uruguay, the *Uruguay 1930* variable to see if a possible outsized regional interest in soccer persisted nearly a century later. The coefficient was significant at the 5 percent level for males but insignificant for females. The coefficient in the model for male soccer players is 1.34. Recall, this is a binary indicator. A county in one of these five states has 1.34 more male soccer players on a Division 1 roster. At the mean of 0.31 male players per county, these states, all other things equal, have 4.3 times the number of college-bound soccer players. We find this apparent persistence of the cultural tradition of playing soccer interesting, and potentially important for future analysis of patterns of sports adoption in a particular area.

We also tested the model for the 18-24 year-old population in a county to see if the relevant population pool increased the precision of the estimation. The R-squared is slightly lower for these models, 0.5538 for men compared with 0.5872 in the previous model, and 0.6098 for women compared with 0.6436 in the previous model. The coefficient estimates are similar. The effect of more population in the 18-24 year-old range is near unitary elastic for males. An increase of 1,000 18-24 year-olds, about 20 percent, produces 17.4 percent more players. For females, a 20.6% increase in 18-24 year-old population yields 20% more players.

Table 4: Soccer players by county

Variable	Men's Soccer Coefficient	t-Stat	Women's Soccer Coefficient	t-Stat
Population	0.0029	6.53	0.0097	14.01
Median HHI	0.0145	3.30	0.0640	4.86
Median Home Value	-0.0006	-0.54	-0.0054	-1.80
Pct. HS Grad	-0.0131	-1.22	-0.0794	-4.43
Pct. Bach	0.0270	3.64	0.0871	4.51
Pop. Per Sq. Mile	-0.0001	-2.83	-0.0004	-5.59
Pct. Black	0.0060	2.70	0.0061	1.34
Pct. Hispanic	0.0006	0.54	-0.0531	-4.01
FBS Schools	0.0053	0.30	0.0183	0.41
Uruguay 1930	1.3428	2.11	1.4686	0.96
Constant	-0.2202	-0.25	4.0226	3.03

(OLS using county level data, 3140 observations. R-squared = 0.5872 for Men's, 0.6436 for Women's. State dummy variables omitted from table.)

Table 5: Soccer players and 18 – 24 population per county

Variable	Men's Soccer Coefficient	t-Stat	Women's Soccer Coefficient	t-Stat
Pop. 18-24 (1,000s)	0.05384	6.02	0.1906	12.43
Median HHI	0.01820	4.05	0.0781	5.90
Median Home Value	-0.00310	-0.25	-0.0039	-1.32
Pct. HS Grad	-0.14298	-1.19	-0.0812	-4.37
Pct. Bach	0.02306	2.88	0.0681	3.49
Pop. Per Sq. Mile	-0.00006	-1.81	-0.0004	-4.23
Pct. Black	0.00649	2.73	0.0076	1.47
Pct. Hispanic	0.00573	0.46	-0.0526	-4.36
FBS Schools	0.00688	0.39	0.0250	0.58
Uruguay 1930	1.39759	2.16	1.6574	1.04
Constant	-0.26257	-0.26	3.6918	2.66

(OLS using county level data, 3140 observations. R-squared = 0.5538 for Men's, 0.6098 for Women's. State dummy variables omitted from table.)

CONCLUSION

We offer empirical evidence that income, population, education, and to some extent, culture and tradition have measurable impacts on an athlete's propensity to play college soccer in the United States. Despite the fact that soccer is arguably the easiest logistically (you can practice with varying numbers of players) and least expensive sport to play in theory (all you really need is a ball and a field to play), the results suggest soccer athletes come from higher income areas. A lack of community support may partially drive this, but that may be the result of a lack of enthusiasm among student-athletes. They do not see soccer as a path to high status or prosperity in the United States. Baseball, basketball, and football players are highly paid, with the top players potentially earning tens or even hundreds of millions of dollars over the course of their careers, and enjoying celebrity status, while the top U.S. male soccer players are relatively unknown.

Therefore, while the sport of soccer has some scholarship opportunities, student-athletes likely do not perceive it as a lucrative career path in the United States. The economic logic suggests it is more of a leisure sport than an investment in future earnings, aside from potential college scholarships. Intuitively, NCAA soccer players are more likely to come from higher income, more highly educated areas, and the data is consistent with this.

We have also demonstrated the relevance of the market environment for soccer. It is a global sport, yet the competition for athletes necessary to deliver the product is a local phenomenon. While Kuper and Szymanski (2014) suggest that the United States should over time produce an elite-level men's soccer team, we suggest it may take longer than Kuper and Szymanski predict it will take to do so. Even a large, rich market must confront the competitive landscape and economic reality that other major sports offer the potential for greater returns for the typical student-athlete, which dampens its appeal in the U.S., relative to most other countries. Incentives matter. This is just a state of the world.

ENDNOTES

[1] The latest Current Population Survey release (2018 – Tables 1-2, 1-4, and 1-6) shows only 73 percent of the Hispanic origin population age 18 and over with at least a high school diploma, while the percentage is 87% and 90% for Black and White populations respectively.

[2] We would like to thank the editor for the suggestion about the Hispanic variable interaction and high school education attainment. We also thank our colleague, Dr. Phillip Njoroge, for the suggestion of cultural difference. Separate effects for women's soccer have also been found in published research. See Congdon-Homan and Matheson (2013).

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